THEORETICAL AND EMPIRICAL ISSUES IN TOURISM DEMAND ANALYSIS

by

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Thesis submitted to the University of Nottingham

for the degree of Doctor of Philosophy, June 2001

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ABSTRACT

The majority of empirical studies of tourism analysis use a static single equation approach to model the demand for tourism of one origin for one or more destination countries. The examination of such studies generally shows that the economic interpretation and policy implications drawn as conclusions are based on mis-specified models, invalid estimation and inference procedures, inconsistent estimates and poor forecasting performance. Static single equation models of tourism demand tend to neglect interdependencies among destinations, ignore nonstationarity, overlook dynamics and, generally, disregard economic theory. Empirical specifications constrained by these flaws are bound to generate biased and inconsistent estimates upon which no reliable economic analysis or policy implication can be based.

In an analytical context that focuses on the UK tourism demand for France, Spain and Portugal in the period 1969-1997, the main objective of this thesis is to demonstrate that consistent elasticities' estimates and reliable forecasts can be obtained from empirical models which are based on the principles of economic theory, and specified and rigorously tested within the rules of sound econometric methodology. The alternative models estimated in chapters 4 to 7 include error-correction autoregressive distributed lag models (ARDL), static and dynamic almost ideal demand systems (AIDS) and cointegrated vector autoregressive models (VAR).

The main findings that emerge from the study are as follow. The battery of diagnostic tests applied to the dynamic error-correction ARDL models provide sufficient evidence to classify them as statistically robust, structurally stable and well-defined specifications. The evidence obtained for the AIDS and VAR systems indicates them as data-coherent and theoretically-consistent models, complying with the utility maximisation hypotheses. The similarity, across models, of the estimates of the long-run structural parameters and the accuracy of the forecasts they provide further support the reliability of these models for explaining and predicting the UK tourism demand behaviour, in contrast to the static single equations estimated in chapter 3. The specifications of chapters 4 to 7 can easily be extended, without loss of generality, to more origins and destinations and can be adapted to alternative contexts such as the demand for specific regions within a country, specific resorts within a region or even specific types of tourism products such as accommodation or leisure facilities, within a local area.

ACKNOWLEDGEMENTS

I am deeply grateful to my supervisor, Professor Thea Sinclair, for her insightful comments, advice and constant encouragement while preparing the thesis. More than an excellent supervisor, Thea is a friend from whom I have learned invaluable lessons, which will help me in my future professional duties and research quest.

My thanks are also due to Mr. Alan Pack and Professors Anthony Thirlwall, Alan Carruth and Manuel Mendes de Oliveira who were always helpful and generous in providing guidance, assistance and advice whenever asked.

I would like to thank Professors Alistair Bruce and Paul Fenn for their comments on several chapters of the thesis which helped me to improve its contents. Thanks are also due to the anonymous referees who commented on the model and results of chapter 5.

I thank all colleagues in the Christel DeHann Tourism and Travel Research Institute for their support and help and particularly, Sarah Gerard, whose competence, proficiency and expertise always makes the most difficult problem appear simple and easy to solve.

I wish to acknowledge the encouragement and support of my dear friends Cristina, Erica, Ria, Sophia, João, Jorge, Miguel, Owen, and Roy who made the life of a PhD student seem almost bearable and, at times, even bright and fun.

I am deeply grateful to my parents, brother and sister and to my son, for taking care of me in the darker hours and for giving me constant support during my study years. Without them this thesis would not have been possible.

Finally, I am grateful to Kevin for sharing the joys and pains of research in economics; for the exciting, inspiring, enjoyable and endless discussions about econometric modelling and economic theory (seems a contradiction in terms, but it is not!); for lifting my spirits high when trouble was in sight, and keeping my feet down when enthusiasm was excessive; for constructive comments on my 'original' ideas and devastating criticisms on my self-indulgent concepts; for brightening my laughs and drying my tears; for listening, reasoning, caring, reassuring, encouraging and comforting; for making my life a wonderful place to live...

CHAPTER 1

INTRODUCTION

1.1. THE IMPORTANCE OF TOURISM DEMAND ANALYSIS

In the last half century, tourism has become one of the world's most important economic activities. Between 1950 and 1997, international tourism receipts increased at a remarkable annual average rate of 12% and although this growth rate has been slowing down since the mid 1970s, it remained 7% over the last decade, more than double the world's GDP growth[1] (WTO, Yearbook of Tourism statistics, 1998, p.2-3). During 1997, 617 million tourists travelled the world spending, $US 447 billion, and the World Tourism Organisation (WTO) predicts that by 2020, there will be 1.6 billion tourists spending $US 2,000 billion. According to the same source, tourism is responsible for more than 220 million direct and indirect jobs and, if complementary economic activities linked to tourism are included, tourism accounts for about 11% of the world's GDP.

These figures illustrate the importance of tourism in the world economy. Indeed, revenue from foreign tourists creates and sustains jobs, generates additional income for private and public entities, alleviates trade deficits, increases foreign exchange reserves and finance imports, acts as a catalyst to investment and, overall, contributes to the economic growth of destination countries. In addition, tourism often requires investment in physical and human capital, and its returns can be realised relatively quickly. Therefore, the activities associated with tourism are generally considered vital to assist development in local and regional areas of developing and industrialized countries.

As tourism emerges to a centre-stage position in regional and national economies, it is important to measure and predict tourists' demand patterns and behavioural features. Whether and how much net economic benefits tourism brings to destinations depends on

[1] The World Economic Outlook Database of the IMF states that the world GDP growth rate in the 1990s is 3.4%. http://www.imf.org/external/pubs/ft/weo/1999/01/data/ngdpd_a.csv.

the precise form and scale of tourism demand. The key role of tourism demand in the success of many economic activities requires knowledge of its main determinants and accurate forecasts of its future levels. Not surprisingly, in the last three decades, research on the economics of tourism has been a growing area of interest for both business and academic sectors.

1.2. THE OBJECTIVES OF THE STUDY

Evaluation of the magnitude and direction of the impacts of tourism expenditure on destinations requires estimates of current and future demand. Hence, tourism demand modelling and forecasting studies have been a growing area in economic literature. However, the majority of studies of tourism demand analysis fail to incorporate the theoretical basis and methodological tools fundamental in the construction of accurate and reliable models for explaining and predicting tourism phenomena.

Empirical studies of tourism demand generally involve the use of econometric models to specify the relationships between the demand levels and its determinants. Econometric modelling provides a good basis for forecasting, which is of considerable value for public policy and an important element in public and private investment decisions. However, modelling tourism demand presents several difficulties, which are mainly linked with its specific features. First, most of the time series used in the estimation of tourism demand are trended or non-stationary. Models that overlook this feature of the data may give rise to spurious regressions, invalidating statistical inference and forecasting procedures. Second, consumer theory hypothesis should be integrated and tested within the quantitative framework adopted to model tourism demand. Models that are theoretically inconsistent do not serve well the purposes they aim to achieve. Third, the inherent dynamic nature of tourism demand and the possible existence of feedback effects requires the explicit incorporation of a time dimension and the consideration of short-run adjustments within a system of equations structure.

The difficulties encountered in the construction of the models are frequently overcome by means of simplifying assumptions, which permit the specification and estimation of quantitative relationships explaining and predicting tourism demand behaviour. Depending on the assumptions underlying the construction of the model, the subsequent specification will differ both in its static or dynamic nature and in the variables included, functional form adopted and estimation methods used. Different models perform differently, being more or less reliable according to the theoretical and methodological frameworks within which their specification is formulated. Tourism demand models constructed under questionable assumptions, overlooking dynamics and feedback effects, ignoring the spurious regression problem, neglecting interdependencies among competing destinations, and lacking the theoretical basis within which testable hypothesis of consumer theory can be included, generally give rise to misspecification bias, unreliable and misleading estimation results and overall invalid statistical inference and forecasting procedures.

The literature concerning tourism demand analysis shows that the large majority of early empirical studies have used a static single equation approach to model an origin's demand for tourism in one or more destinations. These *ad hoc* models tend to be based on implausible assumptions, lack consistency with consumer behaviour theory and, as Witt and Witt (1995, p. 458) observe, "*the quality of the empirical results obtained is questionable*". A few recent studies in tourism demand behaviour have attempted to overcome the problems associated with the traditional single-equation demand models by considering one or more of the problematic aspects of their modelling strategies. However, to the best of our knowledge, none has yet addressed these problems in a systematic way.

Against this background, the main objectives of this thesis are to discuss and implement alternative methodological approaches to tourism demand analysis which contribute both to strengthening the theoretical foundations of currently used models, and to apply and evaluate recent advances in econometric modelling, quality evaluation criteria, hypothesis testing and forecasting procedures in a tourism demand context.

1.3. THE SIGNIFICANCE OF THE STUDY

The increasing importance that, in recent years, has been attached to tourism demand analysis clearly demonstrates the need to extend the theoretical and empirical content of the existing literature and, more important, to substantiate the empirical findings with appropriate and extensive testing. In this context, this study makes several important contributions. The thesis analyses the UK tourism demand for its geographically proximate neighbours, France, Spain and Portugal using data for the period 1969-1997. The choice of the countries involved in the empirical analysis took into account the fact that the UK is a major tourism origin, which is of particular importance to France, Spain and Portugal as destinations. Spain and Portugal are interesting cases for consideration owing to their position as economies in transition during the sample period. At the beginning of the period, in 1969, they displayed classic symptoms of underdevelopment: high dependence on agriculture and fishing, lack of industrialisation, low income and low standards of education, health provision and other indicators of social welfare. By 1997, the final year of the period under study, they had joined the ranks of more developed European economies. France was a high-income country over the whole sample period, allowing for useful comparison between it and its poorer neighbours. The concept of neighbourhood also determined the choice of countries permitting analysis of the destinations' dynamic competitive behaviour and its interdependencies.

It is well known that tourism demand is responsive to such variables as income, relative prices and exchange rates. What is not known is how the responsiveness of demand to changes in these variables alters during a country's economic transition and integration into a wider international community. Relationships of substitutability or complementarity may change over time as lower income destinations emerge from relative poverty to achieve higher levels of development. Little information is available about whether lower income destinations tend to become more or less competitive over this transition period, either relative to other developing countries or relative to more industrialised nations. This study provides some interesting insights concerning these issues.

The vast majority of studies of tourism demand have relied on single equation models within a static context (for example, Loeb, 1982; Uysal and Crompton, 1984; Gunadhi and Boey, 1986, Lee *et al.*, 1996). These models are not derived from consumer demand theory, generally disregard dynamics, the non-stationarity of the time series involved, and potential simultaneity bias. As a result, the estimation results are unreliable, the statistical inference invalid and the forecasting ability of such models is so poor that, as showed in Martin and Witt (1989) and Witt and Witt (1992), even the simplest of univariate time series model – the naïve no-change specification – can supply more accurate forecasts.

In this study, the specification deficiencies, theoretical flaws and technical inadequacies associated with traditional tourism demand modelling are examined and overcome by the construction and estimation of alternative econometric specifications which are both theoretically consistent and empirically plausible. These specifications incorporate dynamics through short-run adjustment mechanisms; consider the interdependences among destinations and test consumer theory hypothesis using system of equations; avoid simultaneity bias using Sims' (1980) vector autoregressive approach, and prove the existence of a long-run equilibrium relationship between the UK tourism demand and its determinants according to the Johansen (1988) cointegrating vector analysis. The models estimated in this study are subject to rigorous quality scrutiny under the rules of the most recent econometric methodologies, such as structural constancy testing, causality and exogeneity testing, cointegration analysis and encompassing. Finally the alternative models are compared in their forecasting ability and reasons for their different performances are analysed.

An important approach to tourism demand analysis involving systems of equations is the Almost Ideal Demand System (AIDS) of Deaton and Muellbauer (1980b). This model embodies the principles of consumer demand theory and is particularly valuable for testing the theoretical hypothesis of homogeneity and symmetry and for the estimation of cross-price elasticities between competing destinations. Some studies, for example, O'Hagan and Harrison (1984) and Syriopoulos and Sinclair (1993), investigate tourism demand using the orthodox static AIDS approach of Deaton and Muellbauer. More recent studies using the AIDS approach add a trend and or other dynamic-like elements to the orthodox model. In contrast with previous findings where homogeneity and symmetry are systematically rejected, these 'unorthodox' models seem to supply a 'quasi-dynamic' structure necessary to support their consistency with the constraints of consumer demand theory, (Papatheodorou 1999; De Mello *et al.* 2001). Nevertheless, the orthodox AIDS model, with or without trend, is derived within a static framework. Specific research on tourism demand using an explicitly dynamic AIDS system of equations is virtually nonexistent and, to the best of our knowledge, only Lyssiotou (2000) addresses this issue in an empirical study concerning the dynamics of adjustment behaviour within a system of equations similar to the AIDS model. In this thesis, we construct and estimate a dynamic AIDS model which, besides the desirable properties already present in its static version, explicitly adds the fundamental time dimension inherent to tourism demand behaviour.

In tourism demand research the modelling of dynamics has generally been confined to the use of error correction single equation specifications based on the Engle and Granger (1987) two-stages approach, (Kulendran, 1996; Kim and Song, 1998; Vogt and Wittayakorn, 1998; Song *et al.*, 2000). One disadvantage of this method is that the usual test statistics to evaluate the quality of model are not strictly valid and cannot be used for inference. Another disadvantage is that this approach does not prove that the cointegrating regression (if one is found) is the unique long-run equilibrium relationship. One alternative method, which may overcome these disadvantages, consists of the derivation of an error correction model based on a general autoregressive distributed lag model (ARDL) as developed by Pesaran and Shin (1995, 1996). The battery of tests proposed in Pesaran *et al.* (1996) can then be applied to confirm or reject the existence of a long-run relationship. In the existing literature, only Song and Witt (2000) supply a condensed example of an application of this method in a tourism demand context. This thesis contributes a chapter to the analysis of the ARDL error correction model.

Although the vector autoregressive (VAR) approach and the cointegration vector autoregressive analysis have been increasingly used in most areas of economic research for the last two decades, researchers on tourism demand analysis have generally ignored this new technique. Exceptions are Kulendran and King (1997) and Kulendran and Witt (2001). To overlook the importance of the VAR modelling approach in tourism demand analysis is to leave out a reliable econometric tool for the estimation and forecasting of the long-run impacts on demand induced by changes in its determinants. Indeed, the VAR approach can overcome such problems as spurious regression, simultaneous bias and identification issues arising from the nature of the variables and the theoretical behavioural features included in the quantitative relationships linking demand to its determinants. In this thesis we use a cointegrated VAR system to estimate the structural coefficients of the UK demand for tourism.

1.4. METHODOLOGY OF THE STUDY

A major debate in the empirical modelling of consumer preferences is associated with data aggregation issues. The advantage of models with data at the individual level is that they avoid potential aggregation bias. However, reliable longitudinal data sets following the same consumers over long periods of time are generally rare and, in tourism contexts, virtually nonexistent. Therefore, some form of aggregation is often unavoidable and, in the case of this study, no other choice could have been made.

Throughout the thesis, a great deal of attention is paid to the analysis and interpretation of events that have affected the time path of the series included in the models. The political events that took place in Portugal and Spain during the 1970s, the changes that occurred in these destinations preceding and following their integration into the EU in 1986, the opening of the Channel Tunnel in 1994, and events, such as the oil shocks in 1973 and 1979, which affected economies worldwide, are analysed from a tourism demand perspective and integrated in the models under plausible hypotheses. The statistical relevance of the variables representing such events suggest that the empirical results obtained are only meaningful and relevant if a robust general knowledge of historical facts affecting the time series involved, is fully and adequately integrated in the modelling procedures. Therefore, an extensive analysis of the time series properties is a fundamental part of the empirical methodology. The descriptive analysis of tourism demand time series is carried out in chapter 2.

The thesis contains a theoretical, econometric and analytical content. Each chapter proceeds with a thorough discussion of the theoretical issues underlying the derivation of the models and econometric methodology applied in their estimation and testing.

The static single equation models used in chapter 3, serve the purpose of demonstrating the methodological flaws of the traditional approach in tourism demand analysis. The theoretical framework of the dynamic single equation error correction models of chapter 4 follows that of Pesaran and Shin (1995, 1996) and Pesaran *et al.* (1996), incorporating the short-run dynamics missing from the static models, and permitting cointegration analysis in an equation by equation basis. A system of equations, however, can provide a more efficient method of modelling interrelationships among destinations, and permits the imposition and testing of theoretical restrictions of consumer demand behaviour. Moreover, the cointegration analysis implemented on a system basis is both more efficient and reliable than that performed on a single equation basis. Hence, a system of equations approach is used to derive the models estimated in chapters 5, 6 and 7.

As pointed out in Granger (1981, 1990), Harvey (1990, 1993), Hendry (1987, 1995), Hendry and Mizon (1978), Hendry and Richard (1982, 1983), Leamer (1983, 1987), Phillips (1986), Sims (1987) and in many other studies, the appropriate econometric modelling of quantitative relationships relies on a thorough examination of the economic time series included, as their statistical properties reflect the features of the data generating process which must be approximated by the empirical specifications. The analysis of the time series included in the models is carried out in chapters 4 and 7 and standard unit root tests of Dickey and Fuller (1979, 1981) are performed to establish the order of integration of the variables involved. However, these tests may suffer from low power in small samples and tend to be biased when the variables contain structural breaks. In such cases, other methods, such as the Phillips and Peron (1988) test, are applied to establish the variables' order of integration.

The selection of functional form to represent consumer preferences is a very important issue in empirical studies of demand behaviour. The ability to model the preferences' structure in an appropriate way relies on choosing a pertinent functional form which is both adequate and tractable without being excessively restrictive. The econometric models derived in this study are based on mainstream economic theory specifications. However, they are modified in several ways to apply to a tourism analysis context, with specific features attached to the relationships between the origin and the destinations considered and among the destinations themselves.

Within a system of demand equations, there are many flexible functional forms that can be used to approximate the consumers' indirect utility or cost functions and these forms may differ substantially in their approximation properties. Although there are many other classifying possibilities (see, for example, Lewbel, 1987), the set of flexible functional forms usually adopted in empirical demand analysis can be divided into three major subgroups (Fisher *et al.*, 2001): locally flexible functional forms, which include the translog models of Christensen et al. (1975) and Jorgenson *et al.* (1980), the AIDS specification of Deaton and Muellbauer (1980) and the generalised Leontief model of Caves and Christensen (1980); globally regular functional forms, which include the minflex Laurent models discussed in, for example, Barnett *et al.* (1985, 1987), the general exponential model of Cooper and McLaren (1996) and the quadratic AIDS model of Banks *et al.* (1997); asymptotically globally flexible forms, which include the Fourier flexible model discussed in Chalfant and Gallant (1985) and the asymptotically ideal model of Barnett and Yue (1998).

Given a specific data set, some functional forms will generally have more desirable approximation properties than others. A knowledgeable and well-founded choice of one would imply the specification and estimation of all for the same data set, and the definition of quality criteria, based on which one functional form could be considered to over-perform the others. Such an extensive analytical effort could be viewed as a subject of research on its own right, and is, therefore, beyond the scope of this study but should be considered in future research.

As Hendry (1995) points out, the existence of a potentially large number of theoretically consistent models, which satisfy the required quality criteria makes model choice a non-trivial problem. The criteria by which a good empirical model is judged are necessary but not sufficient requirements, since the failure of any may indicate inadequacy of the model, but the fulfilment of all gives no guarantee of the model's ongoing applicability. The choice of the functional form for the systems of equations in chapters 5 and 6 had these considerations in mind, as well as the fact that, with the exception of the AIDS model, all the alternative functional forms mentioned above specify non-linear share equations.

The system of equations in chapter 5 is based on the static AIDS model of Deaton and Muellbauer (1980) and the system of equations in chapter 6 is a dynamic AIDS specification based on the models of Anderson and Blundell (1983, 1984). The AIDS model is seen as a particularly convenient specification with considerable attractive features which, with appropriate transformations and restrictions, can nest a variety of alternative models. In addition, the functional form of the AIDS equations gives an arbitrary first-order approximation to any demand system, satisfies the axioms of preferences exactly, permits perfect aggregation over consumers, and allows for simple linear estimation methods and the imposition of linear restrictions to test homogeneity and symmetry.

However, the system approaches of chapters 5 and 6 rests on an *a priori* endogenous-exogenous division of variables that may be questionable, and the time series included in the systems are nonstationary. Hence, the estimation results obtained from these models can be deemed spurious and the statistical inference invalid, if no cointegrated relationship(s) are found linking the variables of these specifications. In the presence of non-stationary time series and potential feedback effects, an efficient approach for estimating long-run relationship(s), must be a system of equations which allows for all variables to appear as dependent variables, and for appropriate cointegration analysis. An econometric methodology with these characteristics was first proposed by Sims (1980) and is used to specify the models of chapter 7. The econometric methodology applied to the VAR models of this chapter draws extensively on the concepts and techniques showed in Engle and Granger (1987, 1991), Granger (1988, 1997), Harris (1995), Johansen (1988, 1996), Johansen and Juselius (1990), Pesaran (1998) and Pesaran *et al.* (1996).

No economic analysis based on quantitative specifications is complete without the examination of their forecasting performance. The predictive ability of the econometric models used is evaluated and compared following the views and procedures described in such studies as Engel and Yoo (1987), Fair (1986), Granger (1981), Granger and Newbold (1986), Newbold and Bos (1994) and Clements and Hendry (1998). All the estimations, statistical tests, inference and forecasting procedures were computed with Pesaran and Pesaran's (1997) *Microfit 4.0.*

1.5. OVERVIEW OF THE CHAPTERS

Chapter 2 explains, by means of basic statistics, graphs and tables, the evolution of the UK tourism demand for its southern neighbouring countries over the sample period 1969-1997. The analysis of the data in this chapter is used in subsequent chapters to provide a basis for the characterisation of relevant variables and help the economic interpretation of the results obtained.

Chapter 3 begins with an overview of early empirical research in tourism analysis explaining and critically evaluating the econometric models that have been used to estimate tourism demand. Although some researchers for example, Little, 1980 and Witt, 1980 have attempted to introduce dynamics by including lagged variables in their otherwise static models, the literature shows that the large majority of investigators such as, Gray (1966), Artus (1972), Barry and O'Hagan (1971), Jud and Joseph (1974), Kliman (1981), Lin and Sung (1983), Papadopoulous and Witt (1985), Gunadhi and Boey (1986), have used a static single equation approach to model an origin's tourism demand for one or more destinations. The chapter examines the estimation results obtained from different static single equation models, focusing on the comparison of alternative specifications of the UK demand for tourism which are estimated using different definitions for the variables and different functional forms. The analysis shows that slightly different models can produce considerably different estimates, thereby providing inconsistent results upon which no reliable conclusions can be based. The disparities seem to emerge from the lack of a sound empirical methodology and/or a consistent theoretical framework within which plausible consumer behaviour hypothesis can be fully integrated and tested. Modelling procedures constrained by these deficiencies are bound to produce inadequate empirical specifications which generate biased and inconsistent estimation results.

In chapter 4, a more reliable approach within the single equation framework is considered by the derivation of an error correction specification, which integrates the dynamic dimension of tourism demand behaviour absent from the static version. Ho·.. the inter-temporal nature of tourism demand is not the only feature missing from static single equation models. The behavioural assumptions identified by the utility maximisation hypothesis and the interdependencies among competing destinations are also neglected and cannot be fully integrated and tested within a single equation framework.

Chapter 5 examines the UK demand for tourism within a system of equations approach based on the AIDS model. This model allows for the integration and testing of the utility maximisation hypothesis and for the estimation of cross-equation effects in a way not possible with other alternative functional forms. The AIDS model is formulated with the introduction of some innovations: the concept of neighbourhood between origin and destinations and among destinations themselves, which is believed to be relevant in the explanation of the competitive behaviour of destinations; the concept of development transition periods, which appears to affect the destinations' ability to capture increasing foreign tourism receipts in different ways; and the consideration of a non-constant coefficient of the expenditure explanatory variable, which is believed to change due to factors that modify the political and economic relationships between the countries involved. Unlike earlier studies using the orthodox static AIDS approach, from which the findings appeared to reject utility theory hypotheses, the "unorthodox" model of chapter 5 is well defined, data-coherent and theory-consistent.

Nevertheless, it is possible that current budget shares of the UK tourism demand for its southern neighbours depend not only on current prices and expenditure levels, but also on the extent of consumption desequilibrium in previous periods. Tourists' preferences may have been unstable and the parameters of their utility function may have shifted over time. In this case, a short-run dynamic mechanism, taking account of the adjustment of demand towards its long-run equilibrium, ought to be considered. These inter-temporal aspects of tourists demand behaviour have been largely ignored in the literature, and previous studies considering system of equations approaches generally concentrate on purely static models in order to test the assumptions of utility theory. More recent research has recognised the importance of dynamics in tourism demand analysis but studies including these aspects in a system of equations framework are still rare. The derivation and estimation of a dynamic AIDS model in chapter 6 contributes to fill this gap.

The dynamic model of chapter 6 permits the estimation of separate long- and short-run effects that changes in prices and expenditure have on UK tourism demand. The long-run estimates obtained from this model are similar to those obtained from the static version and seem to confirm the existence of a steady state equilibrium relationship between the UK tourism demand and its determinants. Nevertheless, the AIDS systems of chapters 5 and 6 assume that the current levels of the regressors in their equations are exogenously determined, although this might not be the case. Furthermore, the time series used in the estimations are nonstationary and, unless the variables in the equations are cointegrated, this may give rise to spurious results. In chapter 7, cointegration estimation techniques are applied to a VAR system of equations, permitting both to overcome problems of identification arising from unfounded assumptions on the endogenous/exogenous nature of the variables involved, and to confirm the existence of structural relationships between the UK tourism demand and its determinants for each destination.

Chapter 8 carries out a comparative quality evaluation of the econometric models of chapters 4 to 7, by contrasting their relative forecasting ability over the out-of-sample period 1994-1997 and analysing reasons for their different performances. The analysis establishes that although all models are good forecasters, the cointegrated VAR model over-performs the others. Chapter 9 presents a summary of the main findings and puts forward some general conclusions, which might have interesting implications for policy purposes and future research in tourism demand analysis.

CHAPTER 2

UK TOURISM DEMAND FOR FRANCE, SPAIN AND PORTUGAL SINCE 1969: FACTS AND FIGURES

2.1. INTRODUCTION

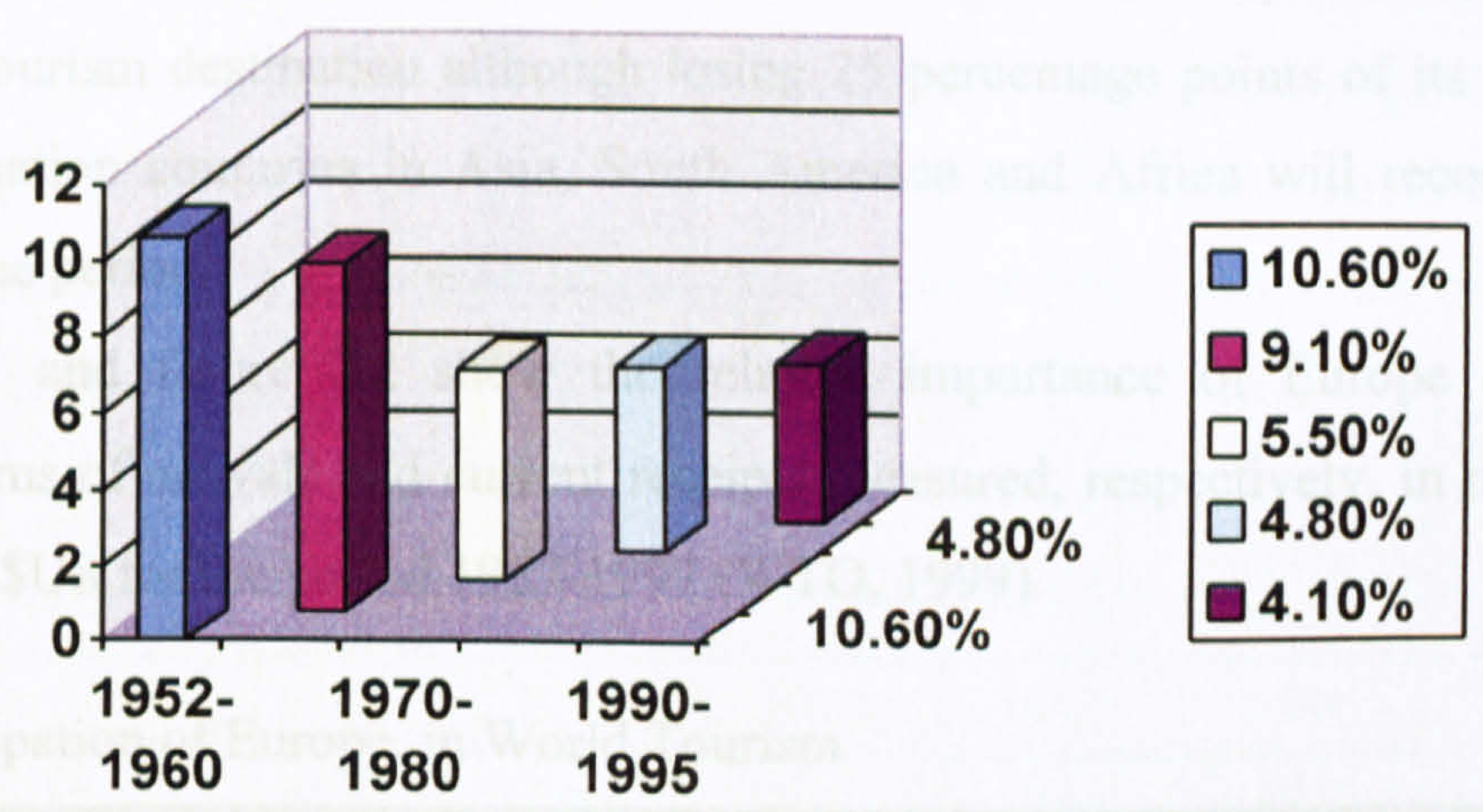
The objective of this chapter is to examine the UK tourism demand for France, Spain and Portugal in the period 1969-1997, using data collected from several sources. The descriptive analysis of the data aims, on the one hand, to give a general view of the UK tourism demand trends within the world and European contexts and, on the other hand, to examine the relevance of tourism destinations such as France, Spain and Portugal relative to other destinations in world and European terms. The analysis focuses on the significance of UK tourism demand relative to that of other origins for France, Spain and Portugal, and on the importance for UK tourists of these countries relative to other tourism destinations. Knowing when, where and how much time and money UK tourists have been spending on their holidays, helps to provide a context for the quantitative approach adopted in subsequent chapters. Indeed, the appropriate characterization of relevant variables, the definition of econometric relationships and the interpretation of results are generally grounded on a good understanding of trends, features and facts affecting the behaviour of the economic time series under study. A thorough examination of available information is therefore, a *sine qua non* for an accurate and reliable quantitative analysis of the UK tourism demand for France, Spain and Portugal.

The structure of the chapter is as follows. Section 2.2 provides an outline of the behaviour of world, Europe and UK tourism demand over time. It also analyses the relative importance of UK demand in world and European terms. Section 2.3 investigates the world and Europe's most important destinations and considers the position of France, Spain and Portugal relative to other destinations in world and European terms. Section 2.4 analyses UK tourism demand for this set of countries in visits and expenditure terms. Section 2.5 concludes.

2.2. WORLD, EUROPEAN AND UK TOURISM DEMAND

Tourism growth has been impressive during recent years. The number of tourist arrivals in all countries increased from 69 million in 1960 to 617 million in 1997 (WTO, 1999). In spite of the slowdown in the growth rate of arrivals since 1960 as depicted in figure 2.1, this variable increased by a factor of 23 between 1950 and 1997.

Figure 2.1: Tourists Arrivals Growth Rate 1952-1995



Source: Yearbook of Tourism Statistics, 1995-1999, WTO.

Although economic and political factors such as the oil crises in the 1970s, the economic recession in the 1980s, and the political instability in the 1990s (Gulf crisis in 1990-91 and war in Yugoslavia in 1993-95) may have had a negative influence on tourism flows, world tourism has continued to grow over the last decades and will continue to grow in the next. Indeed, the World Tourism Organisation (WTO, 1999) predicts an increase of its growth rate in the years 2000-2010 and that by 2020, 1.6 billion tourists visiting countries abroad annually will spent around 2,000 billion $US.

With 15% of the world population and one-third of world GDP, Europe is the largest participant in world tourism, receiving 60% of total arrivals and 52% of total tourism receipts in 1998 (Tourism highlights, WTO, 1999). For decades, Europe has been the world's leading tourism contributor and its share of the global tourism market is still the largest compared with that of other regions, although its growth rate has been declining in recent years. Until 1980, international tourism was typically a North American and European phenomenon but since then, new sources of demand and new competitive destinations have been playing an important role in challenging Europe's leading position. Possible causal factors for this change include the lower growth rate of long distance air travel fares compared with short distance fares; competitive prices and political and economic stability generally offered by new long haul destinations; tourist saturation and environmental degradation in some traditional European destinations and the slowdown of economic growth in main origin countries. However, according to the WTO predictions for the next decade, Europe remains the world's most important tourism destination although losing 25 percentage points of its 1970's share, while new destination countries in Asia, South America and Africa will record substantial growth in the same period.

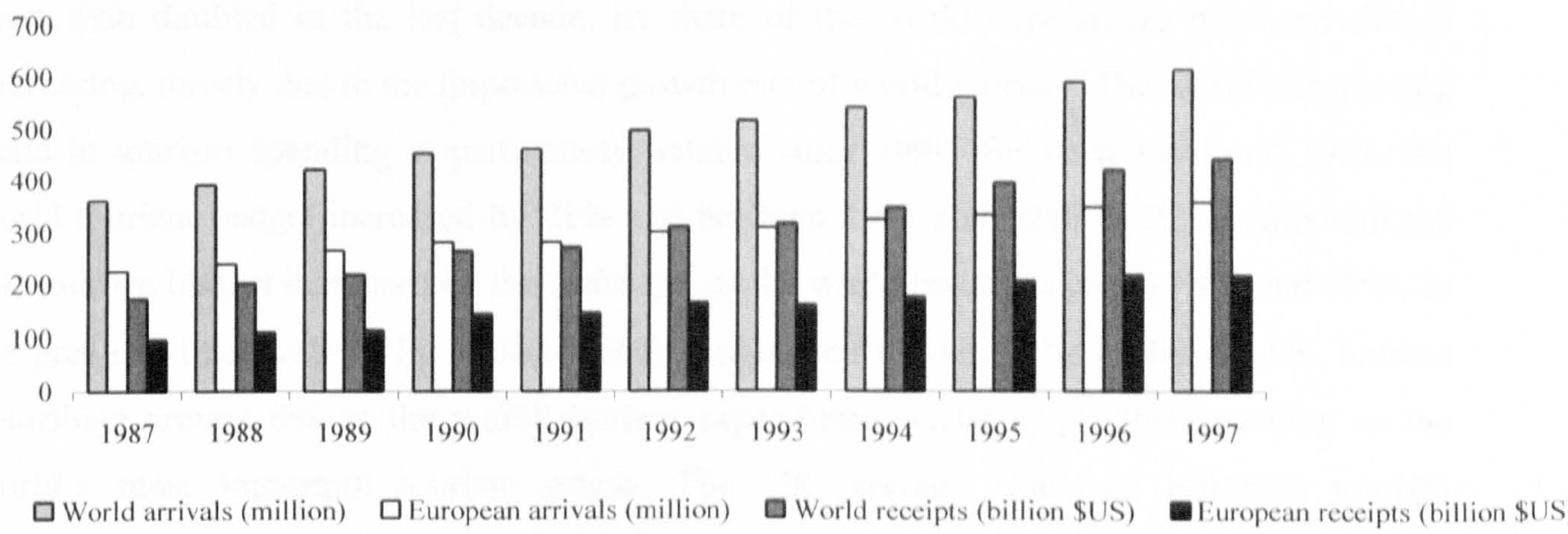
Table 2.1 and figure 2.2 show the relative importance of Europe as a tourism destination in terms of arrivals and current receipts measured, respectively, in million and in thousand million $US for the period 1987-1997 (WTO, 1999).

Table 2.1. Participation of Europe, in World Tourism

	ARRIVALS (million)					RECEIPTS (billion US$)				
	1987	1990	1993	1996	1997	1987	1990	1993	1996	1997
World	362	456	519	595	617	175	266	321	425	447
Europe	226	282	312	351	361	98	147	163	220	218
Europe/World (%)	62.4	61.8	60.1	60.0	58.5	56.0	55.3	50.8	51.8	48.8

Source: Yearbook of Tourism Statistics (1987-1999), WTO

Figure 2.2: World and European arrivals and receipts 1987-1997



Source: Yearbook of Tourism Statistics (1987-1999), WTO

In tourism terms, Europe holds the world's leading position not only as a destination but also as an origin. Around 50% of the world expenditure on tourism originates in Europe and four out of the five most important tourism origins are European countries. Table 2.2 shows the international tourism current expenditure of the world, Europe and that of the five most important origins for the period 1987-1997 (WTO, 1999).

Table 2.2: International tourism current expenditure (billion US$)

	1987	1990	1993	1995	1996	1997
World	175	266	321	399	425	447
Europe	87	135	145	199	208	217
USA	29	37	41	46	48	51
Germany	25	34	41	52	51	46
UK	12	19	19	24	25	28
France	8	12	13	16	18	17
Italy	5	14	14	12	16	16
Europe/World (%)	50%	51%	45%	50%	49%	49%
UK/World (%)	7%	7%	6%	6%	6%	6%
UK/Europe(%)	14%	14%	13%	12%	12%	13%

Source: Yearbook of tourism statistics (1987-1997), WTO

As an origin country generating receipts for other countries, the UK has always been in an important position, both in European and world terms. Although UK tourism expenditure more than doubled in the last decade, its share of the world expenditure has been slowly decreasing, mostly due to the impressive growth rate of world tourism. The world's increasing trend in tourism spending is particularly notable since 1990. Between 1990 and 1993, the world tourism budget increased by 21% and between 1993 and 1996 by 32%. Although the UK tourism budget increased by the same rate as the world budget between 1993 and 1996, in the previous three years, UK tourist spending remained the same. Nevertheless, UK tourists contribute around 6% of the world tourism expenditure, securing the third position as the world's most important tourism origin. The UK average share of Europe's tourism expenditure has been relatively stable and around 13% since 1985. This share indicates the UK as the second most important tourism contributor of Europe.

As shown in table 2.2, roughly half of the world tourism budget has been spent annually by European tourists since mid 1980's. Germany and the UK are the most important European origins and are the second and third origins, respectively, in world terms. Germany and the UK represent more than one-third of Europe's tourism budget. However, since 1995, Germany's tourism spending has been decreasing in both relative and absolute terms. This fact, compounded with the growth rate of UK tourism expenditure between 1987 and 1997 (133%), which largely surpassed that of Germany (85%), indicates that the UK may become the most important origin in Europe.

For a comprehensive analysis of tourism demand behaviour it is also important to investigate how tourists distribute their budgets around the world and which countries are preferred destinations. In the next section, we investigate these aspects, particularly focusing on the importance of France, Spain and Portugal as holiday destinations for European tourists.

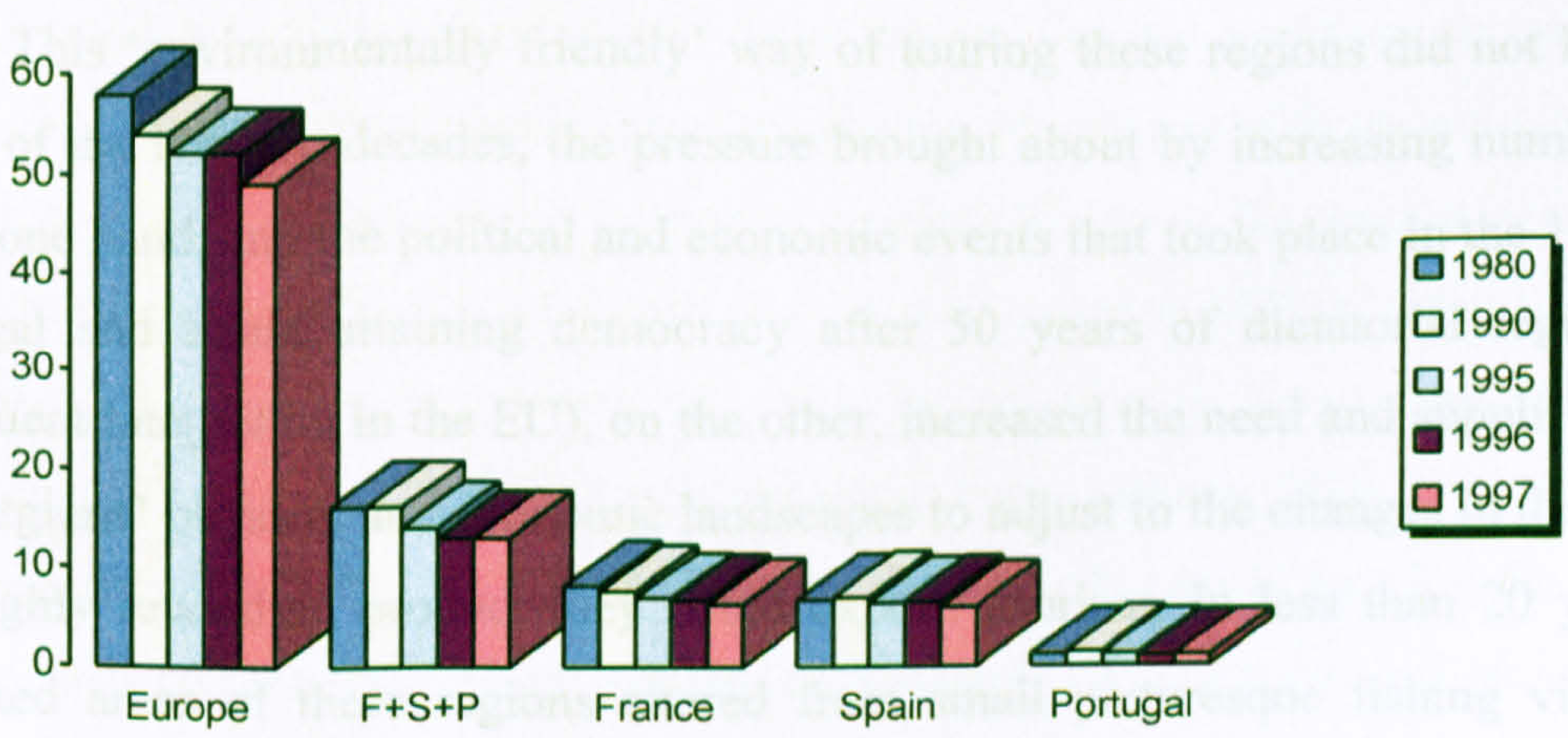
2.3. THE IMPORTANCE OF FRANCE, SPAIN AND PORTUGAL AS DESTINATIONS FOR EUROPEAN TOURISTS

In the late 1990s, European countries were destinations for 60% of world tourists and recipients of half their tourism budgets. However, Europe's share of world tourism has been declining continuously in the last two decades. Since 1980, European destinations have lost around 7% of arrivals and 12% of receipts. Yet, France, Spain and Portugal do not seem to follow Europe's decreasing trend. On the contrary, these destinations' share of world tourism is fairly stable, particularly in terms of arrivals. France, Spain and Portugal together represent close to 19% of world arrivals and 13% of world receipts. These destinations' share of Europe's tourism flows is also impressive, with around 33% of total arrivals and more than 25% of total receipts. The relevance of these destinations is illustrated in table 2.3 and figure 2.3 (WTO, 1999 and WTO Europe, 1997) which show Europe, France, Spain and Portugal's shares of world tourist arrivals and tourism receipts.

Table 2.3: Europe, France, Spain and Portugal's shares of world tourism (%)

	ARRIVALS					RECEIPTS				
	1980	1990	1995	1996	1997	1980	1990	1995	1996	1997
Europe/World	65.1	61.8	59.6	58.9	58.5	60.3	55.3	51.5	51.8	48.8
France/World	10.6	11.5	10.6	10.5	10.8	8.1	7.6	6.9	6.7	6.3
Spain/World	13.4	7.5	7.0	6.9	7.0	6.7	7.0	6.4	6.5	6.0
Portugal/World	1.8	1.8	1.7	1.6	1.6	1.1	1.4	1.1	1.0	1.0
(France+Spain+Portugal)/World	25.8	20.8	19.3	19.0	19.4	15.9	16.0	14.4	14.0	13.3
(France+Spain+Portugal)/Europe	38.4	33.6	32.4	32.3	33.4	26.6	28.8	27.6	27.5	27.1

Figure 2.3: Europe, France, Spain and Portugal's shares of world tourism current receipts



Source:Destination country totals according to a single indicator - arrivals of tourists and tourism receipts (1987-97), WTO and Trends of Tourism Movements and Payments (1980), Europe, WTO

Around 80% of tourists visiting European countries originate from Europe itself. Yet, tourism flows are not evenly distributed across European territory but are instead highly concentrated. In 1992, the twelve European Union (EU) members accounted for 38% of world tourism. With the subsequent membership of Austria, Finland and Sweden, the EU share of world current receipts increased to 41%, reinforcing its position as the most concentrated tourist area in the world. According to the OECD (1996), tourism receipts represent a substantial share of total exports in some EU countries. For example in 1994, the tourism's share of total export was 25% in Greece, 18.4% in Spain and 15% in Portugal. However, for Germany and the UK, these shares are less than 4%. Thus, it is apparent that tourism does not have the same relevance for all EU members. Yet, for some EU countries, tourism is a very important part of their national economies and in some regions of these countries, the main source of income and employment.

In the cases of Spain and Portugal, the impact of tourism is particularly impressive at the regional level. For instance, in the early 1970s, regions like the Algarve in Portugal and Costa del Sol in Spain were underdeveloped areas highly dependent on agriculture and fishing activities, lacking any significant industrialisation, with precarious networks of roads and means of transportation and communications, low income levels and standards of education, health provision and other indicators of social welfare. A typical tourist who visited these regions at that time, usually had to walk long distances to buy provisions in the local market, endure long queues to get daily bread, and go to the local post office to make a telephone call.

This 'environmentally friendly' way of touring these regions did not last long. In the course of the last two decades, the pressure brought about by increasing numbers of tourists, on the one hand, and the political and economic events that took place in the 1970s and 1980s (Portugal and Spain attaining democracy after 50 years of dictatorial regimes, and their subsequent integration in the EU), on the other, increased the need and supplied the means for these regions' physical and economic landscapes to adjust to the changes in the demand of the sole highly rewarding product they could export: tourism. In less than 20 years, the more populated areas of these regions altered from small picturesque fishing villages to busy, modern big cities, from hosting a few thousand visitors to accommodating millions, from the sleepy pace of horse drawn wagons to the frantic highway traffic, from the quiet beer and darts playing in the local pub to the crowded and noisy high-tech discotheques. The speed and extent of this adjustment process are visible in these regions, as they are in those other countries for which economic activities associated with tourism are of primary importance. Tourism receipts in these countries give rise to a variety of repercussions, contributing to the transformation of their economies. Indeed, revenue from foreign tourists creates and sustains jobs, generates additional income and alters its distribution across regions. Tourism receipts act as a catalyst to investment and business activity contributing directly and indirectly to regional economic development.

France, Spain and Portugal are important tourism destinations within the EU, accounting for more than 35% of the region's tourism current receipts in 1997. Although these destinations' share of Europe's tourism current receipts decreased from 29% in 1990 to 27% in 1997, their participation in EU receipts grew more than one percentage point in the period, representing more than one-third of total EU tourism current receipts, as shown in table 2.4 .

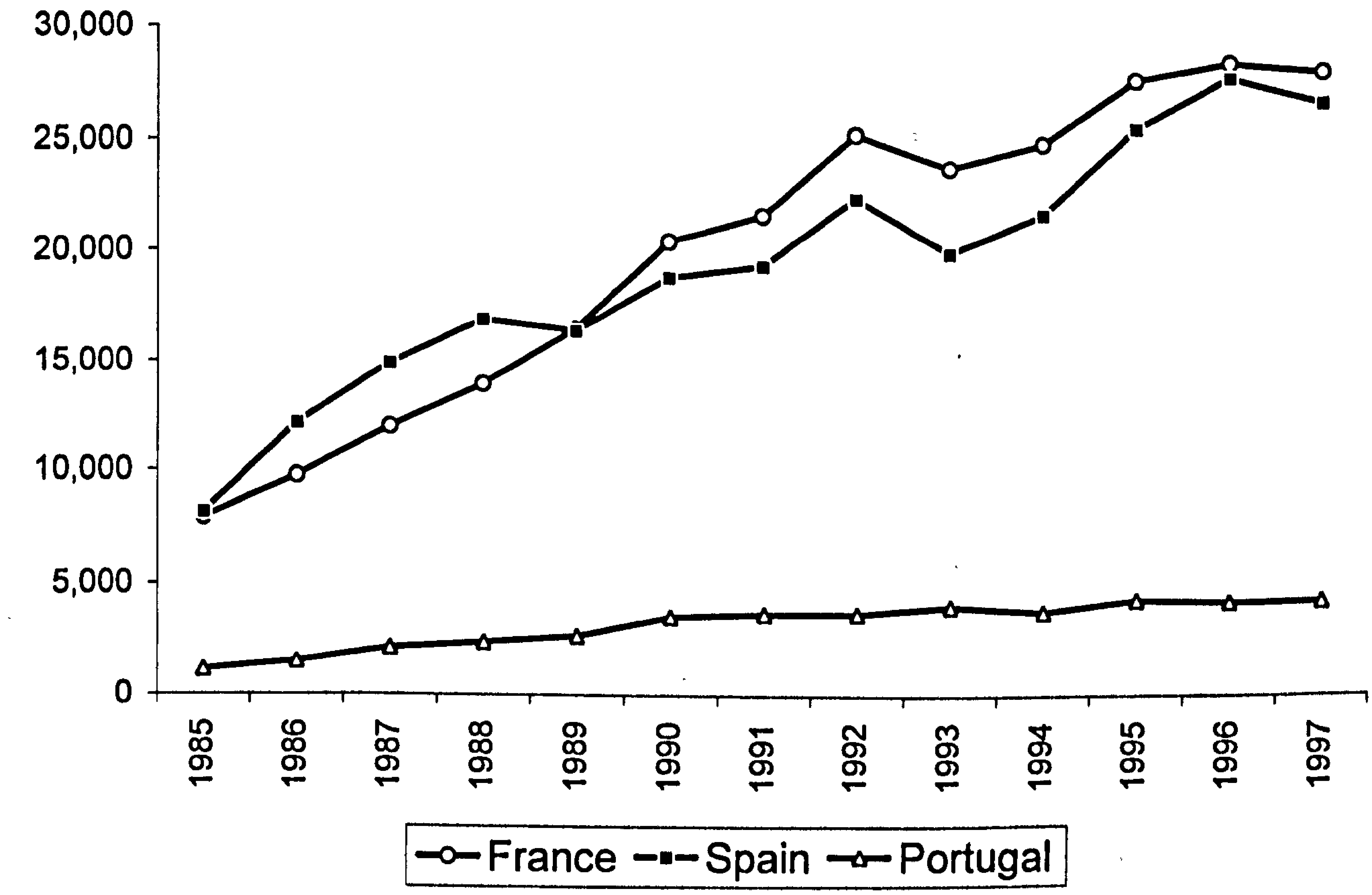
Table 2.4. France, Spain and Portugal share of EU tourism current receipts ($US million)

	1980	1990	1995	1996	1997
EU	50,539	124,075	163,559	168,956	167,070
France+Spain+Portugal	16,416	42,332	57,254	60,294	58,937
(France+Spain+Portugal)/EU	32.5%	34.1%	35.0%	35.7%	35.3%

Source: Destination country totals according to a single indicator – international tourism receipts (1985-1997) and trends of tourism movements and payments (1980), Europe, WTO

France and Spain have always held among the top positions as the world's most preferred destination countries. In terms of arrivals, France is the first destination country and Spain is the third, while in terms of current receipts France is the second destination country and Spain fluctuates between the fourth and the third position (WTO, 1996). In tourism terms, Portugal is a small country compared with its neighbours France and Spain. However, calculations based on WTO data (1999) show that between 1985 and 1997, tourism in Portugal increased faster (except for France in terms of arrivals and for the Netherlands in terms of receipts) than any other Western or Southern continental European country. In fact, tourist arrivals in Portugal increased 60% between 1987 and 1996, only surpassed by France with an increase of 69%, and Portugal's tourism current receipts increased by 275% between 1985 and 1997, only surpassed by the Netherlands with an increase of 288% in the same period. Furthermore, Portugal's share of international tourist arrivals has exceeded 1.5% since 1990, classifying the country as one of the twenty most visited countries in the world (WTO, 1996). Nevertheless, in terms of receipts, the picture is less favourable as Portugal drops to a position below 25 among the most important tourism destinations. Figure 2.4 shows the trends in tourism receipts for France, Spain and Portugal between 1985 and 1997.

Figure 2.4: France, Spain and Portugal tourism current receipts ($US million)



Source: Destination country totals according to a single indicator - international tourism receipts (1985-1997), WTO

Total current receipts in France Spain and Portugal more than tripled between 1985 and 1997. In this period, France's international tourism current receipts increased from 7,943 to 28,009 million $US, implying an average annual growth rate of 19%, Spain's increased from 8,151 to 26,651 implying an average annual growth rate of 17% and Portugal's increased from 1,137 to 4,277 implying an average annual growth rate of 21%. However, if the growth rates are analysed by periods of five years, the picture may appear quite different from that implied by the average for the whole period. For instance, between 1985 and 1990, the average annual growth rate of tourism receipts was 26% for France, 21% for Spain and 35% for Portugal. Yet, for the next five years, these growth rates drop sharply to values of 6% for France and Spain and 4% for Portugal. Between 1995 and 1997 international tourism receipts increased by less than 1% for France, around 2% for Spain and decreased 0.5% for Portugal.

Although it is important to know how France, Spain and Portugal's total current receipts evolve, our interest is to analyse the behaviour over time of receipts from UK tourists and compare their time path with that of total receipts. This is undertaken in the next section.

2.4. UK TOURISM DEMAND FOR FRANCE, SPAIN AND PORTUGAL

The UK has always been a key market for tourism in France, Spain and Portugal. Data for tourist arrivals and tourism receipts can be used to illustrate the importance of UK tourism demand for these destinations relative to that of other origins. For these destinations, table 2.5 shows total arrivals and UK arrivals (million) and current receipts from all international tourists and from UK tourists (billion $US) in the period 1987-1997 (WTO, 1996-1999).

Between 1987 and 1997 in France, total arrivals increased by 81% while UK arrivals increased by 88%. In spite of some fluctuations in the early 1990's, the UK share of total arrivals in France more than recovered its 1987 value in the late 1990's, representing 18% of the total in 1997. In the same period, total arrivals in Portugal increased by 67%. However, UK arrivals in Portugal only increased by 33%. The UK share of total arrivals in Portugal has been relatively stable between 15% and 17%. Given the increase in total arrivals, the stability of the UK share may indicate some diversification of tourism markets for Portugal in the last decade. Spain shows a much slower increase of both total and UK arrivals. While total arrivals in Spain increased by 54% between 1987 and 1997, UK arrivals increased by a modest 9% in the same period. UK arrivals in Spain increased from around 3 million in 1977 to 7.7 million in 1988. However, between 1989 and 1993, this destination lost around 1 million UK arrivals and only recovered its 1987 level in 1996. In the last two decades, total arrivals in Spain never decreased. This, compounded with the fact that the UK share of total arrivals in Spain decreased from 27% in 1987 to 19% in 1997, seems to indicate diversification of origins demanding tourism in this destination, more pronounced than that implied by the numbers for

Portugal. Hence, assuming that arrivals may represent a good proxy for tourism demand, the traditional UK tourists' preference for Spain relative to its neighbours, France and Portugal, seem to have decelerated in the last decade. Furthermore, in terms of UK arrivals, France surpassed Spain for the first time in 1990, and it seems that France's new leading position in the UK tourists preferences has not been reversed.

Table 2.5: Arrivals (million) and current receipts (billion $US) of France, Spain and Portugal

		1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997
France	Total arrivals	37.0	42.7	49.5	52.5	55.0	59.7	60.6	61.3	60.0	62.4	66.8
	UK arrivals	6.4	6.6	7.1	7.0	6.8	8.2	8.2	11.6	11.2	11.5	12.0
	UK/Total.	17%	15%	14%	13%	12%	14%	14%	19%	19%	18%	18%
	Total receipts	11.9	13.8	16.2	20.2	21.4	25.1	23.6	24.7	27.5	28.4	28.0
	UK receipts	1.5	1.7	2.1	2.6	2.8	3.0	2.9	3.0	3.3	3.1	3.7
	UK/Total	13%	12%	13%	13%	13%	12%	12%	12%	12%	11%	13%
Spain	Total arrivals	27.7	29.8	32.5	34.1	34.2	36.5	37.3	43.2	39.3	40.5	43.4
	UK arrivals	7.6	7.7	7.4	6.3	6.2	6.5	6.5	7.7	8.2	7.6	8.3
	UK/Total	27%	26%	23%	18%	18%	18%	17%	18%	21%	19%	19%
	Total receipts	14.8	16.7	16.2	18.6	19.1	22.2	19.7	21.5	25.4	26.7	26.7
	UK receipts	2.7	3.4	2.8	2.7	2.7	3.1	3.2	3.9	4.5	4.2	4.6
	UK/Total	18%	20%	17%	15%	14%	14%	16%	18%	18%	16%	17%
Portugal	Total arrivals	6.1	6.6	7.1	8.0	8.7	8.9	8.4	9.2	9.5	9.7	10.2
	UK arrivals	1.2	1.1	1.1	1.2	1.3	1.4	1.4	1.4	1.6	1.5	1.6
	UK/Total	20%	17%	15%	15%	15%	16%	17%	15%	17%	15%	16%
	Total.receipts	2.1	2.4	2.7	3.6	3.7	3.7	4.1	3.8	4.3	4.3	4.3
	UK receipts	0.4	0.6	0.5	0.5	0.6	0.7	0.6	0.6	0.8	0.7	0.8
	UK/Total	19%	25%	19%	14%	16%	19%	15%	16%	19%	16%	19%

Source: Yearbook of tourism statistics (1987-1997), WTO and Tourism policy and international tourism in OCDE member countries (1990- 1997), OCDE.

Tourist arrivals or visits, as a measure of tourism demand, can be misleading as their levels may not translate into the effective consumption of tourism in the countries visited. Origins' tourism expenditure in a destination or destinations' tourism receipts from an origin, are generally considered more accurate measures of tourists preferences in accordance to the theoretical micro-foundations of econometric models currently used to analyse tourism demand. Moreover, tourism receipts have a direct economic impact on the destinations' economies and, therefore, constitute a more interesting measure for policy purposes.

In table 2.5, the general trend path of tourism receipts appears to follow that of arrivals although in a less obvious way. Between 1987 and 1997, total tourism receipts more than doubled for Portugal and France while for Spain the increase is less than double. Indeed, total receipts increase by 133% for France, 105% for Portugal and 80% for Spain. In the same period, the receipts from UK tourists increased by 147% for France, 100% for Portugal and only 70% for Spain. Since 1990, UK tourists have been contributing, on average, 12% of the total tourism receipts of France, 16% of the total tourism receipts of Spain and 17% of the total tourism receipts of Portugal. These numbers seem to support the view that UK tourists' preferences are slowly but progressively changing from mainly favouring Spain to favouring France and Portugal as holiday destinations.

In order to understand better how UK tourists distribute their budgets among international holiday destinations, table 2.6. shows UK tourism visits and expenditure in the world, Europe, France, Spain and Portugal for the period 1970 to 1997 (Business Monitor MA6 and MQ6, 1967-1993 and Travel Trends, 1997-1999). For the last three decades, the UK demand for tourism abroad has been growing considerably. Between 1970 and 1998, UK tourist visits abroad increased by a factor close to 6 and UK tourism expenditure abroad increased by a factor of around 50. This shows a much faster increase in expenditure than in tourists numbers and provides an indication of the way in which UK residents have been changing their patterns of tourism consumption over time. In 1970, UK tourists' average spending abroad was £45 per visit; by 1997 this average was £408 per visit. Even taking inflation into account, this increase is still impressive.

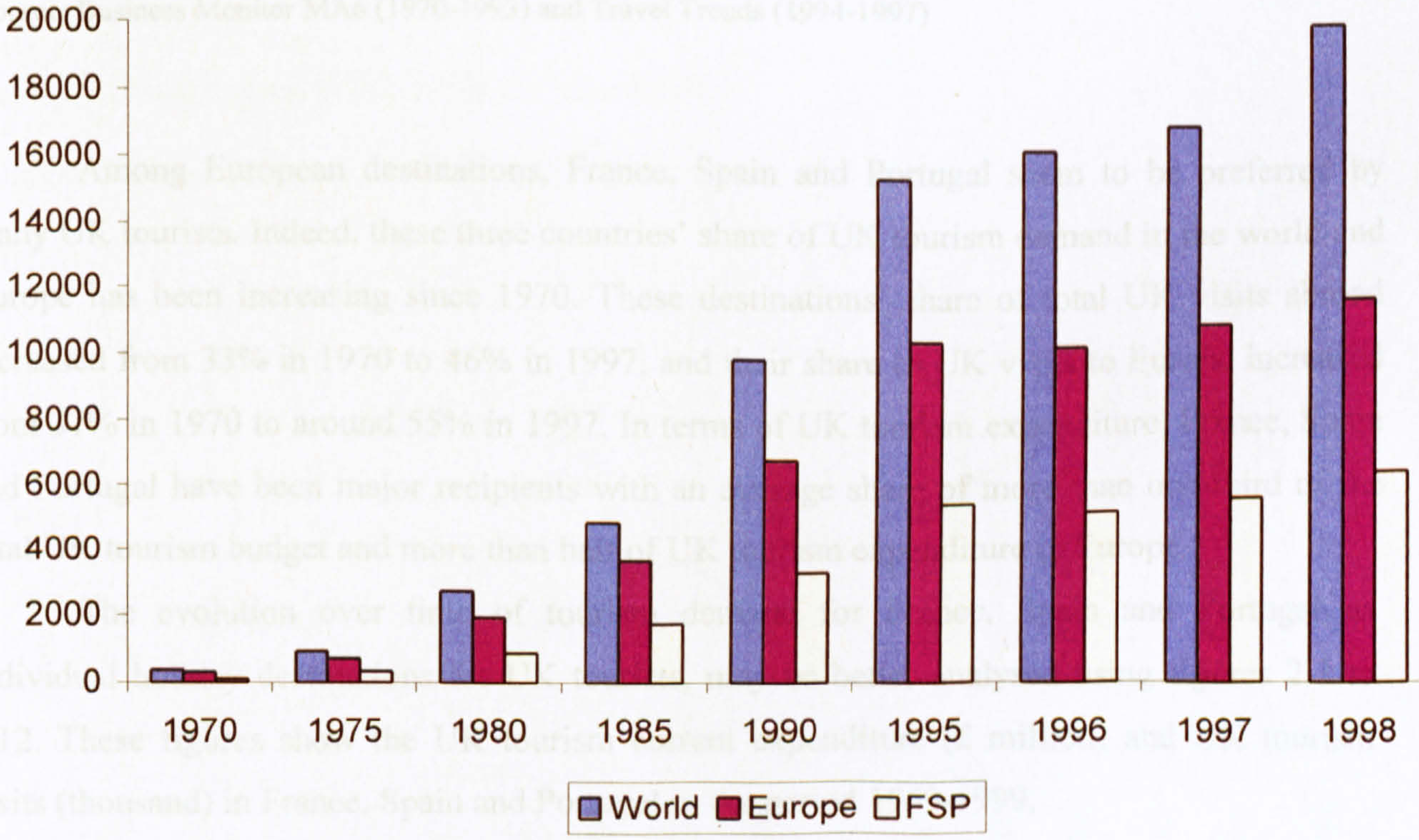
UK tourists' preference for European destinations is confirmed in table 2.6 by Europe's share of UK tourism demand. In 1970, UK visits to Europe represented 90% of total visits abroad and tourism expenditure in Europe represented 83% of the total UK international tourism budget. Although these European shares have been decreasing over the last three decades, Europe's shares of UK tourism are still very important, representing 82% and 64%, respectively, of total visits and expenditure in 1997. Figures 2.5 and 2.6 show the relative magnitudes of UK visits (thousands) and expenditure (£ million) in world, Europe and in France, Spain and Portugal (FSP) as a region.

Table 2.6. UK tourism visits (thousand) and current expenditure (£ million) abroad

	World		Europe		France		Spain		Portugal	
	Visits	Expend	Visits	Expend	Visits	Expend	Visits	Expend	Visits	Expend
1970	8482	382	7662	317	1059	33	1583	73	164	8
1975	11992	917	10283	704	2149	111	2521	207	94	11
1980	17507	2738	14676	1942	3844	375	2617	428	364	67
1985	21610	4871	19181	3687	4523	642	4175	939	709	177
1990	31150	9886	26268	6831	6865	1482	5096	1528	982	306
1995	41345	15386	34418	10422	9645	2107	8239	2877	1211	483
1996	42050	16310	34213	10260	9834	2015	7545	2704	1102	452
1997	45957	17136	37745	10879	11149	2256	8281	2825	1304	492
1998	48800	19900	39500	11700	11518	2663	9650	3236	1299	468

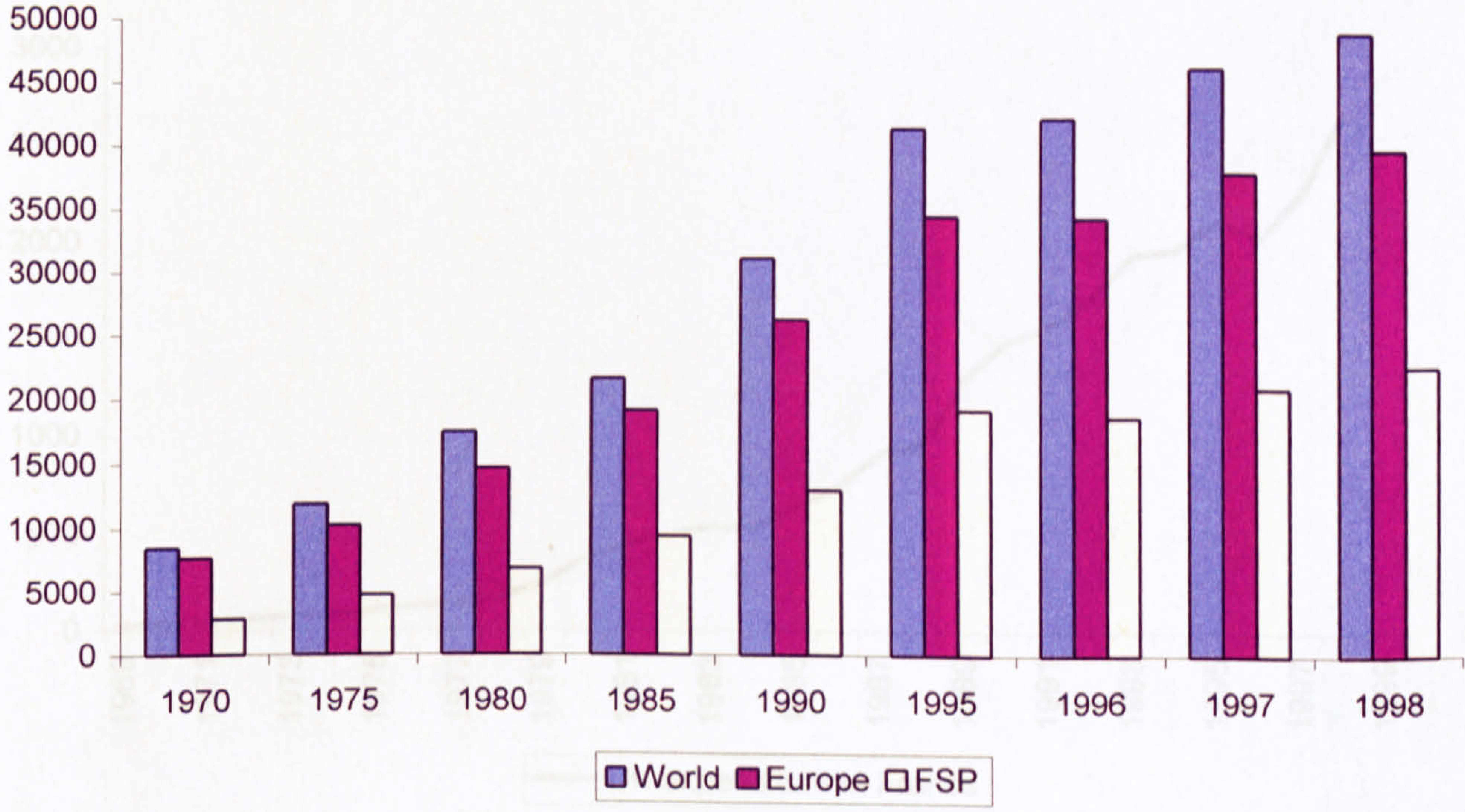
Source: Business Monitor MA6 (1970-1993) and Travel Trends (1994-1997)

Figure 2.5: UK tourism current expenditure in the world, Europe, and in France, Spain and Portugal



Source: Business Monitor MA6 (1970-1993) and Travel Trends (1994-1997)

Figure 2.6: UK tourist visits to the world, Europe and to France, Spain and Portugal



Source: Business Monitor MA6 (1970-1993) and Travel Trends (1994-1997)

Among European destinations, France, Spain and Portugal seem to be preferred by many UK tourists. Indeed, these three countries' share of UK tourism demand in the world and Europe has been increasing since 1970. These destinations' share of total UK visits abroad increased from 33% in 1970 to 46% in 1997, and their share of UK visits to Europe increased from 37% in 1970 to around 55% in 1997. In terms of UK tourism expenditure, France, Spain and Portugal have been major recipients with an average share of more than one-third of the total UK tourism budget and more than half of UK tourism expenditure in Europe.

The evolution over time of tourism demand for France, Spain and Portugal as individual holiday destinations for UK tourists, may be better analysed using figures 2.7 to 2.12. These figures show the UK tourism current expenditure (£ million) and UK tourism visits (thousand) in France, Spain and Portugal in the period 1969-1999.

Figure 2.7: UK tourism expenditure in France (£ million)

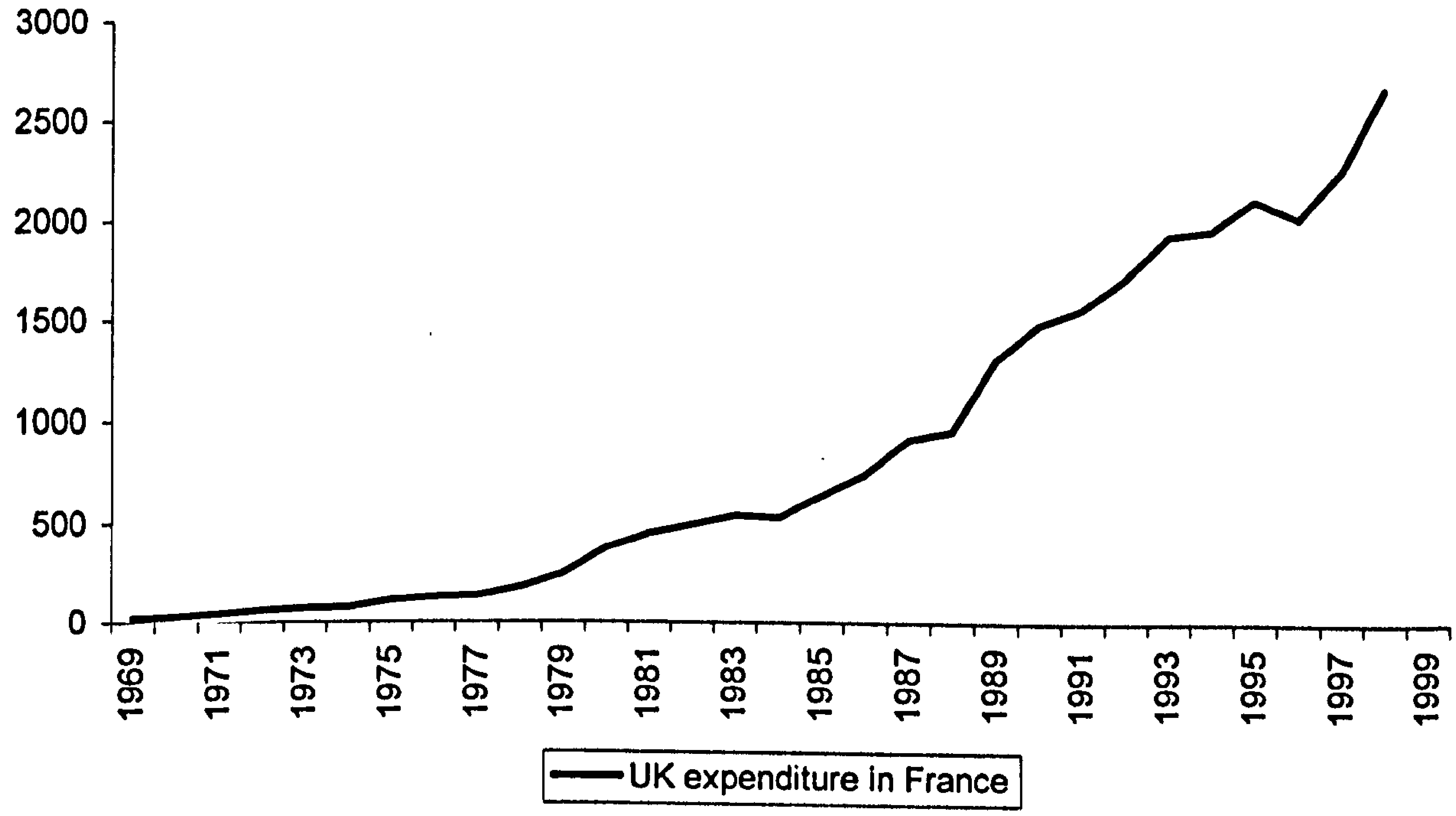
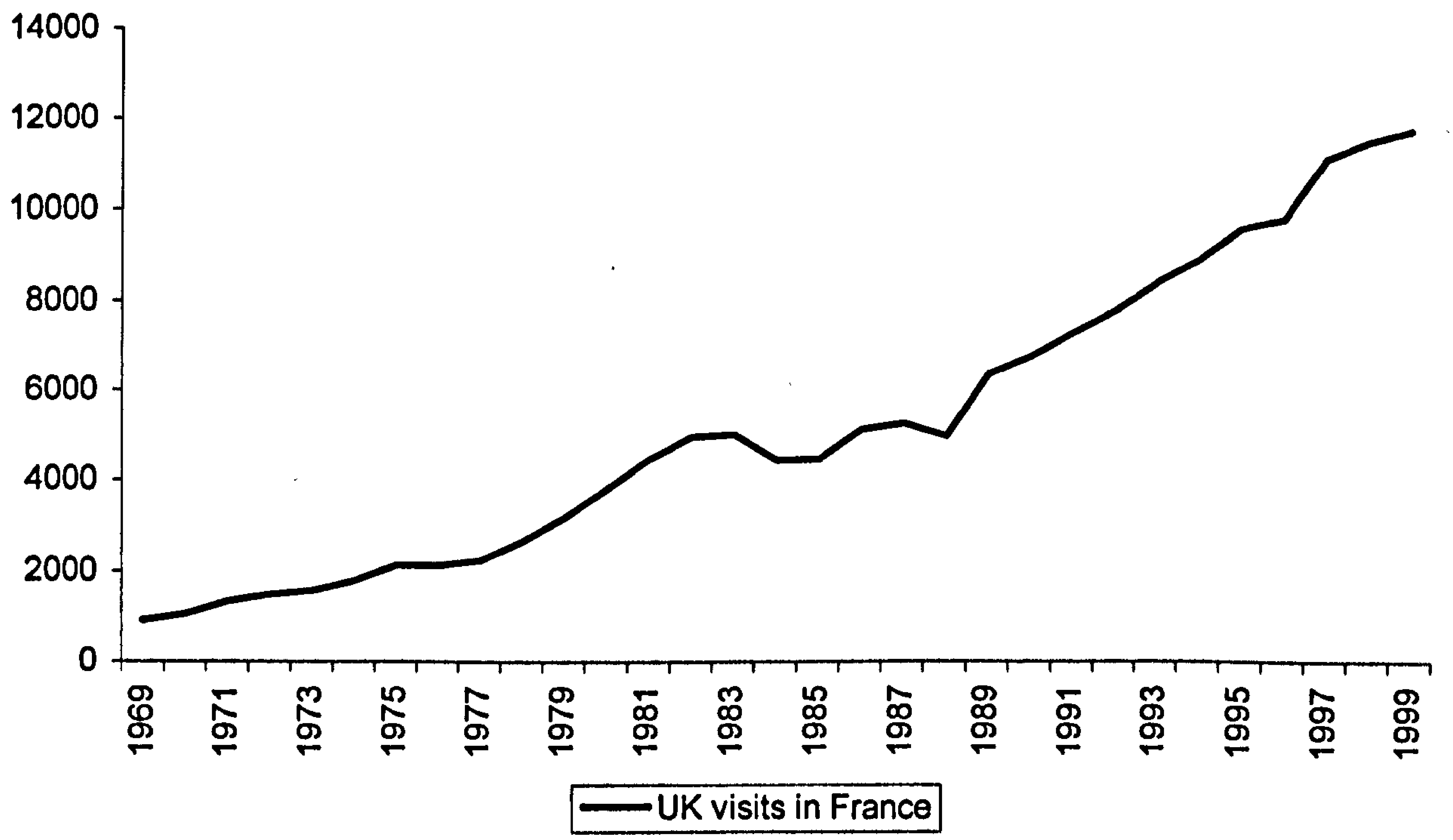


Figure 2.8: UK tourist visits in France (thousands)



Source: Business Monitor MA6 (1969-1993) and Travel Trends (1994-1999)

Figure 2.9: UK tourism expenditure in Spain (£ million)

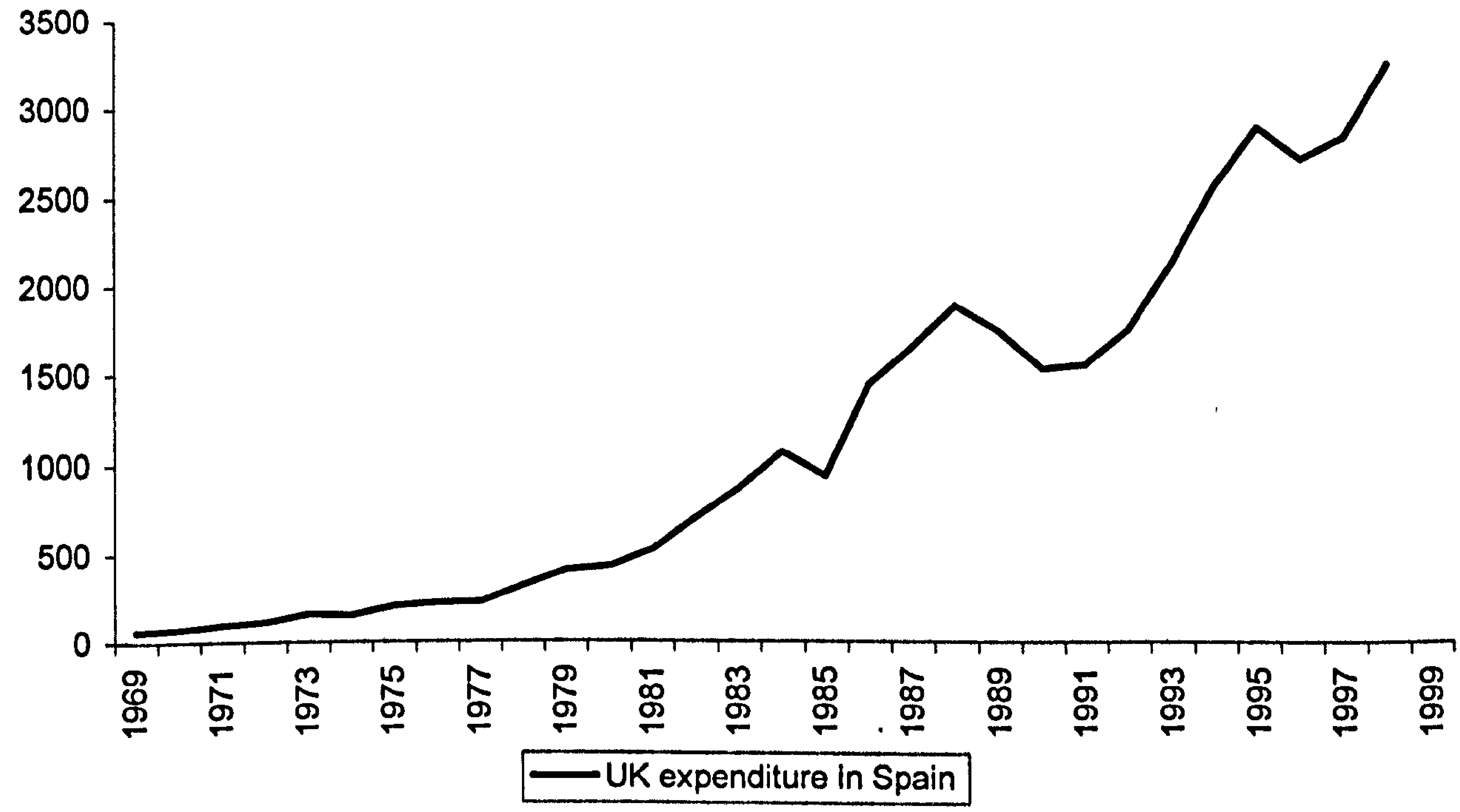
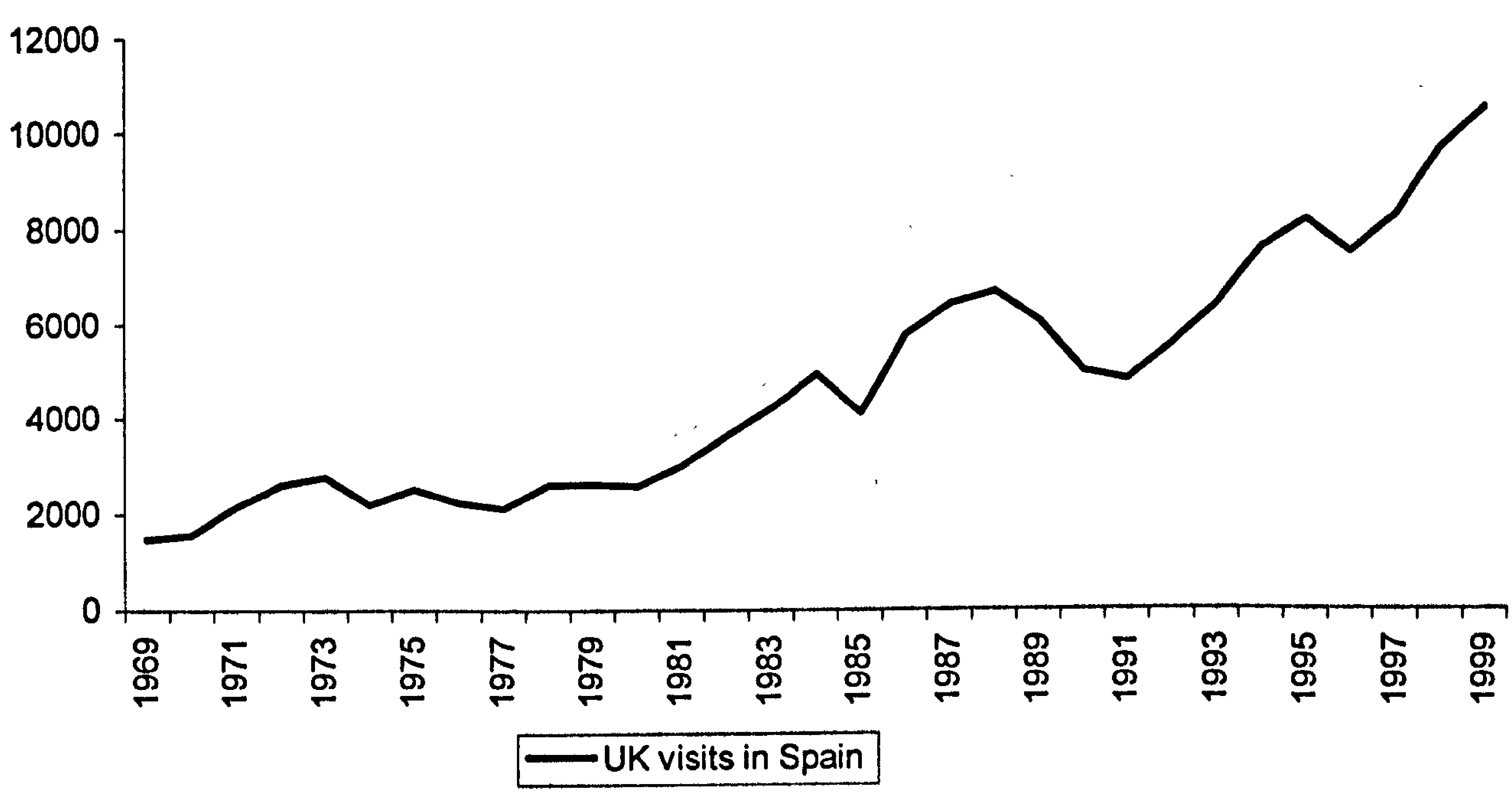


Figure 2.10: UK tourist visits in Spain (thousands)



Source: Business Monitor MA6 (1969-1993) and Travel Trends (1994-1999)

Figure 2.11: UK tourism expenditure in Portugal (£ million)

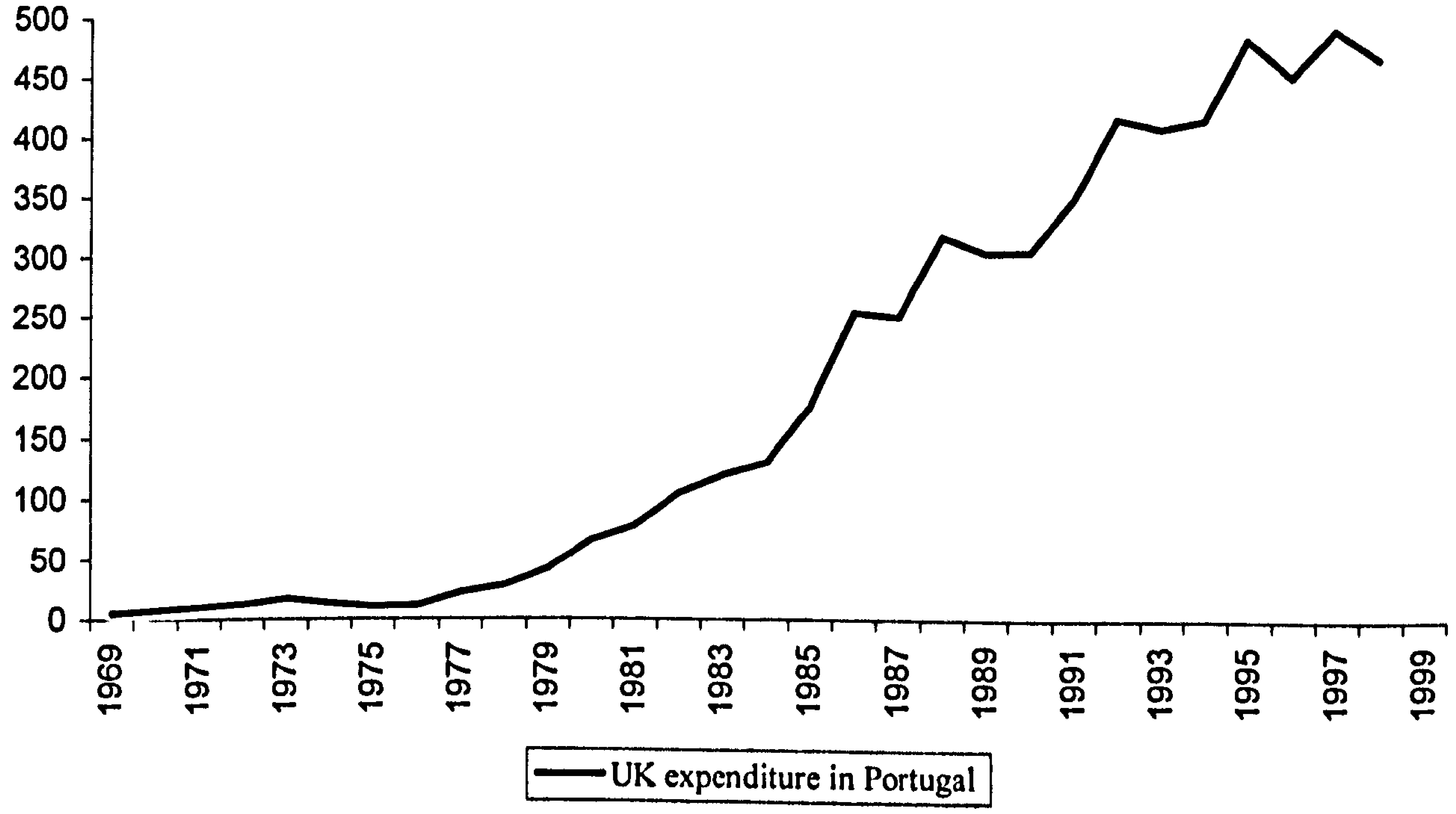
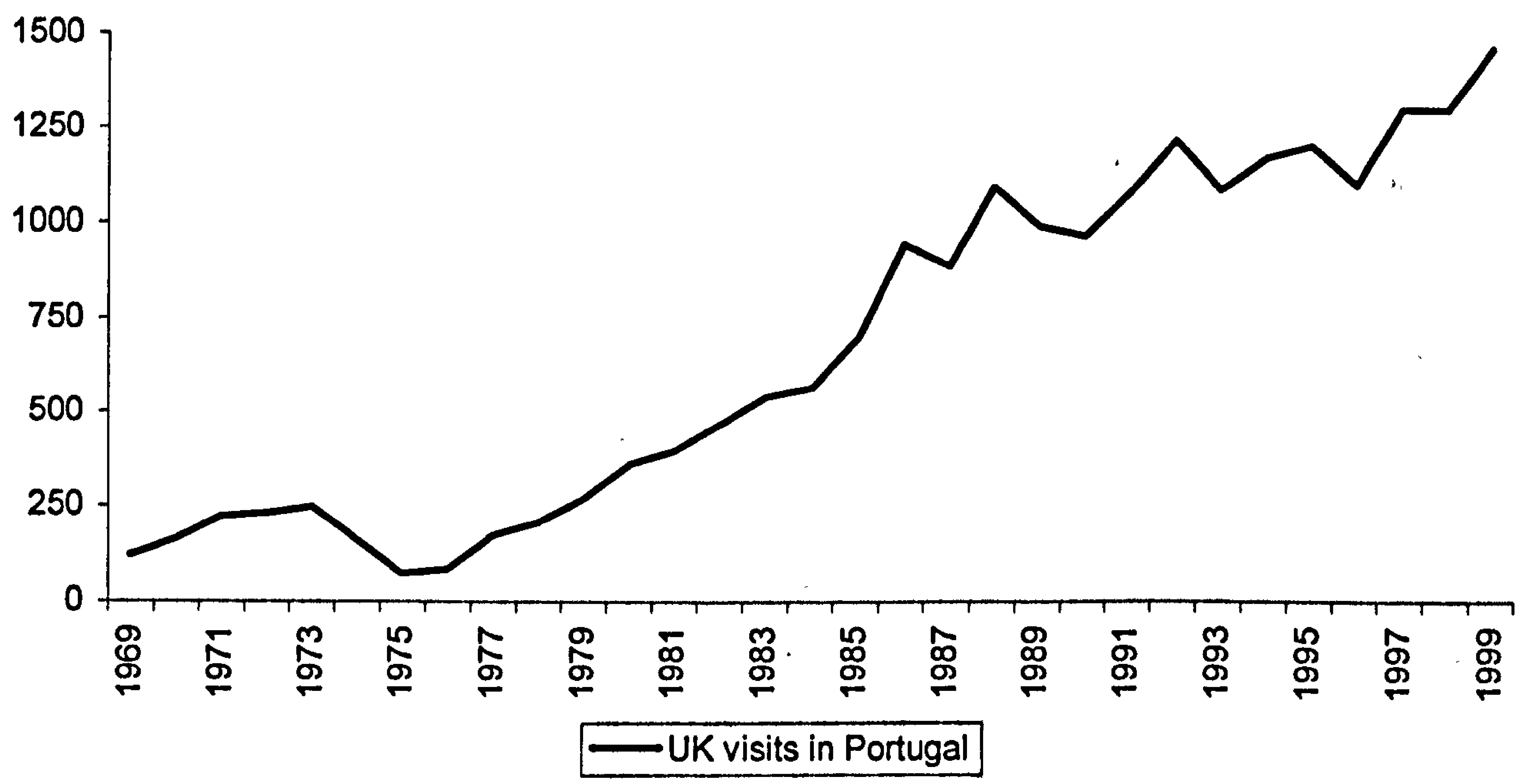


Figure 2.12: UK tourist visits in Portugal (thousands)



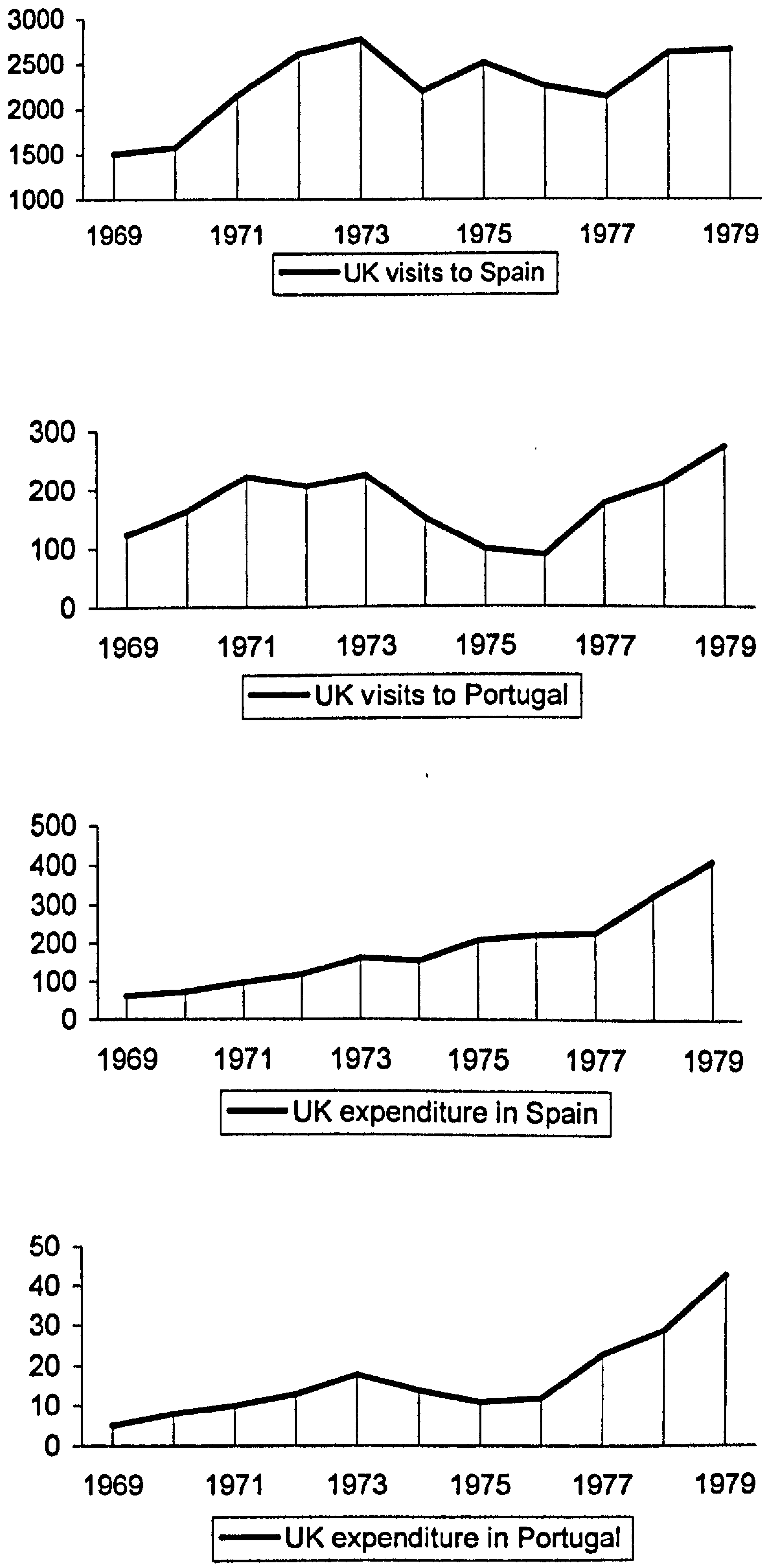
Source: Business Monitor MA6 (1969-1993) and Travel Trends (1994-1999)

Figures 2.7 to 2.12 present increasing trends for both UK expenditure and visits. UK visitors to France increased from around 1 million in 1969 to 12 million in 1999. In the same period, UK visitors to Spain increased from around 1.5 million to 11 million, and to Portugal from little more than 100 thousand to around 1.5 million. A similar trend exists for UK tourism expenditure. The level of UK tourism expenditure in France increased by a factor of 106 between 1969 (£25 million) and 1998 (£2663 million), by a factor of 52 in Spain (from £62 to £3236 million) and by a factor of 94 in Portugal (from £5 to £468 million).

Assuming that France, Spain and Portugal's proximity to the UK gives this region a comparative advantage relative to other destinations in UK tourists' preferences, it is likely that exogenous shocks affecting tourism flows in one country have a significant impact on the demand for the other neighbouring countries. In fact, this appears to be the case for the 1974 revolution in Portugal, the substitution of Franco's dictatorial regime in Spain, and Spain and Portugal's membership of the EU in 1986. The 1974 revolution in Portugal and the political events of 1976-1977 in Spain, led to the substitution of dictatorial regimes by parliamentary democracies in both these countries. Initially, between 1974 and 1979, these events had a negative effect on UK tourism flows to these destinations. However, the consolidation of democracy in both countries and their simultaneous entry into the EU in 1986 seem to have had a positive impact on UK tourism demand in the following decade. Indeed, in 1980-1989, both destinations present almost uninterrupted increasing levels of UK expenditure.

Nevertheless, the 1974 revolution had a major negative impact on the UK demand for Portugal, also affecting Spain. The Portuguese political and social instability of 1974-1976 caused a decrease in UK visits and expenditure in Portugal. In this period, UK demand seems to have been diverted mainly from Portugal to neighbouring Spain. Hence, initially, Spain appears to benefit from the instability in Portugal. However, the subsequent political upheaval in Spain, compounded with the ongoing PREC (communist revolutionary process) in Portugal, diverted a substantial part of UK tourists away from the Iberian Peninsula for the rest of the decade. Only by 1979, did both destinations recover their UK visit levels of 1973. Given the low levels of expenditure and visits of the 1970s, particularly for Portugal, the plots in figures 2.9 to 2.12 do not clearly show these UK tourism demand changes. Therefore, figure 2.13 is constructed to magnify the first decade of the data in order to demonstrate the decrease in UK visits and expenditure in Spain and Portugal, following 1974.

Figure 2.13: UK visits and expenditure in Spain and Portugal 1969-1979



Source: Business Monitor MA6 (1969-1979)

Since the mid 1980's, the number of UK visitors in France has generally exceeded that in Spain by amounts that can reach three million visits. However, the UK tourism expenditure in Spain is generally larger than that in France by amounts ranging between £200 million and £1000 million. This fact illustrates the problems of using tourist numbers as a measure of tourism demand. Although UK tourists visiting Spain are less in number than those visiting France, UK tourism expenditure in Spain generally exceeds that in France, sometimes by substantial amounts. This same feature is detected for Portugal when comparing Spanish and UK visitors and Spanish and UK tourism expenditure in this destination. Spanish tourists come to Portugal in huge numbers but spend very little; UK tourists arrive in smaller numbers but their spending represents a substantial part of Portugal's total tourism earnings.

Smaller numbers of tourists associated with relatively larger expenditure levels, are generally linked to a longer average length of stay and/or a different spending propensity displayed by tourists from specific origins in different destinations. Different behaviours in tourism demand indicated by such measures as the average length of stay, may be due to both individual characteristics of the destinations countries and to specific features of the tourists' demand behaviour. For instance, in the last three decades, the average length of stay in all destinations presents a continuous decrease. Given the general increase in average leisure time across the world and, particularly, in industrialised countries which constitute major sources of tourism demand, the decrease in average length of stay may indicate greater mobility of tourists among preferred destinations. This greater mobility may indicate changes in demand behaviour favouring 'complementary' destinations more than competing ones. Distance between destinations may play a major role in defining these complementarities, from the tourists' point of view. If this is true, then neighbouring countries are likely to attract the attention of cross-border tourists who, having little time to waste, want to spend a part of their holiday in one country and the rest of it in other(s).[0] However, from the destinations' point of view, the important strategy is to keep tourists in their territories for as long as possible. Therefore, the higher mobility of modern tourists may trigger competitive behaviours among neighbouring destinations, more aggressive than those observed between countries far apart.

[0] These are hypothetical assertions which should be supported (or otherwise) by specific research. However, given the increasing average leisure time available to UK tourists and their decreasing length of stay in individual countries associated with increasing budgets allocated to international tourism, it is plausible that UK tourists have been visiting more than one destination during their yearly holidays.

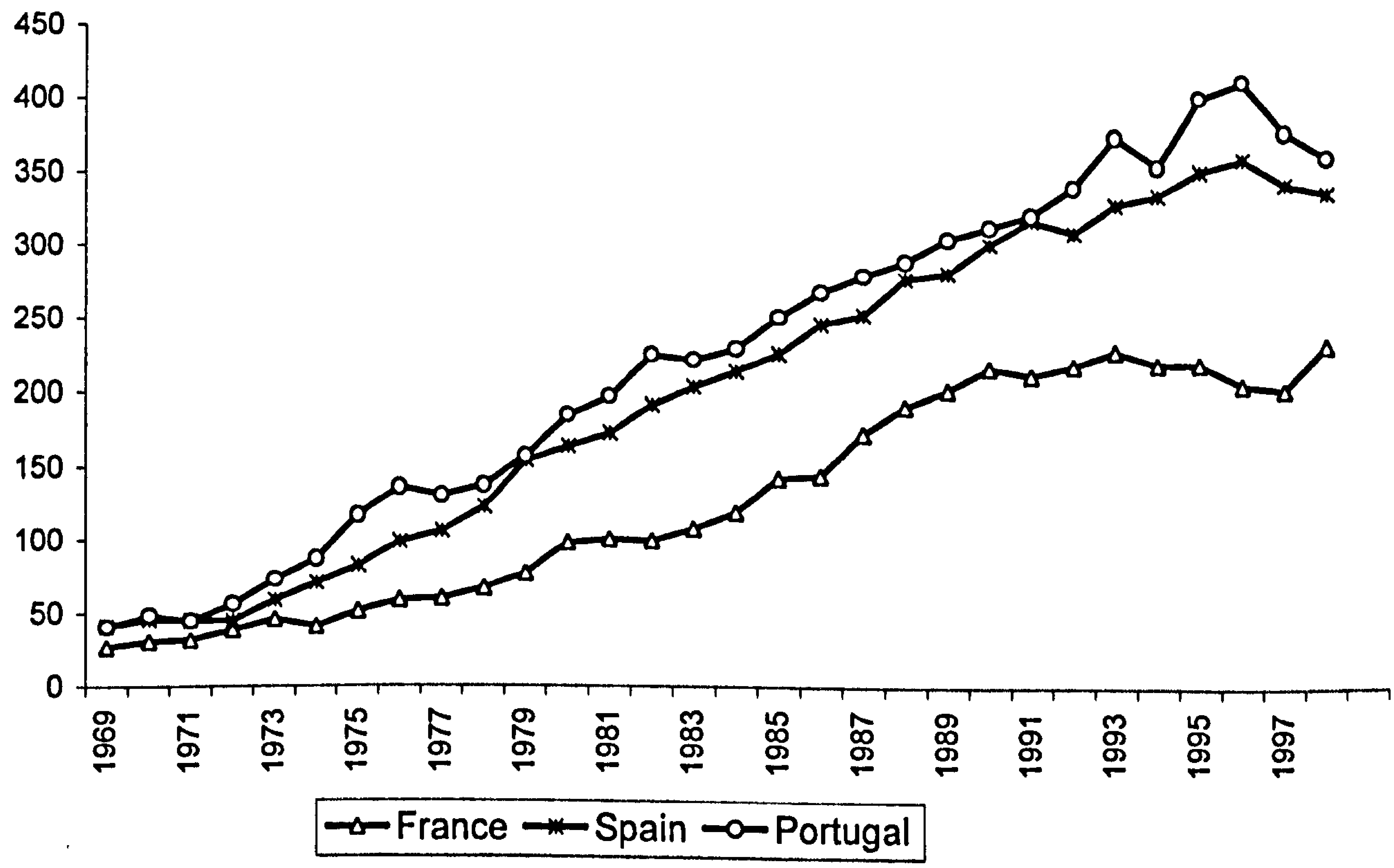
For France, Spain and Portugal, it is interesting to observe that UK tourists' average length of stay, although decreasing in accordance with the general trend observed in all other UK destinations, presents relatively stable values when compared, for example, with the decrease in this measure for Italy or Germany. In fact, UK tourists' average length of stay decreased in France from 8 days in 1975 to 6 days in 1997, in Spain from 13 to 12 days, and in Portugal from 7 to 5 days. However, in Germany this measure fell from 12 to 6 and in Italy from 14 to 9 in the same period (Business Monitor MQ6, 1969-1998).

In general, the longer the average of stay, the smaller the number of visits as the same visit is extended over time. On the other hand, longer stays usually imply more spending. Fewer visits and more spending imply higher expenditure per visit. Therefore, the expenditure per visit is expected to be higher for destinations with longer average length of stay. This may explain why although UK visits to France are larger in number than those to Spain, UK tourism expenditure in Spain generally surpasses that in France.

However, the expenditure per visit in a destination country may reflect more than just an extended average length of stay. In fact, as shown in figure 2.14, the UK expenditure per visit in Portugal is higher than that in Spain or France. Yet, given that UK tourists' average length of stay in Portugal is smaller than in France and Spain, and has decreased most since 1969, and given that costs associated with travelling to any of these destinations are expected to be similar for UK tourists, the levels of expenditure per visit in Portugal must be linked with aspects other than the length of stay.

The main reasons that can explain this apparently odd feature of UK demand for Portugal may be linked with tourism prices and the spending propensity of the visitors. Either tourism price levels in Portugal are substantially higher than those of its neighbours and have been so for the last three decades, or UK tourists are more prodigal when visiting Portugal than when visiting Spain or France. Probably, a combination of these and other factors may explain the differences in UK expenditure per visit in these three destinations. Nevertheless, if prices alone do not explain these differences, then there is scope for a very interesting and innovative field in tourism demand analysis, focusing on the determinants of different spending propensities of tourists from a given origin in different destination countries.

Figure 2.14: UK tourism expenditure per visit in France, Spain and Portugal (£)



Source: Business Monitor MA6 (1969-1993) and Travel Trends (1994-1997)

2.5. CONCLUSION

In the course of the last two decades, the pressure brought about by increasing numbers of tourists and the potential economic benefits and costs that tourism brings to national and regional economies, intensified the need to understand and measure the magnitude and direction of changes in tourism demand behaviour. Whether, how much and for how long tourism receipts contribute net benefits to a region depends, among other things, on the structure of the local economies and on the precise form and scale of the tourism demand for that region. Hence, as tourism emerges as an important factor of development and source of reproducible wealth, it becomes imperative to implement a comprehensive quantitative examination of tourism demand behaviour.

Quantitative empirical studies of tourism demand generally involve the use of econometric models to specify the relationships between the variable measuring demand and its determinants. Econometric modelling generally provides a good basis for economic analysis and policy evaluation. However, appropriate conception and implementation of quantitative analyses requires two basic things: plentiful and reliable data, and adequate methodological tools to build, estimate and test models that can validly describe the past and predict the future of tourism demand levels. Plentiful and reliable data are rarely available to investigators, in particular, in tourism research contexts. Although there has been some improvement recently, the more reliable suppliers of tourism information generally make it available on yearly basis, at aggregate levels, and for relatively short periods of time. Yet, these difficulties should not deter research and although 'big omelettes need big eggs', there is nothing wrong with smaller omelettes as long as they are well cooked.

Whatever the quality and quantity of the information available, the appropriate building of an econometric model and interpretation of its estimates requires a thorough examination of the relevant time series. The data must be inspected in detail, in order to determine how the series behave over time. Data specific features, compounded with knowledge of events which could have had a significant impact in the evolution of the series, can then be integrated in the model building, permitting a more precise specification of the variables interrelationships, within an adequate quantitative framework. Although 'torturing the data until they confess' is not the general idea, every aspect, even minute, can help to construct better models on which reliable inference and forecasting procedures can be based. The analysis of the data in this chapter complies, with these requirements and helps to characterize the relevant variables and define the quantitative framework within which plausible relationships can be modelled and interpreted.

Using the information provided in this chapter, the remainder of the thesis focuses on different methodological approaches to the modelling, estimation and forecasting of tourism demand. Based on these different methodologies, several econometric models are constructed to analyse the behaviour of UK demand for tourism in France, Spain and Portugal over the period 1969 to 1997: a static single equation approach is considered in chapter 3; a dynamic single equation model is estimated in chapter 4; static and dynamic almost ideal demand systems (AIDS) of equations are used in chapters 5 and 6 and a vector autoregressive (VAR) approach is explored in chapter 7. Chapter 8 evaluates the ability of these models to predict the UK tourism demand levels in these destinations. Chapter 9 concludes.

CHAPTER 3

STATIC SINGLE EQUATION MODELLING OF TOURISM DEMAND

3.1. INTRODUCTION

Econometric modelling of tourism demand is not an easy task. The reasons include the complexity of the motivational structure underlying the decision-making process; the multiplicity and heterogeneity of the products and services supplied; the existence of endogenous regressors; the fact that transportation plays a role in tourism consumption; the inter-temporal dependence of current demand; the existence of qualitative factors, often non-measurable, affecting tourists' travelling decisions and the limited availability of reliable data. These conceptual and practical difficulties are frequently overcome by means of simplifying assumptions made to permit the specification of empirical models explaining demand behaviour. Examples of such models are static single equation specifications, which most researchers have adopted, in empirical studies of tourism demand.

Following Johnson and Ashworth's (1990) appeal "*comparisons across studies are urgently needed*" (p. 150), still largely ignored in tourism literature, this chapter analyses the modelling of tourism demand within a static single equation framework, and provides a systematic examination of estimation results obtained from different studies using this approach and a set of observations on the UK demand. The analysis focuses first, on the comparison of estimation results obtained from different single equation models of the UK demand for tourism in France, Spain and Portugal. The differences between models relate to changes in the definition and measurement criteria of the variables included, and to the use of different sets of regressors and linear and log-linear functional forms. Then, we proceed with the comparison of these estimation results with those obtained in other empirical studies involving the UK as an origin country, using the same methodological approach and a similar sample period.

The chapter is structured as follows. Section 3.2. provides a literature review of tourism demand studies which have used the single equation approach. In section 3.3, static single equation models of the UK tourism demand for France, Spain and Portugal are estimated with observations for the period 1969-1997, using different functional forms, different definitions and measurement criteria for some variables, and different sets of regressors. The estimation results of these alternative models are then presented and analysed. In section 3.4, other studies of the UK demand for tourism which use similar empirical methodologies and analogous sample periods are analysed and compared with the results obtained in the previous section. Section 3.5. summarises the main findings and discusses their implications.

3.2. THE SINGLE EQUATION APPROACH IN TOURISM DEMAND MODELLING

3.2.1. THE DIFFICULTIES OF TOURISM DEMAND MODELLING

Studies of tourism demand behaviour are subject to specific problems. The complexity of the motivational structure underlying the decision-making process, and the limited availability of relevant data appear to be two of the main difficulties which have had a significant influence on the way most empirical analysis has been conducted. A quantitative approach in tourism demand analysis requires a formal statistical framework within which the impacts on the demand variable caused by changes in its determinants can be validly and accurately measured. Appropriate econometric modelling provides a good basis for reliable estimation and forecasting which is of considerable value for public policy and private investment decisions. However, appropriate modelling, estimation and forecasting of theoretically consistent and empirically plausible econometric specifications are not easy to achieve in tourism demand contexts.

Rather than constituting a single economic activity, tourism embodies a large set of production and supply activities including transports, communications, catering, entertainment, travel agencies, tour operators, advertising firms, and the production and sale of numerous items. An additional complexity, which distinguishes tourism from other economic sectors, is that tourism products must be consumed in the location of its production.

Tourism demand can be analysed from many different perspectives. From a geographical perspective, tourists' flows can be analysed as domestic or international demand for groups of countries, one sole country, specific regions within a country, specific local areas within a region, specific facilities within local areas. From a motivational point of view influencing the decision to travel and the preference for destinations, the demand for tourism can be viewed as private consumption (personal interest travelling), as a part of the production process (business travelling) or even as a part of the policy-making process affecting this and other sectors of the economy (diplomatic missions). Demand for tourism can also be analysed considering different categories of tourists, grouping them by such criteria as age, gender, nationality or level of income; considering different types of tourism products such as sport, cultural, religious or scientific events, and considering different components of the tourism product such as accommodation, catering or transportation. Irrespective of which characterisation is adopted, tourism analysis remains complex, since it is a demand for heterogeneous goods and services involving, on the supply side, many different industries, production processes and cost structures and, on the demand side, the consideration of a diversity of determinants influencing tourists' behaviour. In addition, and as a result of the unavoidable dislocation process that tourists undertake to consume the goods and services of their choice, tourism demand is much more sensitive to non-economic influences, such as political instability, natural disasters and other special factors, than the majority of other demand behaviour.

These conceptual and practical difficulties, inherent to the research within this context, justify some of the simplifying assumptions that investigators undertake in their attempts to specify econometric models explaining tourism behaviour. However, when such assumptions are inadequate or questionable, they lead to the construction of models which provide unreliable estimates for the responses of demand to changes in its determinants. For instance, assumptions concerning an infinitely elastic supply and the inter-temporal separability of consumption may prove inadequate if prices and demand are simultaneously determined and if decisions to purchase tourism are interrelated over time. Econometric models constructed under these assumptions do not take into account the presence of endogenous regressors and/or that of a dynamic structure, leading to identification, simultaneity and serial correlation bias. Another example of inappropriate specification is to ignore the possibility that the decision to consume specific types of tourism may be related to the consumption of other goods and services, as well as other types of tourism. Other potential problems include those of inappropriate aggregation over the preference structure and/or over the components of tourism products since different individuals and groups may display different tastes or behaviours, and different types of tourism may be consumed in the same destination.

In addition to misspecifications resulting from the lack of a theoretical basis within which reasonable and testable hypotheses can be included, the econometric methodology used in the estimation of tourism demand models can also add problems to the estimates' accuracy. This is generally the case when the estimation procedures do not take into consideration the problem of spurious relationships arising from regressing nonstationary data, the presence of severe collinearity, residual serial correlation problem, inaccuracies associated with small sample estimations and the possible existence of feedback effects and lagged structures.

The difficulties in modelling tourism demand are broadened by the existence of unquantifiable factors influencing demand, as well as the inaccuracy or unavailabily of data on the objectively measurable ones. The scarcity of reliable statistics on tourism, added to the inadequacies of much of the required data mainly compiled from secondary sources or through non-representative sampling surveys, are important limitations that must be taken into consideration when specifying tourism demand equations. Given these difficulties, it is not surprising that econometric studies of tourism demand have not been reported more frequently in the economics literature. Nevertheless, those that have been undertaken constitute a basis for further research, allowing for comparison of results and critical analysis of differences in the estimates obtained and statistical inference provided.

3.2.2. SINGLE EQUATION MODELS OF TOURISM DEMAND

The majority of studies in tourism demand analysis use a concept of demand which includes the entire bundle of goods and services that tourists usually purchase (transport, accommodation, catering, entertainment and related services). Most of these studies use time series data to regress single equation models focusing on the determinants of tourist flows at the national level. Single equation models in studies of tourism demand generally concern different origin and destination countries, different time periods, different sample sizes, different measurement criteria and different variables. They also vary in relation to many aspects of the models' specification. For example, the assumptions underlying the models, the definition of the dependent variable and that of the regressors, the functional form adopted and the econometric methodology used for estimation. Despite the differences, these studies have a common aim, which is to evaluate the sensitivity of tourism demand to changes in its determinants. Consumers, firms and governments depend on the reliability of such information for decision and policy making. However, the precision of the estimates obtained and the validity of the statistical inference and forecasting procedures, depends crucially on the robustness of the theoretical framework within which the models are specified, and on the use of a sound econometric methodology for the estimation of the relationships between the dependent variable and its determinants.

Quantitative formulations of tourism demand can roughly be grouped into two kinds of models: single equation, and system of equations models both in static and dynamic contexts.[1] Reviews of tourism demand studies, focusing on the modelling, estimation methodologies and/or analysis of results are given in Johnson and Ashworth (1990), Sheldon (1990), Tansel (1991) and Lim (1997).

The majority of earlier quantitative studies used a static single equation approach to explain demand behaviour. A typical single equation model of a static tourism demand function can be described as follows:

$$D_{ij} = f(Y_i, P_{ij}, P_{ik}, E_{ij}, E_{ik}, TR_{ij}, TR_{ik}, Z)$$

where D_{ij} is tourism demand by origin i for destination j; Y_i is income of origin i; P_{ij} and P_{ik} are prices in origin i relative to destination j and relative to competitor destinations k; E_{ij} and E_{ik} are exchange rates between origin i and j and between i and competitor destinations k. TR_{ij} and TR_{ik} are transport costs between i and j and between i and competitor destinations k. Z represents a set of qualitative variables affecting the origin's tourism demand.

[1] Recently, studies of statistical simulation of tourism demand regarding habit persistence modelling contribute to enlarging the scope of this classification, see, for example, Darnell and Johnson (2001).

The functional form adopted in the large majority of single equation studies in tourism demand research, is the log-linear form. Although the linear form has also been used, the popularity of the log-linear form is related to both its convenient property of supplying direct estimates of the relevant elasticities and to its claimed (Witt and Witt, 1992; Lee *et al.*, 1996) good empirical performance relative to the linear form. Most studies involve time series analysis on a yearly basis, although there are also examples of the use of pooled and cross sectional data (Jud and Joseph, 1974; Mak *et al.*, 1977; Kliman, 1981; Lin and Sung, 1983; Trembley, 1989; Yavas and Bilgin, 1996; Romily *et al.*, 1998).

The empirical specification of single equation model is defined within the framework of a set of assumptions. Some of the most common underlying this category of models are: independence between tourism consumption and the consumption of other goods and services; independence between decisions to purchase tourism in destination j and to purchase it elsewhere; intertemporal separability of consumption; invariant tastes; perfectly elastic supply; aggregation across consumers and across types of tourism; and predetermined or exogenous explanatory variables. Although some of these assumptions can be deemed reasonable or, at least, unavoidable, others may be considered controversial. Additional assumptions can be found in different contexts, depending on specific features of the origins/destinations involved and on the relevance of particular aspects that researchers want to enphasize in their analysis.

3.2.3. VARIABLES INCLUDED IN SINGLE EQUATION MODELS OF TOURISM DEMAND

The definition and measurement of variables included in single equation models also vary across studies. For example, the dependent variable, representing the demand for tourism, can be measured by total number of visits, arrivals or tourists, as in Kliman (1981), Lin and Sung (1983) and Gunhadi and Boey (1986); visits per head of the origin country population, as in Witt (1980a, 1980b) and Witt and Martin (1985); total real expenditure or receipts, as in Gray (1966), Barry and O'Hagan (1971), Jud and Joseph (1974), Loeb (1982), Lin and Sung (1983), Uysal and Crompton (1984); and per capita real expenditure, as in Artus (1972) and Loeb (1982). However, total or per capita tourism expenditure is generally considered as an appropriate proxy of tourism consumption because, unlike arrivals, visits or tourist numbers, it depicts fluctuations in tourism spending patterns, revealing the factors responsible for changes in expenditure levels. In some circumstances, increasing numbers of tourists can be accompanied by decreasing expenditure due to higher inflation rates, tourists' lower spending propensity or lower average length of stay. Hence, tourism expenditure is likely to be a more accurate measure for tourism demand and, given its direct effects in the destinations' economies, a more interesting variable for policy purposes.

The independent variables are also subject to different definitions and measurement criteria in tourism demand studies. The variables representing the origin's income are usually measured by total or per capita real (disposable) income with few exceptions, notably Lin and Sung (1983) and Gunadhi and Boey (1986), who use the real per capita national income, Jud and Joseph (1974), who uses Gross Domestic Product, and Little (1980) who uses real per capita consumption. The real disposal income of the origin country is viewed as an adequate measure since tourism is considered a final good.[2] Whether total or per capita real income is more appropriate depends on the equation specification and restrictions imposed.

Several considerations related to the price determinants of demand must also be addressed. First, the demand for tourism concerns a bundle of goods and services ranging from air travel to theatre tickets. A tourism-specific price index would be an appropriate measure for this determinant but the difficulties involved in its construction usually preclude its use. Second, demand for tourism in one destination can be sensitive to changes in the relative prices of other destinations and, therefore, these variables should be included. Finally, tourists decide to purchase tourism in one destination by, among other things, comparing their own domestic prices with the foreign prices. As domestic and foreign prices are expressed in different currencies these must be translated into a common currency to provide a clearer idea of the relative purchasing power. Hence, both prices and exchange rates of the origin relative to the destination and to its competitors are generally considered relevant determinants of demand.

Due to the unavailability of tourism-specific price indexes, the majority of investigators accept the consumer price index as a good proxy, although a few, for example Gunadhi and Boey (1986), construct special tourism price indexes for the demand analysis they conduct. Except for Grey (1966), who excludes relative prices from his model on the basis that

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[2] As noted in Song and Witt (2000), if the research focuses on business tourism demand, a more general income variable such as GDP should be used.

"changes in foreign prices, other than those caused by changes in th e exchange rate are unlikely to influence demand ..." (p. 86), most investigators use relative prices as relevant determinants of tourism demand. As shown by Martin and Witt (1987), the relative consumer price index, with or without the consideration of exchange rates, can be a reasonable proxy for the cost of tourism, but exchange rates alone are not acceptable. Hence, prices are present in tourism demand studies either in the form of relative prices between origin and destination with or without competitors' relative prices (Jud and Joseph, 1974; Witt, 1980; Kliman, 1981; Witt and Martin, 1985; Uysal and Crompton, 1984), in the form of relative prices and exchange rates as separate regressors, (Artus, 1972; Little, 1980; Loeb, 1982) or in the form of relative prices adjusted by exchange rates (Kliman, 1981; Gunadhi and Boey, 1986; Martin and Witt, 1987), usually denominated effective (or adjusted) prices. Whether it is more appropriate to consider relative prices and exchange rates as separate regressors or effective prices, has been discussed in the literature by a few authors (O'Hagan and Harrison, 1984; Syriopoulos, 1995; Sinclair and Stabler, 1997). The arguments for and against are generally based on the short-run and long-run effects that changes in these explanatory variables may have on tourism demand. Models specified within a static single equation approach generally intend to describe a long-run equilibrium relationship. In this case, it seems theoretically more appropriate to include effective prices rather than relative prices and exchange rates as separate determinants.

Since transportation costs may represent an important part of the total price of visiting a destination, some studies (e.g. Jud and Joseph, 1974; Kliman, 1981), consider these costs as a separate determinant of tourism demand. The price of tourism consumption can be disaggregated into three essential elements: the price of travelling, the price of commodities and services purchased after arrival, and the price of the destination's currency. The generally adopted definition of tourism demand implies that the price variables included in a demand equation embody the first two of these three elements, making unnecessary the inclusion of a separate price for transportation. However, if the price variables in the right-hand side of demand regressions do not account for transportation cost, this cost should be included as a separate regressor unless, as is the case in various contexts, this cost can be considered irrelevant or not a major influence in the decision-making process. Ruling out a theoretically relevant variable on the grounds that data are difficult to obtain and its inclusion would decrease the degrees of freedom (Uysal and Crompton, 1984), that there are complexities involved in the construction of such variables (Gunadhi and Boey, 1986), or that other studies had found it insignificant (Loeb, 1982), do not seem reasonable criteria for omitting relevant variables. When a transport cost variable is found to be, on a theoretical basis, a separate relevant determinant of tourism demand for a particular set of origins and destinations, it should be included in the model. For instance, if research concerns the analysis of UK demand for destinations far apart such as Ireland and Australia, a transport cost variable is likely to be a relevant determinant of preferences. However, the appropriate form of its inclusion is neither evident nor simple. The reason concerns the difficulties attached to the construction of a comprehensive and accurate measure for the cost of travelling. Generally, an acceptable travel cost variable can be constructed on an average basis, that is, transport costs can be measured as the weighted mean price of all types of transport used to move tourists from an origin to the destinations. Therefore, a meaningful transportation cost variable is difficult to obtain due to the complexity of the fare structure, changes in the route network, seasonal frequency of departures and, generally, the lack of reliable information about tourism traffic. In addition, if transport cost is an important variable in tourists' decision-making process, tourists' choice among alternative destinations, should take account of this determinant. Hence, if the cost of travelling is a relevant determinant of tourism demand for a given destination, the cost of travelling to its competitors ought to be included in the demand equation. However, alongside the problems stated above, it is not always clear which competitors should be considered.

Given these difficulties, most studies which include a separate transport cost variable, measure it as the economy class air fare of a return trip between the origin and destination capital cities. Yet, this is not always advisable since, in many cases, a significant share of the incoming tourists uses other means of transportation. Moreover, a large share of the air traffic is covered by charter flights with complex seasonal cost and transfer structures, which depart from locations other than the origin's capital city and arrive at airports in areas other than that proximate to the destinations' capital city. In these cases, the inaccuracies of the measure used to proxy the travel cost cause the estimates of the respective elasticity to be biased upward or downward, depending on the under- or over-statement of the transport variable. In addition, it has been observed that the transport cost variable, measured as airfares, shows a high negative correlation with the income variable causing collinearity problems. These are often revealed by the insignificance of one or both of the correlated variables and by wrongly-signed coefficients attached to those variables.

However, the omission of a relevant transport cost variable on these grounds is hardly a solution, since collinearity can be overcome by such means as additional data containing more variation in the independent variables. In contrast, estimation bias caused by relevant variables omission cannot be overcome by such means.

Unavailability of data is not an acceptable reason for omitting a relevant transport cost variable, or any other pertinent variable for that matter. However, when that must be the case, researchers should be aware of, and explicitly refer to, the estimation bias inserted in their misspecified models. For instance, it can be shown that the effect of omitting a relevant variable negatively correlated with a variable that is included can cause an upward bias in the estimated coefficient of the included variable.[3] Moreover, the omission of any relevant variable always leads to (false) autocorrelation in the estimated model, with all its known adverse consequences for the validity of the estimation and inference procedures.

Given these considerations, it is not surprising that several tourism demand studies, which have included a transport cost variable found it insignificant (Gray, 1966; Little, 1980). However, there are also examples of a significant negative relationship being found between the transport cost variable and the demand for tourism, although the validity of the estimates may be challenged due to the models' misspecifications. For example, in Kliman (1981), the estimates of the transport cost elasticities of the Canadian demand for tourism ranged between –0.94 (Italy) and –3.09 (Portugal), and in Jud and Joseph (1974) the estimated travel cost elasticity of the world-wide demand for tourism in seventeen Latin American countries was –0.665, while that of the US tourists' demand was –2.022. It should, nevertheless, be noted that, even if these estimates were sanctioned by appropriate modelling and inference procedures, the coefficients of travel cost variables measured as airfares, represent partial elasticities as they can only indicate the effect of a specific air fare change on the tourism demand for a given country or region, *ceteris paribus*. The effect of changes in transport costs other than that specific airfare cannot be inferred from the estimates obtained.

Changes in tastes can have a significant effect on tourism demand. However, due to the difficulties attached to this variable measurement in aggregate terms, most empirical studies use trends to proxy its behaviour. Due to the link between individual preferences structure and consumers' tastes, the explicit modelling of this variable in a tourism context should take place at a disaggregated level of analysis. Nevertheless, the appropriate integration of such a capricious variable in tourism demand models still has a long way to go in empirical research.

Tourism demand can be responsive to advertising campaigns designed to attract visitors to specific destinations. When this is the case, relative marketing expenditures by the destinations should be included as an explanatory variable. For example, Uysal and Crompton (1984) found that this variable had a significant positive effect on the German and Spanish tourism demand for Turkey, ranging between 0.094 (Germany) and 0.596 (Spain). A review of demand models that include marketing variables is provided in Witt and Martin (1987a).

Tourism demand can also be sensitive to special factors not included in any of the variables discussed above: political disturbances, natural disasters, sporting events, religious meetings, international fairs are examples. When these factors are clearly defined they can easily be included in the demand equations and their effects have a straightforward interpretation through their estimated coefficients. Some examples of earlier empirical studies covering several of these special effects within a static single equation approach are Little (1980), Loeb (1982), Uysal and Crompton (1984), Gunadhi and Boey (1986).

3.2.4. STATIC AND DYNAMIC SINGLE EQUATION MODELLING OF TOURISM DEMAND

With few exceptions, a specification with an explicitly dynamic structure of the demand for tourism, distinguishing between short- and long-run effects, has rarely been estimated in the single equation contexts of early research in tourism analysis. Generally, earlier empirical studies are characterised by the absence of an explicit theory of consumers' behaviour, mostly by overlooking its inter-temporal nature; that is, the formation of expectations and habit persistence features. These studies assume that the current level of demand depends only on the current levels of its determinants and ignore the possibility of past demand patterns affecting current consumption. They provide no explanation of the process by which tourism demand occurs over time, and neglect the possibility of an adjustment mechanism propelling the current level of tourism demand to its long-run equilibrium value. Therefore, these studies can only be viewed as specific cases within the wide range of plausible demand behaviours, and not as a general comprehensive analysis of tourists' conduct.

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[3] See Gujarati, 1995, pp. 457-8, for example.

As a result, important aspects of consumer theory have been disregarded and the differences between short- and long-run responses of tourism demand to changes in its determinants have generally been ignored. For example, the theory of inter-temporal choice states that consumers decide how to allocate present and future consumption according to their rate of time preference, meaning that current consumption can depend on any combination of current, past or future values of its determinants. Consumption may be a function of the discounted present value of expected future income, and uncertainty about future income can deter consumption or reduce its current value by inducing precautionary savings. Moreover, consumption decisions are often made in a context of imperfect information, uncertainty and liquidity constraints that restrict current demand. Hence, the level, symmetry and accuracy of information, as well as the degree of imperfection in the relevant markets, are important ingredients in the formation of expectations. Consequently, the explanation of current demand often requires the formulation of a theory of expectations, taking into account if consumers are "forward-looking" or "backward-looking". If consumers are "backward-looking", the demand equation should include lagged values of the income variable; if they are "forward-looking", a suitable expectations process should be taken into account (Sinclair and Stabler, 1997).

Tourism demand specifications modelled as appropriate dynamic forms also permit the examination of important theoretical aspects concerning the sensitivity of demand to variations in prices and exchange rates. For instance, it is generally assumed that, in the short-run, tourists are more aware of exchange rates than relative prices. However, favourable variations in exchange rates can be offset by higher inflation rates and these compensatory movements of prices take time before they are fully acknowledged by potential tourists. Hence, in the short-run, the investigation of exchange rate effects separately from price effects assumes particular importance while in the long-run, the effective price variable (relative prices adjusted for exchange rates) are more likely to be relevant for tourism demand than the separate variables of prices and exchange rates. Further issues concerning the dynamic specification of tourism demand responses to changes in price variables are related to the fact that tourism purchases can be made in advance, as well as simultaneously, with their actual consumption. This implies that lagged price-change variables, in addition to their current values, may be appropriate in certain cases. On the other hand, expected future changes in relative prices and/or exchange rates are less likely to be relevant determinants of tourism demand due to the lack of information (unpredictability) generally attached to fluctuations in these variables' values.

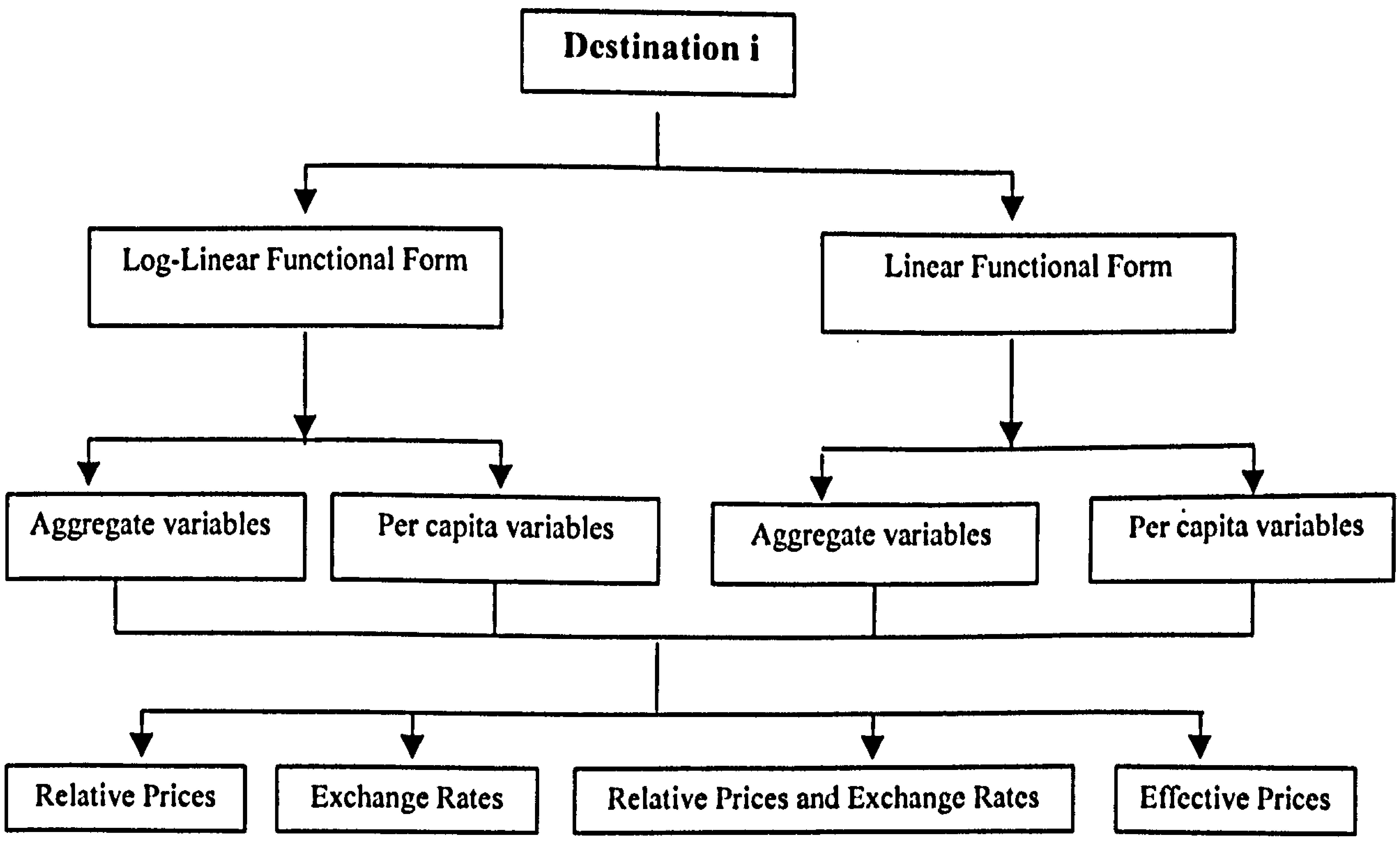
The few early studies that introduce dynamic elements in their single equation specifications, mostly do so by including a lag structure in the variables, without a clear explanation of the theoretical principles within which that structure can be justified (e.g. Witt, 1980 and Martin and Witt, 1988). More recent studies do consider a theory-based dynamic approach in the formulation of their models and explicitly show how the empirical models are derived from the principles of economic theory (e.g. Syriopoulos, 1995, Vogt and Wittayakorn, 1998, Kim and Song, 1998 and Song *et al.*, 2000).

The above discussion suggests that most empirical research using single equation quantitative approaches to study tourism demand fails to consider several important theoretical and methodological matters related to the modelling, estimation and inference procedures. Empirical inadequacies such as omission of relevant variables, incorrect functional form, or ignoring the problems attached to regressing nonstationary data, are bound to result in mis-specified models which produce inconsistent estimates, invalid inference and unreliable forecasts. These models' limitations are summarised in Syriopoulos' (1995, pp. 318-9) remark that "*the most critical weakness of these studies has been the general failure to pay attention to the 'dynamics' of tourism demand, as these works rest upon a static (and, therefore, incomplete) theoretical and empirical framework.*"

An appropriate approach within the single equation framework would have to consider a flexible dynamic econometric model, involving all possible relevant determinants as well as an equilibrium correction mechanism. Such a model would allow for the separate examination of short- and long-run impacts of the independent variables' changes on tourism demand levels. The construction of a dynamic model could be implemented according to the "general to specific" methodology discussed in, for example, Hendry and Mizon (1978), Hendry and Richard (1982) and Hendry (1987, 1995). Unlike the traditional "specific to general" approach applied in most tourism demand studies, where the manipulations carried out on the 'simple' model in order to attain a more general one are both theoretically unfounded and econometrically inappropriate, the general to specific approach consists of modelling a theory-based demand equation in the most general terms possible, and attains a final parsimonious specification which is legitimately derived from the general form by testing plausible restrictions imposed on the variables' coefficients.

3.3. STATIC SINGLE EQUATION MODELS OF THE UK TOURISM DEMAND IN FRANCE, SPAIN AND PORTUGAL

We demonstrate the limitations of the static single equation approach in modelling tourism demand, by estimating the UK demand for tourism in France, Spain and Portugal using alternative specifications, and comparing the results obtained. The specifications are defined by considering different functional forms, different aggregation levels for the expenditure and income variables and different sets of regressors. The specifications used are summarised in the following diagram:



Hence, for each destination, sixteen different models are estimated. The analysis of the statistical performance, data-coherence, and theoretical consistency of these models allows for a systematic evaluation of the methodological tools applied in their derivation, and provides a legitimate basis for comparison of the estimates obtained. In the next section, we start by discussing the assumptions underlying the modelling framework of the UK demand for tourism in France, Spain and Portugal over the period 1969-1997. A full description of the variables included is given. The static single equations resulting from different specifications are estimated with the OLS method and an interpretation of the estimation results is provided. Finally, we compare these results with those of other similar studies and analyse the implications of this comparison exercise.

3.3.1. Static single equations of the UK tourism demand for France, Spain and Portugal

3.3.1.1. Assumptions

The models are based on a number of assumptions. Separability between tourism consumption and labour supply is assumed, since it is postulated that the decisions to engage in paid work are made prior to the decision to engage in tourism consumption. Hence, the former are not related to those underlying the spending of the earned income on tourism products. Separability between UK consumers' expenditure in tourism and their expenditure on other goods and services is also assumed. This implies the separation of consumers expenditure into groups of commodities where each group of preferences is independent of the preferences underlying other groups. Therefore, it is assumed that the expenditure allocated to France, Spain and Portugal by UK tourists is independent from the expenditure on other goods and services, including tourism budgets assigned to other destinations.

UK tourism in France, Spain and Portugal can be viewed as a clearly defined product, with relatively constant motivations underlying the tourists' decision to visit these countries, over the period under study. Indeed, not only is the UK demand for these destinations predominantly a demand for holidays by private households (business travellers comprise only a small proportion of total flows), but also this group of destinations has specific common attributes which separate them from others, either in Europe or in the world.[4]

[4] The most important are climatic factors, relatively small cultural differences and their proximity to the origin country rendering transportation costs relatively unimportant in this context.

We also assume separate demand equations for UK tourism expenditure in France, Spain and Portugal.[5]

In the long-run, France, Spain and Portugal's supply of goods and services to UK tourists may reasonably be assumed to be perfectly elastic, since the level of UK tourists' consumption at any period, can be viewed as relatively small compared with that of national residents and other tourists in most areas. In fact, investment in the tourism sector is generally undertaken to satisfy both current and future demand, and tourism-related supplying activities are particularly vigorous in destinations such as the ones under consideration, where tourism revenue is an important part of both regional and national economies. Furthermore, a shift from accommodation in hotels to self-catering facilities has been observed in these countries, contributing to overcome any potential accommodation constraints. Therefore, "*the spectre of full capacity can reasonably be ignored*" (Gray, 1966, p.86). As a result, prices may be assumed exogenous or, at least, predetermined explanatory variables.[6]

Since the static equations of the UK tourism demand for France, Spain and Portugal specified in the next section included nonstationary time series, the regressions may be spurious unless the variables are cointegrated, that is, unless a genuine long-run equilibrium relationship exists among the time series involved. Consequently, for the time being, we assume that such a relationship exists, implying that the static equations specified represent the steady state equilibrium relationship between the UK demand for tourism and its determinants.[7]

3.3.1.2. Functional form and variables definition

The modelling of a structural relationship between the UK demand for tourism and its determinants requires the discussion of further issues such as the functional form adopted and

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[5] Although these destinations can be viewed as a separate group from other destinations, the assumption of separate equations of UK tourism demand for France, Spain and Portugal can be regarded as dubious, given their previously assumed common characteristics which make these destinations interdependent in the first place. This assumption is inherent in the single equation framework and, therefore, inevitable in this context. However, it will be relaxed in chapters 5, 6 and 7 where systems of equations are considered.

[6] Tourism prices are frequently determined in advance of the actual consumption (e.g. package holidays and pre-booking conditions) and do not ordinarily respond to the level of demand in the short-run. Nevertheless, the assumption of exogenous prices is relaxed in chapter 7.

[7] This assumption is given empirical support in chapter 4, where the cointegration analysis of the relevant variables within a dynamic single equation framework is carried out.

the definition of the variables included in the single equations to be estimated. Since theory does not provide a clear indication about which functional form should be used, there is no *a priori* certainty about whether the functional form should be linear or log-linear. Although in the linear form, the coefficients are interpreted as the absolute change in the dependent variable per unit absolute change in its determinants, while in the log-linear form the coefficients represent the relative change of the dependent variable per unit relative change in its determinants, both forms give important information concerning the impacts that changes in the regressors have on tourism demand levels. However, no direct comparison can be made between these forms using the usual econometric criteria since in the first case, the dependent variable is in the linear form, and in the second in the log form. Moreover, since the variables' coefficients represent different things (absolute changes in the linear case, and relative changes or elasticities in the log-linear case), no direct comparison can be made between them as well. Nevertheless, an indirect comparison can be made if the relevant elasticities are computed from a linear form model regressing, say, Y on X, using the formula $\varepsilon = (\partial Y/\partial X).(X^*/Y^*)$, where ε represents the elasticity of Y with respect to X, and Y^* and X^* are the average values of the dependent variable Y and the independent variable X.[8]

Although some empirical studies have shown that a number of tourism demand relationships can be reasonably approximated by a linear functional form (Smeral *et al.*, 1992), the large majority of studies applying a static single equation approach to tourism demand analysis use a log-linear functional form for their models. This choice is mainly justified on the grounds of the same procedure being used previously, or inferior results being obtained with the linear form. However, in the particular modelling exercise we are involved with here, no clear-cut decision could be made concerning the appropriate functional form. Indeed, the analysis of the data, functional form tests and goodness of fit measures, indicate both as equivalent. Hence, the estimation results of both the log-linear and linear specifications are reported, to permit a thorough comparison of statistical performances.

The dependent variable expressing the UK demand for tourism in France, Spain and Portugal is the total expenditure UK tourists allocate to each destination. Data for the UK total

[8] The elasticities are computed for the average values of Y and X. In this way, estimates of the *average* elasticities are obtained. Hence, in the linear form the elasticities are not constant as their magnitudes depend on the chosen values of Y and X, while the log-linear form provides constant elasticities.

tourism expenditure, disaggregated by destinations and measured in million pounds sterling, were obtained from one common source, *The Business Monitor MA6 (1970-1993)* continued as *Travel Trends (1995-1998)*, so reducing potential data inconsistencies. The values of the dependent variable in real terms are computed by deflating the UK tourism budget allocated to each destination by its consumer price index.

The variables considered to be relevant determinants of the UK demand for tourism are the origin's real disposal income (in both aggregate and per capita terms) and the price of tourism which is represented by its two separable components: the origin's consumer price index, both relative to that of the destination and to that of its competitors, and the exchange rate between the origin's currency and that of both the destination and its competitors. Data for the UK population, price indexes and exchange rates were obtained from the *International Financial Statistics* (IMF, Yearbook 1980-98).

Aggregate and per capita real expenditure and income were chosen to represent, respectively, the dependent variable and one of its determinants in alternative specifications. However, since the UK population did not vary much over the sample period (from 55 million in 1968 to 58 million in 1997), any measurement errors are not likely to be serious and the consideration of aggregate or per capita expenditure and income are not expected to have a substantial effect on the estimated results.

Given that differences in the distance between the UK and any of the destinations under consideration are relatively small, it can be argued that, over time, little variability is to be expected in the relative costs of travelling. Hence, a transport cost variable is not expected to add much explanatory power to the model.

Other special events, which appear to be relevant to the explanation of the UK demand for tourism in France, Spain and Portugal, are taken into account by dummy variables included in the demand equations. Such events are the political upheaval and economic instability in Portugal following the revolution of April 1974, and the political changes that occurred in Spain following the death of Franco in 1976. It is assumed that these events affected the three destinations in different periods. For instance, the Portuguese revolution influenced UK tourism demand in the region more intensively and for a longer period in the case of Portugal (1974-1979) than in the cases of its neighbouring countries. The UK demand for tourism in Spain was affected by both the Portuguese revolution and Spain's own process of political change but their effects were less durable in this destination. Hence the dummy variable included in the equation for Spain is defined for the period 1974-1977. This same period is also considered in the equation for France, assuming that the political change in Spain rather than that in Portugal, had greater effects on the UK demand for France. However its pertinence in this destination equation is expected to be secondary.[9]

In the static version of the single equation model, no intertemporal dependence of tourism consumption and no explicit interdependencies among competing destinations are allowed for. Hence, the equations are estimated separately and it is assumed that only the contemporary values of the determinants have relevant effects on the current demand for tourism. The omission of such features from the modelling of tourism demand is likely to cause misspecification bias in the models' estimation results. Hence, evidence of serial correlation is to be expected in the estimation of these static single equation models.

Given these considerations, the static form of a log-linear single equation model of the UK demand for tourism in destination i is:[10]

$$LREX_i = \alpha_i + \beta_i LRDI + \sum_j \gamma_{ij} LRP_j + \sum_j \delta_{ij} LER_j + \eta_i D_i + u_i \tag{3.1}$$

where i = F (France), S (Spain) and P (Portugal); $LREX_i$ is the logarithm of UK real tourism expenditure in destination i; LRDI is the logarithm of UK real disposable income; LRP_i and LER_i are, respectively, the logarithms of tourism relative price in destination i and the exchange rate between the UK and destination i currencies. These variables were constructed as follows:

$$LREX_i = \ln(REX_i) = \ln\left(\frac{EX_i}{CPI_i}\right)$$

[9] Other events, which are believed to have had a relevant impact in the UK tourism demand for these destinations, are the oil crises of 1973 and 1979, Portugal and Spain joining the EU in 1986 and the opening of the Channel Tunnel in 1994. However, the dummy variables corresponding to these events were not statistically significant when included in a static single equation context. Therefore, these variables were omitted from these models. Yet, as will be shown in following chapters, the significance of these events may be captured by dummy variables defined in different ways within a dynamic single equation and system of equations' specifications.

[10] In tables 3.1 to 3.12 showing the estimation results below, the superscript "a" and "c" over the relevant variables denote the representation of the aggregate and per capita forms of the expenditure and income variables respectively,. The log-linear form is indicated by adding the letter "L" added to the names of all variables in equation (3.1), and the linear form of the model is denoted by the omission of this letter.

where REX_i is the UK real expenditure and EX_i is the UK nominal expenditure allocated to destination i, and CPI_i is the consumer price index of destination i;

$$LRDI = \ln(RDI)$$

where RDI is the UK real disposable income;

$$LRP_i = \ln(RP_i) = \ln\left(\frac{CPI_i}{CPI_{UK}}\right)$$

where RP_i is the relative price in destination i and CPI_{UK} is the UK consumer price index;

$$LER_i = \ln(ER_i) = \ln\left(\frac{CUR_i}{CUR_{UK}}\right)$$

where ER_i is the exchange rate between the UK and destination i currencies, CUR_i is the national currency of destination i and CUR_{UK} is the UK currency.

D_i is a dummy variable which takes the value of unity for observations in the period 1974-1979 in the equation for Portugal, and 1974-1977 in the equations for France and Spain, and zero otherwise; u_i is a well-behaved stochastic disturbance term in the i^{th} equation.

Hence, the single equation models of the UK demand for tourism in France, Spain and Portugal are:

$$LREX_F = \alpha_F + \beta_F LRDI + \gamma_{FP} LRP_P + \gamma_{FS} LRP_S + \gamma_{FF} LRP_F + \delta_{FP} LER_P + \delta_{FS} LER_S + \delta_{FF} LER_F + \eta_F D_F + u_F$$
$$LREX_S = \alpha_S + \beta_S LRDI + \gamma_{SP} LRP_P + \gamma_{SS} LRP_S + \gamma_{SF} LRP_F + \delta_{SP} LER_P + \delta_{SS} LER_S + \delta_{SF} LER_F + \eta_S D_S + u_S$$
$$LREX_P = \alpha_P + \beta_P LRDI + \gamma_{PP} LRP_P + \gamma_{PS} LRP_S + \gamma_{PF} LRP_F + \delta_{PP} LER_P + \delta_{PS} LER_S + \delta_{PF} LER_F + \eta_P D_P + u_P$$

where the disturbances are assumed to be white noise stochastic processes.

The estimation of these equations is performed using OLS and four different sets of explanatory variables, giving rise to four different regression equations: in regression 1, only the relative prices are considered; in regression 2, the exchange rates alone are included; in regression 3, both relative prices and exchange rates are present; in regression 4, only effective prices are included.[11]

[11] The effective price of tourism in destination i is defined, in the log form, as: $LP_i = \ln(P_i) = \ln(RP_i/ER_i)$, where LP is the logarithm of the effective price (P) and RP and ER are defined as above. The linear form of the model omits the letter "L" in this variable's name.

3.3.2. ANALYSIS OF THE ESTIMATED RESULTS

The estimated results are presented in tables 3.1 to 3.12 labelled Portugal, Spain and France. There are four tables for each destination. Each of the four tables pertaining to one destination identifies a different specification of the models: the first table, specifies a log-linear functional form with aggregate values for the expenditure and income variables; the second, specifies the same functional form with per capita values for the expenditure and the income variables; the third, specifies a linear functional form with aggregate expenditure and income variables; the fourth specifies a linear functional form with per capita expenditure and income.

In some specifications, one or more variables are deleted due to their insignificance. The resulting estimation results are labelled with capital letters to distinguish them from those obtained with all the variables included. For example, in the equation for Portugal, the estimation of regression 4 is carried out in two different ways: regression 4(A) includes the effective prices of Portugal, Spain and France and regression 4(B) excludes the effective price of France. In the equation for France, the estimation of regressions 3 and 4 are performed in three different ways. Regression 3(A) includes all relevant variables in each particular model; regression 3(B) omits the dummy and the relative price and exchange rate of Portugal; regression 3(C) omits the exchange rate of Spain. Regression 4(A) includes all relevant variables in each particular model; regression 4(B) excludes the dummy variable and the effective price of Portugal; regression 4(C) omits the effective price of Spain.

The tables display the estimates (t values in brackets) of the intercept (INT) and coefficients of the variables included in each model. The last four rows of each table show the adjusted R^2 statistic for each model's goodness of fit, the F statistic for the overall significance test, and the Durbin-Watson (DW) and Lagrange Multiplier (LM) statistics for the test of residual serial correlation. The LM statistic follows a chi-square distribution and for this statistic the correspondent p values are shown in brackets. The symbols •, °and ' represent, respectively, the 1%, 5% and 10% significance levels.

3.3.2.1. Presentation of the estimated results

PORTUGAL

Table 3.1: Log-linear functional form and aggregate expenditure and income

	1	2	3	4	
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER	Effective Prices	
				A	B
INT	-9.415 (-1.29)	-0.894 (-0.11)	1.610 (0.18)	-1.924 (-0.27)	-1.176 (-0.21)
LRDI[a]	1.318 (2.07)°	0.568 (0.88)	1.222 (2.08)°	0.579 (1.61)	0.538 (1.92)'
LRP_P	-1.420 (-3.29)•		-1.696 (-4.19)•		
LRP_S	3.541 (3.32)•		8.051 (4.10)•		
LRP_F	-0.213 (-0.45)		-1.981 (-2.31)°		
LER_P		0.200 (0.76)	-0.758 (-1.226)		
LER_S		-0.345 (-0.53)	-2.132 (-3.68)•		
LER_F		0.417 (0.75)	2.399 (2.92)•		
LP_P				-0.999 (-2.29)°	-0.942 (-3.06)•
LP_S				0.774 (1.22)	0.824 (1.46)
LP_F				0.133 (0.19)	
D_P	-0.323 (-3.05)•	-0.254 (-1.48)	-0.457 (-3.66)•	-0.296 (-2.16)°	-0.297 (-2.22)°
$\overline{R}^2$	0.694	0.557	0.804	0.665	0.679
F	13.67	8.05	15.39	12.12	15.78
DW	0.863	0.846	1.517	1.013	1.005
LM	4.2 (0.04)	5.6 (0.02)	0.4 (0.52)	3.6 (0.06)	3.6 (0.06)

Table 3.2: Log-linear functional form and per capita expenditure and income

	1	2	3	4	
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER	Effective Prices	
				A	B
INT	-8.759 (-1.71)	-3.122 (-0.53)	2.457 (0.35)	-3.559 (-0.62)	-2.886 (-0.61)
$LRDI^c$	1.403 (2.03)°	0.628 (0.87)	1.234 (1.98)'	0.554 (1.43)	0.504 (1.66)
LRP_P	-1.430 (-3.41)•		-1.684 (-4.33)•		
LRP_S	3.519 (3.34)•		8.047 (4.09)•		
LRP_F	-0.196 (-0.42)		-1.989 (-2.34)°		
LER_P		0.177 (0.65)	-0.768 (-1.26)		
LER_S		-0.339 (-0.52)	-2.129 (-3.67)•		
LER_F		0.474 (0.86)	2.401 (2.92)•		
LP_P				-1.021 (-2.37)°	-0.958 (-3.10)•
LP_S				0.767 (1.20)	0.824 (1.45)
LP_F				0.150 (0.22)	
D_P	-0.326 (-3.08)•	-0.246 (-1.43)	-0.460 (-3.72)•	-0.291 (-2.15)°	-0.292 (-2.20)°
$\overline{R}^2$	0.672	0.521	0.790	0.639	0.654
F	12.46	7.08	14.14	10.93	14.21
DW	0.863	0.838	1.517	1.013	1.005
LM	4.23 (0.04)	5.82 (0.02)	0.40 (0.53)	3.6 (0.06)	3.6 (0.06)

Table 3.3: Linear functional form and aggregate expenditure and income

	1	2	3		4	
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER		Effective Prices	
			A	B	A	B
INT	-287.3 (-1.07)	359.7 (1.49)	-482.0 (-1.66)	-306.2 (-1.95)'	149.5 (1.18)	174.0 (2.02)'
RDI^a	0.004 (1.88)'	-0.0005 (-0.27)	0.005 (2.64)°	0.005 (2.86)•	0.003 (1.35)	0.001 (1.49)
RP_P	-421.7 (-3.06)•		-601.2 (-4.05)•	-625.1 (-4.67)•		
RP_S	670.9 (3.24)•		1234.7 (3.48)•	1011.2 (3.70)•		
RP_F	-34.1 (-0.37)		-85.6 (-0.70)			
ER_P		0.406 (0.94)	0.05 (0.08)	0.49 (1.01)		
ER_S		0.060 (0.08)	-2.24 (-2.87)•	-1.74 (-2.89)•		
ER_F		-8.704 (-0.65)	17.91 (1.04)			
P_P					-60687.0 (-2.24)°	-55459.4 (-2.97)•
P_S					50231.2 (1.68)	53062.3 (1.93)'
P_F					427.7 (0.27)	
D_P	-87.4 (-3.34)•	-88.4 (-2.08)°	-121.6 (-3.79)•	-134.0 (-4.73)•	-84.1 (-2.47)°	-83.9 (-2.51)°
$\bar{R}^2$	0.690	0.525	0.755	0.765	0.643	0.657
F	13.49	7.20	11.76	16.17	11.09	14.41
DW	1.008	0.934	1.447	1.406	1.049	1.044
LM	4.29(0.04)	4.81(0.03)	1.39(0.24)	1.16(0.28)	3.83(0.05)	3.98(0.05)

Table 3.4: Linear functional form and per capita expenditure and income

	1	2	3		4	
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER		Effective Prices	
			A	B	A	B
INT	-5.62 (-1.18)	5.86 (1.27)	-8.85 (-1.70)	-5.69 (-1.99)'	2.88 (1.26)	3.42 (2.29)°
RDI^c	0.005 (1.95)'	-0.0004 (-0.17)	0.006 (2.60)°	0.005 (2.79)•	0.001 (1.15)	0.001 (1.20)
RP_P	-7.70 (-3.33)•		-10.44 (-4.22)•	-10.92 (-4.87)•		
RP_S	11.71 (3.26)•		21.65 (3.47)•	17.69 (3.66)•		
RP_F	-0.49 (-0.30)		-1.54 (-0.71)			
ER_P		0.006 (0.75)	-0.0002 (-0.01)	0.008 (0.91)		
ER_S		0.0013 (0.10)	-0.04 (-2.84)•	-0.030 (-2.83)•		
ER_F		-0.114 (-0.47)	0.32 (1.05)			
P_P					-1110.1 (-2.34)°	-1004.5 (-3.02)•
P_S					861.9 (1.60)	921.8 (1.86)'
P_F					8.90 (0.32)	
D_P	-1.58 (-3.41)•	-1.49 (-1.95)'	-2.18 (-3.83)•	-2.40 (-4.76)•	-1.45 (-2.42)°	-1.44 (-2.45)°
$\bar{R}^2$	0.670	0.470	0.736	0.747	0.611	0.626
F	12.38	5.96	10.76	14.76	9.80	12.70
DW	1.005	0.895	1.435	1.394	1.039	1.033
LM	4.26 (0.04)	5.36 (0.02)	1.46 (0.23)	1.20 (0.27)	3.9 (0.05)	4.1 (0.04)

SPAIN

Table 3.5: Log-linear functional form and aggregate expenditure and income

	1 Relative Prices (RP)	2 Exchange Rates (ER)	3 RP and ER	4 Effective Prices
INT	-10.733 (-1.96)'	-7.602 (-1.41)	-12.434 (-1.41)	-19.990 (-3.47)•
LRDI[a]	1.581 (3.31)•	1.123 (2.37)°	1.640 (2.72)°	1.944 (6.57)•
LRP_P	-0.805 (-2.25)°		-0.872 (1.84)'	
LRP_S	2.865 (2.99)•		2.307 (1.27)	
LRP_F	0.634 (1.91)'		0.941 (1.12)	
LER_P		0.025 (0.12)	0.225 (0.36)	
LER_S		0.429 (0.92)	0.103 (0.20)	
LER_F		-0.085 (-0.22)	-0.328 (-0.40)	
LP_P				-0.870 (-2.02)'
LP_S				-0.624 (-1.37)
LP_F				1.280 (2.09)°
D_P	0.248 (2.55)°	0.134 (1.14)	0.257 (2.08)°	0.195 (1.54)
$\overline{R}^2$	0.859	0.798	0.839	0.819
F	34.98	23.15	19.19	26.35
DW	1.743	1.244	1.69	1.238
LM	0.06 (0.80)	3.73 (0.05)	0.15 (0.70)	3.45 (0.06)

Table 3.6: Log-linear functional form and per capita expenditure and income

	1	2	3	4
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER	Effective Prices
INT	-8.618 (-2.24)°	-7484 (-1.88)	-9.902 (-1.45)	-16.648 (-3.49)•
LRDI[c]	1.614 (3.12)•	1.176 (2.24)°	1.657 (2.58)°	2.022 (6.31)•
LRP_P	-0778 (-2.23)°		-0.823 (1.82)'	
LRP_S	2.799 (2.94)•		2.270 (1.24)	
LRP_F	0.633 (1.90)'		0.915 (1.09)	
LER_P		0.012 (0.06)	0.191 (0.31)	
LER_S		0.432 (0.93)	0.121 (0.24)	
LER_F		-0.068 (-0.18)	-0.319 (-0.39)	
LP_P				-0.822 (-1.95)'
LP_S				-0.644 (-1.39)
LP_F				1.252 (2.07)°
D_P	0.246 (2.53)°	0.134 (1.14)	0.252 (2.05)'	0.190 (1.51)
$\bar{R}^2$	0.843	0.778	0.821	0.800
F	31.17	20.60	17.08	23.45
DW	1.740	1.248	1.68	1.234
LM	0.07 (0.79)	3.69 (0.06)	0.17 (0.68)	3.49 (0.06)

Table 3.7: Linear functional form and aggregate expenditure and income

	1	2	3	4
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER	Effective Prices
INT	-4162.3 (-3.67)•	-0.644 (-0.00)	-3739.3 (-2.54)°	-1081.2 (-2.04)°
RDI[a]	0.027 (3.07)•	0.012 (1.49)	0.023 (2.51)°	0.033 (6.82)•
RP_P	-1409.3 (-2.43)°		-1305.9 (1.96)'	
RP_S	3817.7 (4.16)•		3529.1 (1.98)'	
RP_F	897.7 (2.52)°		1316.3 (2.31)°	
ER_P		1.102 (0.64)	0.110 (0.04)	
ER_S		4.419 (1.43)	0.811 (0.52)	
ER_F		-47.241 (-1.00)	-75.7 (-0.94)	
P_P				-337383.5 (-2.18)°
P_S				-177229.5 (-1.47)
P_F				18810.8 (2.47)°
D_P	297.4 (2.54)°	125.1 (0.82)	226.9 (1.66)	237.2 (1.30)
$\bar{R}^2$	0.871	0.800	0.864	0.810
F	38.73	23.43	23.28	24.9
DW	1.768	1.176	1.603	1.080
LM	0.19 (0.67)	5.46 (0.02)	0.65 (0.42)	5.64 (0.02)

Table 3.8: Linear functional form and per capita expenditure and income

	1	2	3	4
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER	Effective Prices
INT	-72.4 (-3.58)•	-0.481 (-0.03)	-65.6 (-2.48)°	-20.81 (-2.17)°
RDI[a]	0.027 (2.84)•	0.011 (1.23)	0.023 (2.35)°	0.035 (6.54)•
RP_P	-23.6 (-2.44)°		-22.2 (2.03)'	
RP_S	65.7 (4.14)•		61.3 (1.97)'	
RP_F	15.5 (2.46)°		22.8 (2.27)°	
ER_P		0.020 (0.63)	0.001 (0.02)	
ER_S		0.076 (1.40)	0.032 (0.52)	
ER_F		-0.804 (-0.95)	-1.3 (-0.92)	
P_P				-5588.2 (-2.11)°
P_S				-3246.9 (-1.51)
P_F				322.5 (2.43)°
D_P	5.2 (2.52)°	2.2 (0.83)	4.0 (1.66)	4.1 (1.28)
$\bar{R}^2$	0.856	0.778	0.848	0.789
F	34.26	20.58	20.56	22.00
DW	1.771	1.182	1.607	1.083
LM	0.18 (0.67)	5.36 (0.02)	0.62 (0.43)	5.59 (0.02)

FRANCE

Table 3.9: Log-linear functional form and aggregate expenditure and income

	1	2	3		4	
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER		Effective Prices	
			A	B	A	B
INT	-3.983 (-0.65)	-10.340 (-1.50)	-10.319 (-1.16)	-9.021 (-1.58)	-27.762 (-3.75)•	-20.885 (-4.40)•
LRDI[a]	0.979 (1.83)'	1.480 (2.46)°	1.465 (2.40)°	1.506 (3.48)•	3.125 (8.22)•	2.803 (10.39)•
LRP_P	0.475 (1.19)		0.089 (0.19)			
LRP_S	0.079 (0.07)		0.259 (0.14)	1.495 (2.46)°		
LRP_F	-2.349 (-6.33)•		-2.564 (-3.01)•	-2.951 (-6.90)•		
LER_P		0.796 (3.18)	0.320 (0.50)			
LER_S		-0.335 (-0.57)	-0.452 (-0.88)	-0.637 (-1.45)		
LER_F		-0.879 (-1.81)'	0.622 (0.75)	1.033 (2.41)°		
LP_P					-0.661 (-1.20)	
LP_S					1.458 (2.49)°	1.660 (3.12)•
LP_F					-1.430 (-1.82)'	-2.100 (-4.03)•
D_P	-0.062 (-0.58)	0.080 (0.54)	0.034 (0.27)		0.161 (0.99)	
$\overline{R}^2$	0.971	0.947	0.973	0.976	0.951	0.952
F	189.04	100.37	127.12	226.25	110.25	186.51
DW	0.752	0.698	1.236	1.297	1.342	1.220
LM	9.41(0.00)	9.40(0.00)	3.37(0.07)	2.01(0.16)	2.43(0.12)	3.51(0.06)

Table 3.10: Log-linear functional form and per capita expenditure and income

	1	2	3		4	
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER		Effective Prices	
			A	B	A	B
INT	-4.082 (-0.95)	-8.637 (-1.70)	-8.066 (-1.16)	-7.057 (-1.58)	-19.808 (-3.19)•	-14.954 (-3.77)•
LRDI[a]	0.981 (1.70)	1.509 (2.25)°	1.434 (2.21)°	1.519 (3.13)•	3.272 (7.81)•	2.978 (9.94)•
LRP_P	0.473 (1.21)		0.143 (0.31)			
LRP_S	0.083 (0.08)		0.260 (0.14)	1.516 (2.42)°		
LRP_F	-2.348 (-6.30)•		-2.608 (-3.07)•	-2.954 (-6.83)•		
LER_P		0.797 (3.11)•	0.280 (0.44)			
LER_S		-0.325 (-0.55)	-0.444 (-0.85)	-0.632 (-1.43)		
LER_F		-0.891 (-1.84)'	0.640 (0.77)	1.019 (2.37)°		
LP_P					-0.530 (-0.96)	
LP_S					1.447 (2.40)°	1.629 (3.01)•
LP_F					-1.536 (-1.94)'	-2.065 (-3.90)•
D_P	-0.063 (-0.58)	0.081 (0.54)	0.027 (0.22)		0.143 (0.87)	
$\bar{R}^2$	0.970	0.944	0.972	0.974	0.947	0.949
F	180.97	95.74	120.90	214.55	101.74	175.67
DW	0.752	0.703	1.229	1.293	1.346	1.235
LM	9.43(0.00)	9.32(0.00)	3.48(0.06)	2.07(0.15)	2.43(0.12)	3.37(0.07)

Table 3.11: Linear functional form and aggregate expenditure and income

	1	2	3			4		
	Relative Prices (RP)	Exchang e Rates (ER)	RP and ER			Effective Prices		
			A	B	C	A	B	C
INT	1771.7 (3.49)•	-1513.4 (-2.87)•	1569.2 (2.55)°	-1280.9 (-2.94)•	1305.9 (2.37)°	-1188.9 (-4.23)•	-1506.8 (-5.17)•	-1190.1 (-4.20)•
RDI[a]	0.013 (3.37)•	0.032 (6.42)•	0.013 (3.40)•	0.034 (7.74)•	0.014 (3.71)•	0.026 (10.02)•	0.032 (15.00)•	0.028 (14.00)•
RP_P	1716.0 (6.61)•		1946.4 (7.00)•		1883.1 (6.99)•			
RP_S	-2464.2 (-5.99)•		-2411.3 (3.24)•	-115.7 (-0.19)	-1943.1 (-3.45)•			
RP_F	-791.7 (-4.97)•		-1030.7 (-4.33)•	-593.4 (-1.76)'	-1117.0 (-5.08)•			
ER_P		0.385 (0.35)	-1.597 (-1.23)		-1.94 (-1.56)			
ER_S		-1.034 (-0.52)	1.408 (0.96)	-1.381 (-0.60)				
ER_F		-3.227 (-0.11)	25.452 (0.76)	47.562 (1.16)	48.32 (2.05)°			
P_P						277328.2 (3.39)•		262890.0 (3.21)•
P_S						79175.4 (1.24)	46797.6 (0.66)	
P_F						-13625.9 (-3.79)•	-3403.6 (-1.12)	-11183.2 (-3.14)•
D_P	-191.2 (-3.64)•	-62.2 (-0.64)	143.1 (2.51)°		-138.70 (-2.44)°	-211.7 (-2.19)°		-174.1 (-1.98)'
$\bar{R}^2$	0.985	0.953	0.986	0.958	0.987	0.970	0.958	0.969
F	374.63	115.51	256.40	129.24	294.01	180.22	214.82	219.94
DW	1.467	0.869	1.705	0.931	1.792	1.492	0.904	1.226
LM	2.1(0.15)	9.7(0.00)	0.6(0.42)	7.1(0.01)	0.3(0.60)	1.6(0.21)	8.3(0.04)	4.0(0.05)

Table 3.12: Linear functional form and per capita expenditure and income

	1	2	3			4		
	Relative Prices (RP)	Exchange Rates (ER)	RP and ER			Effective Prices		
			A	B	C	A	B	C
INT	31.9 (3.55)•	-27.5 (-2.60)•	27.7 (2.50)°	-24.7 (-2.99)•	23.6 (2.36)°	-22.2 (-4.30)•	-28.8 (-5.27)•	-22.5 (-4.33)•
RDI^c	0.012 (2.87)•	0.033 (5.49)•	0.013 (3.06)•	0.035 (6.70)•	0.013 (3.29)•	0.027 (9.42)•	0.033 (13.69)•	0.029 (14.09)•
RP_P	30.3 (7.02)•		33.6 (7.34)•		32.8 (7.36)•			
RP_S	-42.0 (-5.94)•		-41.3 (3.17)•	-0.547 (-0.05)	-33.9 (-3.44)•			
RP_F	-14.5 (-5.20)•		-18.5 (-4.41)•	-10.5 (-1.70)'	-20.0 (-5.20)•			
ER_P		0.010 (0.47)	-0.026 (-1.11)		-0.031 (-1.43)			
ER_S		-0.017 (-0.46)	-0.022 (0.88)	-0.024 (-0.56)				
ER_F		-0.132 (-0.24)	0.467 (0.79)	0.795 (1.06)	0.836 (2.00)'			
P_P						5066.1 (3.57)•		4839.0 (3.40)•
P_S						1433.3 (1.24)	875.3 (0.67)	
P_F						-249.7 (-3.51)•	-62.6 (-1.12)	-206.2 (-3.29)•
D_P	-3.3 (-3.62)•	-1.05 (-0.59)	-2.5 (-2.44)°		-2.4 (-2.40)°	-3.8 (-2.24)°		-3.2 (-1.93)'
$\bar{R}^2$	0.984	0.947	0.985	0.952	0.986	0.967	0.952	0.966
F	352.42	100.64	237.33	111.33	274.10	163.10	187.13	199.05
DW	1.458	0.847	1.688	0.900	1.770	1.508	0.904	1.244
LM	2.2(0.14)	10.2(0.00)	0.7(0.39)	7.7(0.01)	0.4(0.55)	1.5(0.23)	8.4(0.04)	3.7(0.05)

3.3.2.2. Analysis of the estimated results

In regressions 1, 2 and 3 of all equations, collinearity problems are likely to arise due to the high correlation between the income variable and the relative prices of Portugal and Spain (respectively, 0.96 and 0.94 in the log form and 0.97 and 0.93 in the linear form), between the Portuguese exchange rate and the relative prices of Portugal and Spain (respectively, 0.97 and 0.99 in the log form and 0.96 and 0.98 in the linear form) and between the relative prices of Portugal and Spain (0.99 in the log form and 0.97 in the linear form). These problems might be expressed through unexpected signs and individual non-significance of the estimated coefficients, associated with overall significance of the model and relatively high R^2 values. Collinearity is rather apparent in regressions 2 and 3 for Spain. Indeed, regressions 2 and 3 for Spain are overall significant and have high explanatory power; however, most of the coefficients are individually non-significant and some of their signs are not as expected. Yet, these symptoms seem to disappear when effective prices, instead of relative prices and exchange rates, are included as regressors.

In the presence of collinearity, the regression coefficients tend to show large standard errors implying that they cannot be estimated with great precision. Yet, as collinearity violates no classic assumption, it can be shown that unbiased consistent estimators still occur. So, even in the presence of severe collinearity, the OLS estimators still remain BLUE (best linear unbiased estimators). The only direct effect of multicollinearity is to induce large standard errors. But, this is also the effect of small sample sizes and independent variables with little variability. The sample size and the variability of the regressors depends on the data availability and not on the researchers' choice. Therefore, as Achen (1982, p.83) points out, the question of 'what should be done about multicollinearity?' is equivalent to the question 'what should be done if the sample size is small?' for which "*no statistical answer can be given*".[12]

Since multicollinearity is a feature of the sample and not of the population "*we do not 'test for multicollinearity' but can, if we wish, measure its degree in any particular sample*" (Kmenta, 1986, p.431). There are no *methods* of detecting multicollinearity; "*what we have are rules of thumb, some informal and some formal, but rules of thumb all the same*" (Gujarati,

[12] According to Leamer (1983, p. 300-1), high collinearity is "*a fact of life*", not a 'problem'; and *ad hoc* solutions that have been used such as stepwise regression and ridge regression "*can be disastrously inappropriate*".

p.335). These rules include the recognition of the 'symptoms' described above (individual insignificance of regressors associated with high R^2 values and overall significance of the regression, wrong signs of the coefficients and high sensitivity of the estimates to small changes in the sample data), as well as some measurement entities such as the coefficient of correlation between regressors and the variance-inflating factor (VIF) defined as $VIF=1/(1-r_{ij}^2)$, where r_{ij}^2 is the coefficient of correlation between the explanatory variables i and j. As r_{ij}^2 approaches unity, the VIF approaches to infinite giving a measure of the speed with which the variance and covariance of the estimators increase with collinearity. Other measures include the high pair-wise correlation, the condition index and the tolerance index (see Gujarati, 1995).

The Durbin-Watson (DW) statistic and/or the Lagrange multiplier (LM) test for serial correlation, detect this problem at the 5% significance level in regressions 1, 2 and 4 for Portugal, and in the log form of regression 1 and all regressions 2 for France. Therefore, no reliable conclusions can be derived from the estimates of these specifications. The DW test is inconclusive in all regressions for Spain. However, the LM test detects serial correlation at the 5% level in regressions 2 and 4 in the linear form and in regression 2 in the log form for Spain. No serial correlation is detected in all regressions 3, and 1 and 4 of the log form for Portugal; in regressions 1 and 3 of the linear form for Spain or in regressions 3 and 4 of the log form and in regressions 1, 3(A), 3(C) and 4(A) of the linear form for France. Hence, only regressions 3 for Portugal in both functional forms, 1 and 3 in the linear form and 1, 3 and 4 in the log form for Spain, and regressions 1, 3 and 4 in the linear form and 3 and 4 in the log form for France are considered as providing the best results, given the reservation that a static single equation may not be the ideal means to conduct a reliable analyses of the UK tourism demand.

The following analysis is carried out at several levels: first, each destination country is considered separately; then, only the reliable regressions, according to the econometric quality criteria described above, are considered; finally, the interpretation of the coefficients' estimates is given separately for the log-linear and for the linear functional forms. Although for both functional forms, the estimates obtained with the per capita and aggregate specifications do not differ much, the goodness of fit and overall significance are slightly superior for the models with the aggregate values of the expenditure and income variables. Hence, the analysis focuses on the aggregate form of the models.

Demand Equation for Portugal

The log form of regression 3 for Portugal is significant overall, with an F statistic of 15.39. The model explaining 80% of the UK demand variations. Except for the Portuguese exchange rate, all other explanatory variables are individually significant at the 5% level or less and the signs of the coefficients' estimates are as expected. The estimated income elasticity for UK tourism demand in Portugal is 1.2, meaning that if the UK real disposable income increases by 1%, the demand for tourism in Portugal increases by 1.2%, *ceteris paribus*. The UK demand for Portugal is expected to decrease by 1.7% for each 1% increase in its relative price, *ceteris paribus*. The relative prices of Spain and France have a significant although opposite effect on the UK tourism demand for Portugal. While a 1% increase in the relative price of Spain induces an increase of 8% in the UK tourism demand for Portugal, the same price increase in France, causes a 2% decrease, *ceteris paribus*. This classifies tourism in Spain as a substitute and, in France, as a complement of tourism in Portugal. [13]

The Portuguese currency exchange rate is not significant and has the 'wrong' sign attached to its coefficient estimate. If significant, the negative sign on this coefficient would mean that the cheaper the escudo is to UK tourists, the less they would be willing to visit Portugal. If tourism in Portugal is a normal good, this result is incongruent. Yet, the non-significance of the Portuguese exchange rate means that UK tourists are not sensitive to changes in the relative value of the escudo, which is plausible since the escudo has always been a weak currency relative to sterling. Nevertheless, the coefficients of the exchange rates for Spain and France are significant at the 1% level, and their signs confirm the substitutability and complementarity of Spanish and French tourism, respectively. Hence, if the ratio between the Spanish (French) currency and the pound increases by 1%, meaning that the national currency would be weakening relative to sterling, the UK demand for Portugal would decrease (increase) by 2.1% (2.4%), *ceteris paribus*.

The interpretation of the coefficient estimate of the dummy variable is linked to that of the intercept and must take into account the fact that there is no logarithm attached to the variable. The intercept itself, has no sensible economic interpretation as it represents, in the

[13] These results appear to be inconsistent with empirical evidence. In addition, the estimated sensitivity of the UK tourism demand for Portugal to changes in the relative price of Spain seems excessive and can only be explained by the mis-specification problems stated above.

form $e^{1.61}$=5.00, the average UK tourism demand for Portugal in the year base (1990), when UK real disposal income is £1million and the value of sterling equals the values of all other currencies. However, in the period 1974-1979, the intercept shifts downward and this shift is significant. Hence, the value $e^{(1.61-0.46)}$=3.16 must be subtracted from the intercept value 5.00, which gives £1.84 million. This value can be interpreted as the average decrease in the UK tourism demand for Portugal over the period 1974-1979.

In the linear form of regression 3 for Portugal, no autocorrelation is detected. Both versions A and B, are significant overall and explain around 76% of the UK demand variations. Considering the version where the price and exchange rate of France are omitted,[14] the effect of an absolute unit increase in the UK income (£1 million) is 0.005, meaning that the UK demand for Portugal would increase by an average amount of £5000, *ceteris paribus*. The own-relative price impact is negative, as expected, and significant at the 1% level. However, its interpretation is not simple given the relative price variable's definition. The relative price is a ratio between two price indexes. Hence, if this ratio increases from, say, 1 in the year base to 1.01 in the following year, this would mean that prices in the destination increased by 1 percentage point (pp), given that prices in the UK remained constant. In this case, the impact of this relative own-price change on the UK demand for Portugal would be a decrease of £6.3 million, *ceteris paribus*. Similar interpretation can be given to the relative price coefficient for Spain: if prices in Spain increase by 1 pp, assuming that prices in the UK do not change, the demand for tourism in Portugal increases by £10.1 million.

The interpretation of the exchange rate coefficients also has to take into consideration the variable's definition. The UK demand for Portugal does not respond significantly to changes in the value of the pound relative to the escudo, but it does respond significantly to changes in the relative value of the peseta. Hence, if the value of the pound relative to the peseta increases by 1, the UK demand for tourism in Portugal is expected to decrease, on average, by £1.7 million, *ceteris paribus*. Given the magnitude of UK tourists' spending every year in these destinations, one peseta more or less per pound makes a considerable difference for the choice of the country to be visited.

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[14] In version A, the relative price and exchange rate of France are individually non-significant. Their omission improves the goodness of fit and overall significance of regression B. The F test confirms the omitted variables as irrelevant.

In the case of the linear functional form, the interpretation of the dummy variable's coefficient is direct. Its value means that the UK demand for Portugal in the period 1974-1979 experienced an average decrease of £134 million. This value seems extremely high and differs sharply from the equivalent information obtained with the log-linear form.

Comparison between estimates of the log-linear and linear forms of regression 3 for Portugal can be carried out if average elasticities are calculated for the linear form using the formulae given above. Table 3.13 shows the elasticities estimates obtained from regression 3 in the linear and log-linear forms.

Table 3.13: Elasticities estimates of regression 3 log-linear and linear forms for Portugal

Regression 3	RDI	RP_P	RP_S	RP_F	ER_P	ER_S	ER_F
Log-linear	1.22	-1.70	8.05	-1.98	-0.76	-2.13	2.40
Linear	1.40	-1.45	3.09		0.28	-1.09	

The elasticities estimates for the UK demand for Portugal vary considerably depending on which functional form is used. The elasticity estimate for the relative price of Spain is 8.1 in the log-linear form, and less than half that value in the linear form. The elasticity for the exchange rate of Spain is estimated to be around -2 with the log-linear form, and around -1 with the linear form. In the log-linear form, the relative price and exchange rate of France are indicated as relevant explanatory variables of the UK demand for Portugal and their coefficients qualify France as a complement destination of Portugal. Yet, in the linear form, neither of these coefficients is significant, indicating that UK tourism in Portugal is not sensitive to changes in the relative price and/or exchange rate of France.

Demand Equation for Spain

Regression 3 in the log-linear form shows several symptoms of collinearity: the majority of the variables are insignificant and/or present unexpected signs, the value of the adjusted R^2 is high (0.84) and the regression is significant overall with an F statistic value of 19.2. The estimated income elasticity is 1.64 and significant at the 5% level. The relative price of Portugal is significant at the 10% level and its coefficient estimate indicates that, *ceteris* *paribus*, a 1% increase in that price leads to a 0.87% decrease in UK demand for Spain. This indicates tourism in Portugal as a complement to, rather than a substitute of, tourism in Spain. The dummy's coefficient is significant at the 5% level but its magnitude is irrelevant.

In the log-linear form of regression 4, the symptoms of collinearity seem to disappear and, although still insignificant, the own-effective price coefficient has now the expected sign. Yet, it should be mentioned that the LM test would detect serial correlation at the 6% level. The model explains 82% of the dependent variable variability and it is significant overall with an F statistic of 26.4. Changes in the effective price of France have a significant and positive influence on the UK demand for Spain, indicating tourism in France as a substitute for tourism in Spain. Tourism in Portugal appears, once again, as a complement of tourism in Spain, as a 1% increase in the effective price of Portugal induces a decrease of 0.87% in the demand for tourism in Spain, *ceteris paribus*. The income elasticity is 1.94 and significant at the 1% level, but the dummy variable is now insignificant.

As mentioned above, relative prices and exchange rates are supposed to be relevant explanatory variables of the UK demand for tourism. However, in the case of Spain, exchange rates do not appear to have any relevance to the explanation of the UK demand for tourism. Regression 1 in the log-linear form does not include the exchange rates as regressors but, nevertheless, presents features of a good fit: it explains 86% of the dependent variable variations, is significant overall with an F statistic of 35, does not present evidence of serial correlation and, except for the own-relative price, all variables are individually significant and present the expected signs. The income elasticity is significant at the 1% level, positive and above unity. A 1% increase in the relative price of Portugal (France) induces a decrease (increase) in the UK demand for Spain of 0.81% (0.63%), *ceteris paribus*. The dummy variable coefficient is significant at the 5% level but its magnitude is not relevant. The estimate of the relative price elasticity for Spain is significant at the 1% level, positive and around 2.9. Yet, it indicates that increases in the relative price of Spain induce increases in the UK demand for Spain, which is an odd result. Collinearity and/or misspecification bias may be responsible for the unexpected sign, significance and magnitude of this coefficient.

The linear form regression 1, shows the income coefficient as significant at the 1% level. Its magnitude indicates the average increase in UK demand for Spain, measured in £million (£27000), per £1 million increase in the UK disposable income, *ceteris paribus*. The coefficients' estimates of the relative price of Portugal and France are significant at the 5% level, and their magnitudes indicate that a 1 pp increase in the relative price of Portugal (France), assuming prices in the UK unchanged, induces an approximate decrease (increase) of £14 (£9) million in the UK demand for Spain, *ceteris paribus*. These values indicate tourism in Portugal and in France to be, respectively, a complement to and substitute of tourism in Spain. The estimate of the dummy variable's coefficient is significant and indicates that the UK demand for Spain increased by £297 million in the period 1974-1977. The coefficient's estimate of the relative price of Spain is significant and positive, indicating the same peculiar result as before.

In the linear form of regression 3, symptoms of multicollinearity are detected, probably due to the inclusion of the three (irrelevant) exchange rate variables. The interpretation of the estimation results is similar to that given for regression 1. The elasticities estimates for the log-linear and linear forms of regressions 1, 3 and 4 are presented in the table 3.14.

Table 3.14: Elasticities estimates for the log and linear forms of regressions for Spain

	Regression	RDI	$(R)P_P$[1]	$(R)P_S$[1]	$(R)P_F$[1]	ER_P	ER_S	ER_F
Log-linear	1	1.58	-0.81	2.87	0.63			
	3	1.64	-0.87	2.31	0.94	0.23	0.10	-0.33
	4	1.94	-0.87	-0.62	1.28			
Linear	1	1.51	-0.66	2.35	0.71			
	3	1.29	-0.61	2.17	1.04	0.01	0.10	-0.55

[1] Indicates relative prices for regressions 1 and 3 and effective prices for regression 4.

Important differences can be observed in the estimates of the own-price elasticities, which range between –0.62 and 2.87, and in those of the price elasticity of France, which range between 0.63 and 1.28. The other elasticities estimates seem to be fairly similar.

Demand Equation for France

In the log-linear functional form, only regressions 3 and 4 present acceptable quality according to the usual criteria. In regression 3(A) the dummy variable and the relative price and exchange rate of Portugal are manifestly insignificant variables and were omitted from the model. Version (B) of regression 3 explains 98% of the variations of the dependent variable and is significant overall with an F statistic of 226. Except for the exchange rate of Spain, all variables are significant at the 1% or 5% levels, and their estimated coefficients have the expected signs. The income elasticity estimate is 1.5, indicating that an increase by 1.5% in the UK demand for France follows a 1% increase in the UK real income, *ceteris paribus*. The own-relative price elasticity estimate is 2.95, indicating that the UK demand for France decreases by around 3% for each 1% increase in the own relative price, *ceteris paribus*. The demand for tourism in France also responds significantly to changes in the relative price of Spain as a 1% increase in this price induces, *ceteris paribus*, an increase of 1.5% in the UK demand. Hence, Spain is a substitute destination of France for UK tourists. A close to unity elasticity estimate is associated with changes in the exchange rate for France suggesting that, if the French franc becomes 1% cheaper relative to the pound, the UK demand for tourism in France increases by around 1%, *ceteris paribus*.

In regression 4(B), all coefficients are significant at the 1% level. The model explains 95% of the variability of UK demand for France, and is significant overall with an F statistic of 186.5. However, it should be noted that the LM test for serial correlation would detect this problem at the 6% level, suggesting that the omitted variables may be relevant regressors. The income elasticity estimate is 2.8 meaning that, *ceteris paribus*, the UK demand for tourism in France increases by around 3% per 1% increase in the UK real income. The own effective price elasticity estimate is 2.1 which indicates that a 1% increase in the price of France induces a 2% decrease the in UK demand for France, *ceteris paribus*. The UK demand for tourism in France increases by 1.7% if prices in Spain increase by 1%. This result indicates Spain as a substitute destination of France.

The linear form of regressions 1, 3 and 4, present acceptable quality according to the usual criteria. In regression 1, no signs of correlation are detected, the model explains 98.5% of the variations of the dependent variable, it is significant overall with an F statistic of 374 and all the coefficients are individually significant at the 1% level, presenting the expected signs. A positive unity change (£1 million) in the UK disposable income has an estimated positive impact on demand of £13,000. The coefficient for the relative price of Portugal is highly significant and indicates that if prices in Portugal increase by 1 pp, while prices in the UK remain constant, the UK demand for tourism in France increases by £17 million, *ceteris paribus*. The coefficient sign indicates that Portugal is a substitute destination for France. Conversely, the estimated coefficient for the relative price of Spain indicates this destination as a complement to, rather than a substitute of, tourism in France. Indeed, a 1 pp increase in the relative price of Spain induces a decrease of around £25 million in the UK demand for France, *ceteris paribus*. However, if prices in France increase by 1 pp, the estimated decrease in the UK demand will only be £7.9 million, *ceteris paribus*.

Regression 3(B) excludes the dummy, relative price and exchange rate of Portugal. However, both the DW and LM tests detect serial correlation in this model indicating the possibility of relevant variables being omitted. Version (C) of this model omits only the exchange rate of Spain. Its increased explanatory power, overall significance and absence of serial correlation reveals this choice as correct. Except for the exchange rate of Portugal, all the other variables are individually significant at the 1% or 5% levels. The estimated coefficient of the exchange rate for France indicates that an increase in the value of the pound by 1 franc induces an increase in the UK demand for France by £48 million. Given the relatively strong value of the franc and the magnitude of the UK tourists' expenditure in France (e.g. £1958 million in 1997), a 1 franc per pound increase in the relative value of the UK currency is likely to make a significant difference to the decision to visit France. The economic interpretation of the other coefficients is similar to that given for regression 1.

In the linear form of regression 4, the omission of explanatory variables appears to cause serial correlation as detected by the DW and/or the LM tests. Hence, only version (A) seems to be acceptable. *Ceteris paribus*, the estimated impact of a £1million increase in UK disposable income is a £26 million increase in the demand for France. The effective price of Spain is not a significant explanatory variable in this model but the effective prices of Portugal and France are. The economic interpretation of these estimates is complicated due to the way in which the effective price variable is defined. Hence, an example seems to be appropriate to explain these estimates.

Consider the following values extracted from the data, for the relative price, exchange rate and effective price of France in the period 1990-1991:

	$RP_F = \frac{CPI_F}{CPI_{UK}}$	$ER_F = \frac{CUR_F}{CUR_{UK}}$	$P_F = \frac{RP_F}{ER_F}$
1990	1.000	9.72	0.1029
1991	0.975	9.98	0.0980

The relative price of France (RP_F) is defined as the ratio between the price index of France and the price index of the UK. The exchange rate of France (ER_F) is defined by the amount of francs that 1 pound sterling can buy. The effective price of France (P_F) is defined by the ratio between the relative price and the exchange rate of France. Consider the exchange rate of 1990 (9.72 francs per £1), and assume that in 1991 the relative value of the franc increased by 1 pp. Then, in 1991, £1would be 9.6228 (9.72-0.0972=9.6228) francs as a 1pp increase means an increase of 0.0972 in the franc's relative value. If no change occurs in the relative price of tourism in France, the effective price in 1991 would be 0.10392. Hence, the change in the effective price of France in 1991 relative to its value in 1990 is 0.10392–0.1029=0.00102. This change multiplied by the coefficient's estimate of the effective price of France (-13625.9) is -£13.9 million. Hence, -13.9 represents the impact on the UK tourism demand for France when the relative value of the franc increases by 1 pp, given that the own-relative price did not change. On the other hand, assume that the exchange rate is constant but the relative price of France increases by 1 pp. Then, its value increases from 1 in 1990, to 1.01 in 1991. The effect of this variation in the effective price is 1.01/9.72=0.10391. Then, the effective price would increase by 0.00101 from 1990 to 1991. This increase multiplied by the estimated coefficient is -£13.9 million, representing the impact on the UK demand for France of a 1pp increase in the relative price of France, given the exchange rate. Therefore, changes in the exchange rate and in relative prices cause changes in the effective price which impact on demand in the same way.

Consider now the following values of the same variables for Portugal:

	$RP_P = \frac{CPI_P}{CPI_{UK}}$	$ER_P = \frac{CUR_P}{CUR_{UK}}$	$P_P = \frac{RP_P}{ER_P}$
1990	1.000	254.41	0.0039
1991	1.052	255.64	0.0041

Assume that the value of the escudo relative to the pound increases by 1 pp from 1990 to 1991. Hence, the value in 1991 is 251.87 escudos per pound. Given a constant relative price of Portugal, its effective price in 1991 is now 1/251.87=0.00397. The change in the effective price would then be, 0.00007. This change multiplied by the estimated coefficient of the effective price of Portugal (277328.2) is £19.4 million. This value represents the impact on the UK tourism demand for France when the exchange rate of Portugal increases by 1 pp, provided that the relative price does not change. On the other hand, if the exchange rate is constant but the relative price of Portugal increases by 1 pp, its value will be, say, 1 in 1990 and 1.01 in 1991. The effect of this variation in the effective price would be 1.01/254.41=0.00397. Then, the effective price increases by 0.00007 as before. This change, multiplied by the coefficient's estimate, is £19.4 million which is the same impact on the UK tourism demand for France caused by an increase of 1 pp in the exchange rate of Portugal. Table 3.15 shows the elasticities estimates obtained with the log-linear and the linear functional forms of regressions for France.

Table 3.15: Elasticities estimates for the log and linear forms of regressions for France

	Regression	RDI	$(R)P_P$[1]	$(R)P_S$[1]	$(R)P_F$[1]	ER_P	ER_S	ER_F
Log-linear	3 (A)	1.47	0.09	0.26	-2.56	0.32	-0.45	0.62
	3 (B)	1.51		1.50	-2.95		-0.64	1.03
	4 (A)	3.13	-0.66	1.46	-1.43			
	4 (B)	2.80		1.66	-2.10			
Linear	1	1.18	1.30	-2.46	-1.02			
	3 (A)	1.18	1.48	-2.40	-1.32	-0.29	0.29	0.30
	3 (C)	1.27	1.43	-1.94	-1.44	-0.36		0.57
	4 (A)	2.36	1.30	0.44	-1.77			

[1] Indicates relative prices for regressions 1 and 3 and effective prices for regression 4.

The estimates of the log and linear forms in regressions for France present important discrepancies. For example, the income elasticity ranges between 1.18 and 3.13. The price of Portugal variable is not significant in the log form but is significant in the linear form. The elasticity estimates of the price of Spain range between -2.46 and 1.66. The own price elasticity estimate ranges between 1.02 and 2.95. The estimate for the exchange rate of France elasticity ranges between 0.30 (non-significant), and 1.03 (significant).

An overall view of the models' estimation results, is provided in tables 3.16, 3.17 and 3.18, showing the elasticities' estimates obtained with each regression. Table 3.16 presents the results for Portugal, table 3.17 for France and table 3.18 for Spain.

Table 3.16: Elasticities estimates of the demand equation for Portugal

PORTUGAL			RDI	$(R)P_P$	$(R)P_S$	$(R)P_F$	ER_P	ER_S	ER_F
1	Log-linear	Aggregate	1.32	-1.42	3.54	-0.21			
		Per capita	1.40	-1.43	3.52	-0.20			
	Linear	Aggregate	1.12	-0.98	2.05	-0.13			
		Per capita	1.39	-1.02	2.04	-0.11			
2	Log-linear	Aggregate	0.57				0.20	-0.35	0.42
		Per capita	0.63				0.18	-0.34	0.47
	Linear	Aggregate	-0.14				0.23	0.04	-0.31
		Per capita	-0.11				0.19	0.05	-0.23
3	Log-linear	Aggregate	1.22	-1.70	8.05	-1.98	-0.76	-2.13	2.40
		Per capita	1.23	-1.68	8.05	-1.99	-0.77	-2.13	2.40
	Linear	Aggregate (A)	1.39	-1.39	3.78	-0.34	0.03	-1.40	0.64
		Aggregate (B)	1.39	-1.45	3.09		0.27	-1.09	
		Per capita (A)	1.67	-1.38	3.77	-0.35	-0.01	-1.43	0.65
		Per capita (B)	1.39	-1.44	3.08		0.26	-1.07	
4	Log-linear	Aggregate (A)	0.58	-1.00	0.77	0.13			
		Aggregate (B)	0.54	-0.94	0.82				
		Per capita (A)	0.55	-1.02	0.77	0.15			
		Per capita (B)	0.50	-0.96	0.82				
	Linear	Aggregate (A)	0.83	-0.86	0.86	0.17			
		Aggregate (B)	0.28	-0.79	0.91				
		Per capita (A)	0.82	-0.90	0.84	0.19			
		Per capita (B)	0.28	-0.82	0.90				

Table 3.17: Elasticities estimates for the demand equation for France

FRANCE			RDI	$(R)P_P$	$(R)P_S$	$(R)P_F$	ER_P	ER_S	ER_F
1	Log-linear	Aggregate	0.98	0.48	0.08	-2.35			
		Per capita	0.98	0.47	0.08	-2.35			
	Linear	Aggregate	1.18	1.30	-2.46	-1.02			
		Per capita	1.10	1.32	-2.41	-1.07			
2	Log-linear	Aggregate	1.48				0.80	-0.34	-0.88
		Per capita	1.51				0.80	-0.33	-0.89
	Linear	Aggregate	2.91				0.07	-0.21	-0.04
		Per capita	3.02				0.11	-0.20	-0.09
3	Log-linear	Aggregate	1.47	0.09	0.26	-2.56	0.32	-0.45	0.62
		Per capita	1.51		1.50	-2.95		-0.64	1.03
		Aggregate (A)	1.43	0.14	0.26	-2.61	0.28	-0.44	0.64
		Aggregate (B)	1.52		1.52	-2.95		-0.63	1.02
	Linear	Per capita (A)	1.18	1.48	-2.40	-1.32	-0.29	0.29	0.30
		Per capita (B)	1.27	1.43	-1.94	-1.44	-0.36		0.57
		Aggregate (A)	1.19	1.46	-2.37	-1.37	-0.28	-0.26	0.31
		Aggregate (B)	1.19	1.43	-1.94	-1.48	-0.33		0.56
4	Log-linear	Per capita (A)	3.13	-0.66	1.46	-1.43			
		Per capita (B)	2.80		1.66	-2.10			
		Aggregate (A)	3.27	-0.53	1.45	-1.54			
		Aggregate (B)	2.98		1.63	-2.07			
	Linear	Per capita (A)	2.36	1.30	0.44	-1.77			
		Per capita (B)	2.55	1.24		-1.45			
		Aggregate	2.48	1.36	0.46	-1.80			
		Per capita	2.66	1.30		-1.49			

Table 3.18: Elasticities estimates for the demand equation for Spain

SPAIN			RDI	$(R)P_P$	$(R)P_S$	$(R)P_F$	ER_P	ER_S	ER_F
1	Log-linear	Aggregate	1.58	-0.81	2.87	0.63			
		Per capita	1.61	-0.78	2.80	0.63			
	Linear	Aggregate	1.51	-0.66	2.35	0.71			
		Per capita	1.52	-0.63	2.30	0.70			
2	Log-linear	Aggregate	1.12				0.03	0.43	-0.09
		Per capita	1.18				0.01	0.43	-0.07
	Linear	Aggregate	0.67				0.13	0.56	-0.34
		Per capita	0.62				0.13	0.55	-0.33
3	Log-linear	Aggregate	1.64	-0.87	2.31	0.94	0.23	0.10	-0.33
		Per capita	1.66	-0.82	2.27	0.92	0.19	0.12	-0.32
	Linear	Aggregate	1.29	-0.61	2.17	1.04	0.01	0.10	-0.55
		Per capita	1.29	-0.59	2.15	1.03	0.01	0.23	-0.54
4	Log-linear	Aggregate	1.94	-0.87	-0.62	1.28			
		Per capita	2.02	-0.82	-0.64	1.25			
	Linear	Aggregate	1.85	-0.98	-0.60	1.48			
		Per capita	1.97	-0.92	0.64	1.43			

The differences in the magnitude of the elasticities estimated from models with variables defined in aggregate terms, and models with variables defined in per capita terms never exceed 0.3. This suggests that, in terms of elasticities estimates, there is practically no difference between models using aggregate or per capita variables, either in the log-linear or in the linear forms. Then, at least from this point of view, the results are consistent. However, there are striking disparities in the estimates when different sets of regressors or different functional forms are considered. For example, in the equation for Portugal, the elasticity estimate for the price of Spain ranges between 0.77 and 8.05; in the equation for Spain, the own-price elasticity estimate ranges between -0.64 and 2.87 and in the equation for France, the estimates for the price of Spain and income range, respectively, between -2.46 and 1.66 and between 1.10 and 3.27.

Given that all regressions were estimated using the same sample period, the same data set, the same origin and destination countries, within the same static single equation framework, the discrepancies found in the estimation results do not permit strong or convincing inferences about the UK tourism demand behaviour in these destinations. Furthermore, as the time series involved in the regressions are nonstationary, the estimation results may be spurious and the statistical inference invalid. Hence, any policy directives undertaken on the basis of these results would be controversial.

Nonstationarity aside, the results suggest that the functional form adopted, the set of regressors included and the definition of the price variables play a crucial role in the magnitude, signs and significance of the coefficients' estimates. Consequently, comparison between these results and those obtained in other studies is likely to be qualified as a futile and meaningless exercise. Nevertheless, even if only for pedagogical reasons, there are important lessons to be learned from the comparison of the methodologies used, the interpretation of results offered and the conclusions inferred from studies concerned with the investigation of tourism demand. Our next task is to analyse several of these studies and, when possible, to compare methodologies, results and conclusions, as well as their implications.

3.3.3. COMPARISON AND IMPLICATIONS OF THE ESTIMATED RESULTS

The models selected for analysis must have similarities both between themselves and with the models used in this study, as consideration of completely different methodologies serves no fruitful purpose. Hence, the selection focuses on studies of the UK demand for tourism within a static single equation approach, using comparable functional forms and sample periods, and similar explanatory variables which include relative or effective prices and/or exchange rates. Under these sets of conditions, the studies selected are Witt's (1980b) analysis of the UK demand for tourism in 16 main destination countries; Loeb's (1982) examination of the UK tourism demand in the US; Uysal and Crompton's (1984) study of the UK tourism demand for Turkey; Papadopoulos and Witt's (1985) estimation of the UK demand for tourism in Greece; Gunadhi and Boey's (1986) analysis of the UK demand for tourism in Singapore and Witt and Martin (1985) and Martin and Witt's (1987) studies of the UK demand for tourism in several major destinations.

Witt (1980b) examines the UK demand for tourism in sixteen destinations, using pooled data for the period 1965-1972, the log-linear functional form and the OLS method of estimation. Witt's study separates tourism demand by mode of travel (air and sea/land) and by type of holiday (independent and inclusive), using dummy variables. The dependent variable is measured by the number of visits and the income variable is measured in per capita terms. The coefficients' estimates for some of the explanatory variables in three of the specified models are displayed in the following table:

MODELS	Dependent variable	Per capita income	Relative price	Relative cost of tourism	Cost of travel	Travel time	Holiday type	Travel mode	Lagged dependent variable
Model 1	Visits	0.52	-0.24		-0.74*	-1.00*	1.39*	1.32*	
Model 2	Visits	1.39*		-0.05	-0.20	-0.17*	0.16	0.15	0.91*
Model 3	Δ Visits	1.45		-0.69*	0.36	-0.27			

* indicates significant at the 5% level

In Model 2, travel time, per capita income and the lagged dependent variable seem to provide all the explanation the UK demand for tourism needs. No other explanatory variable is significant. Hence, the cost of travel, holiday type and travel mode, which have a significant impact on the demand for tourism in Model 1, vanish as determinants of that same demand behaviour in Model 2. Moreover, the impact of the travel time variable in Model 2 is less than 20% of its estimated value in Model 1. It is possible that the level variables used in Models 1 and 2 share trends which make them spuriously related. In Model 3, first differences of the variables are used, in what seems an attempt to avoid the problems of nonstationary data. In this case, however, the long-run relationship between the UK demand and its determinants is difficult to analyse, since models in differences are intended to depict short-run behaviour. The first difference specification indicates a significant role for the tourism cost variable, but no other variable is statistically relevant.

Loeb (1982) studies the UK tourism demand for the USA using time series data for the period 1961-1978, a log-linear functional form and OLS. The dependent variable is real tourism expenditure. Loeb specifies two alternative models: using aggregate expenditure and income; and using per capita terms for the dependent variable and the income variable. Some of the estimated elasticities are reported in the following table:

Models	Dependent variable	Per capita real income	Aggregate real income	Relative price	Exchange rate	R^2
Model 1	Per capita expenditure	1.04*		-6.36*	4.07*	0.95
Model 2	Aggregate expenditure		0.87	-5.25*	2.64*	0.86

* indicates significant at the 5% level

The income variable is significant when measured in per capita terms but not so when measured in aggregate terms. The estimates for the exchange rate elasticity are considerably different in both models. If the R^2 is the main criterion for measuring the quality of a regression, the specification with per capita expenditure and income is superior. However, the R^2, by itself, is not a dependable criterion of statistical quality. In contrast with the results obtained from the models estimated previously, Loeb's results show that variables measured in aggregate or per capita terms can produce substantial differences in the coefficients estimates. Given that per capita variables can be obtained by simple division of the variables' aggregate levels by the UK population and that the UK population did not vary much over the sample period considered, the estimates and goodness of fit differences reported may indicate misspecified regressions.

Papadopoulos and Witt (1987) estimate the UK demand for tourism in Greece for the period 1972-1982 using a log-linear functional form, the Cochrane-Orcutt (CO) estimation method and number of visits per capita of the UK population as the dependent variable. Their explanatory variables are: per capita income, relative cost of tourism, travel cost, advertising expenditure and a dummy variable. Their estimation results are reported in the following table:

Dependent variable	Per capita real income	Relative cost of tourism	Travel cost	Advertising expenditure	Dummy	DW statistic	R^2
Number of visits per capita	6.67*	-1.67	-0.28	0.26	-0.40	2.49	0.92

* indicates significant at the 5% level

The only relevant variable in this regression is per capita income and this fact alone speaks for the quality of the model. The CO method is used to deal with serial correlation problems detected previously. However, if the serial correlation detected is due to misspecifications in the model, the CO method does not solve the problem. Furthermore, the DW statistic is not a valid method of detection whenever the CO estimation is performed. With such a specification and few degrees of freedom for the estimation procedure, the magnitude of the R^2 is meaningless, as an R^2 of 0.30 with 1000 degrees of freedom is always preferable to a R^2 of 0.99 with 10 degrees of freedom.

Uysal and Crompton (1984) estimate the UK demand for tourism in Turkey for the period 1960-1980 using a log-linear functional form. The dependent variable is defined in two alternative ways: "number of tourists visiting Turkey" and "expenditure in Turkey by tourists". The model with the first dependent variable (Model 1) is estimated with OLS while the model with the second dependent variable (Model 2) is estimated with CO. The independent variables in both models are per capita income, relative price index, exchange rate and promotional expenditure. The estimation results are presented in the following table:

Models	Dependent variable	Per capita income	Relative price	Exchange rate	Promotional expenditure	R^2	DW
Model 1	Number of tourists	-0.064	-1.57*	1.68*	0.28*	0.87	1.64
Model 2	Total expenditure	2.09	-1.49	2.84	0.28	0.93	1.59

* indicates significant at the 5% level

According to the authors "*whenever serial correlation was detected by the DW statistic, a Cochrane-Orcutt procedure was used in an attempt to alleviate the problem*" (p. 293). Yet, misspecification bias cannot be "alleviated" by the use of CO procedure. Strong collinearity is likely to be present in Model 2 where all coefficients are individually insignificant. In an attempt to improve the model, the authors omit the variable 'promotional expenditure' (PE). However, omitting this variable does not seem to be the solution in spite of the authors' claim "*the removal of variable PE increased the coefficients of the income and the exchange rate variables*" (p.296). The word "*increase*" in the text refers to the magnitude of the coefficients' estimates but not their significance. An increase in magnitude can occur even when the omitted variable is relevant which would mean that, instead of solving the problem, it makes it worse by the omission of a relevant variable. The DW statistic in Model 2 is meaningless since the CO procedure is used. In Model 1, the income elasticity estimate is insignificant with the wrong sign, and the DW statistic indicates an inconclusive test for the detection of serial correlation. Inconclusive means doubt, not certainty! Hence, the authors' remark, "*that most of the estimated equations were free from autocorrelation*" (p. 296), is not enough to dismiss the presence of serial correlation in 15 out of 20 equations for which the DW statistic was in the inconclusive zone.

Gunadhi and Boey's (1986) study the UK tourism demand for Singapore in the period 1965-1981, with a log-linear functional form and OLS. The dependent variable in the model is tourist arrivals and the explanatory variables are real per capita income, relative shopping prices, relative hotel prices and exchange rate. Their estimates are showed in the following table:

Models	Dependent variable	Per capita income	Relative shopping prices	Relative hotel prices	Exchange rate	R^2	DW
Model 1	Tourists arrivals	3.74*	-0.41	-0.20	-0.01	0.94	1.27
Model 2	Tourists arrivals	7.30*				0.91	1.91

* indicates significant at the 5% level

In Model 1, only the income elasticity estimate is significant which, considering the high R^2 value, may indicate the presence of multicollinearity. However, the authors' solution for the problem of individual insignificance is to estimate a second model (Model 2) using a stepwise estimation procedure in which all explanatory variables but one, the per capita income, are omitted. The authors then claim that "*the DW statistic proved conclusive only for Indonesia and the UK, implying that for these two equations neither a misspecification nor an omission of significant variables has been made*" and they add, four lines below, this contradictory statement: "*Although the DW statistic for the UK equation indicates the rejection of first-order autocorrelation, the relatively low R^2 value and the significance of income alone, suggests that* *may be other explanatory variables capable of increasing the explanatory power of the model. ... their omission has led to their effects being picked up by the income variable thus inducing an over-estimate of income elasticity*" (pp. 245-46). Yet, one could say, instead, that their omission may have led to misspecification bias, rendering the model invalid and useless for inference proposes. No valid conclusions can be retrieved from such models.

Witt and Martin's (1985) estimate the UK tourism demand for several destinations in the period 1965/8-1983, using a log-linear functional form and the OLS and CO methods. As in Witt (1980), the authors separate the UK tourism demand by mode of travel (air and surface) and by type of holiday (independent and inclusive). The dependent variable is visits per capita and the explanatory variables are per capita real income, relative cost of tourism, cost of travel, exchange rate, lagged dependent variable and a trend. Their estimation results for destination countries France, Italy and Spain are presented in the following table.

Dependent variable	Destination	Per capita income	Relative cost	Cost of Travel	Exchange rate	Lagged (V/H)	Trend	DW	R^2
Visits/head (V/H) independent, by air	France OLS	1.43*	-0.21	-1.16	0.75	0.31		2.06	0.82
	Italy OLS	2.69*	-0.32		1.08*		-0.05*	2.00	0.93
	Spain OLS	1.46	-0.35	-0.02	0.72*	0.62*		2.00	0.97
Visits/head (V/H) inclusive, by air	France CO	3.81*	-0.12	-0.11	1.23				0.74
	Italy CO	5.55*	-0.11	-0.03	0.77*		0.11*		0.88
	Spain CO	0.87	-0.75*			0.52*			0.91
Visits/head (V/H) independent, by surface	France CO	1.35*	-0.44*			0.72*			0.97
	Italy CO	0.94	-0.51	-0.36		0.46*	-0.04		0.55
	Spain CO	0.73	-0.46	-1.60*		0.31			0.66
Visits/head (V/H) inclusive, by surface	France CO	2.54*	-0.91			0.67*			0.94
	Italy OLS	2.12*	-0.54			0.82*		2.08	0.64
	Spain CO	2.88	-0.46	-0.08	3.68*				0.74

* indicates significant at the 5% level

The inclusion of an explanatory variable in one equation and its omission from another similar one, is not explained by the authors and can only be viewed as related to the statistical performance of the models, which is not a sound basis for the use of different sets of variables. Such criteria for excluding or including variables "*leave a puzzle over economic interpretation that requires resolution if the studies are to provide insights into economic behaviour, and if they are not to degenerate into a trawling of the data simply to obtain the best 'fit'*" (Johnson and Ashworth, 1990, p. 149). Again the CO estimation method seems to be used to 'correct' serial correlation which may be linked to misspecified regressions.

The results reported in the studies selected for comparison and in many others, differ considerably when the author(s) eliminate, add, and/or change the definition of one or more variables. However, no clarifying comments are included to justify satisfactorily such differences. In addition, when serial correlation is detected there is no discussion of the possibility of its being caused by misspecification errors. The serial correlation detected is often assumed to be inherent to the disturbances and, therefore, 'corrected' by the Cochrane-Orcutt estimation method. In some studies, the dependent variable is measured in aggregate terms while the income variable is considered in per capita terms. For the sake of consistency, expenditure and income should both be measured in the same terms. In addition, some authors interpret insignificant variables as if they were significant determinants of the dependent variable, and others draw conclusions and policy implications from models that would not pass a more rigorous quality examination.

3.4. CONCLUSION

Research on tourism analysis has largely been focused on the demand side, attempting to establish its determinants and quantify the effects of changes in them on the dependent variable. The studies generally find the main determinants of tourism demand to be the origin's real income, exchange rates, relative prices, and a number of other qualitative and quantitative factors, depending on specific circumstances of the countries analysed. The literature shows that the majority of investigators use a static single equation approach to model tourism demand behaviour. These models generally include different origins and destination countries, different sample sizes, different measurement criteria and definitions for the variables involved, and different estimation methods. In spite of these differences the researchers' common aim is to estimate the sensitivity of tourism demand to changes in its determinants. Consumers, firms and governments depend on the accuracy of such information for decision and policy making. However, the precision of the estimates and the validity of inference and forecasting procedures depends crucially on the robustness of the theoretical framework underlying the specifications and on the use of a sound econometric methodology for the modelling, estimation and evaluation of the quantitative relationships specified.

The main contribution of this chapter is to show how small differences in the specifications of tourism demand within a static single equation context can affect the coefficients' estimates magnitudes, signs and significance, providing inconsistent results upon which no reliable conclusions or policy implication can be based. Taking the UK as an origin and France, Spain and Portugal as destinations, we modelled the origin's demand for tourism using different functional forms (the log-linear and the linear forms), different definitions for the dependent variable and income variable (aggregate and per capita values), and different sets of explanatory variables (exchange rates alone, relative prices alone, exchange rates and relative prices together and effective prices alone). A wide range of elasticities estimates was obtained by simply changing the functional form of the models or by changing the set of explanatory variables, as in the case of using relative prices and exchange rates separately or, instead, using them combined to form effective prices. This instability of the estimates was further confirmed when earlier studies of tourism demand were analysed. While in some studies, real income has a positive and significant effect on demand, in others, this regressor is insignificant and/or has a negative impact on the dependent variable; while some estimations present exchange rates alone as significant regressors, others claim this role for relative prices alone and still others find that both should be included as relevant explanatory variables. Even when studies do agree on the significance and sign of the determinants' effects for similar origin/destination pairs, their magnitudes can differ so sharply that neither a consistent comparison can be established nor a sound judgement can be made from the estimation results. Moreover, either evidence of collinearity and autocorrelation is ignored and insignificant results are interpreted as significant, or misspecification bias, such as omission of relevant variables, are 'blindly' tackled as true serial correlation and 'solved' by the application of alternative estimation methods such as the Cochrane-Orcutt.

Which variables should be included? Which functional forms should be used? Which estimates should be trusted? Econometric models are formal quantitative relationships which link theory and data to allow for the understanding of economic behaviour. The building of econometric models is generally associated with the objective of providing a consistent description of the data generating process and a reliable means of predicting its behaviour. However, by nature, models are simplifications of "the real thing" and as such, they can differ radically in their empirical relevance. That is, they are not equally useful or reliable to portray the phenomena they intend to explain. Indeed, as Leamer (1987, p.1-2) points out, "*models, stochastic or otherwise, are merely metaphors. More importantly, from a practical standpoint, when sensitivity analysis does indicate that inferences depend substantially on the choice of the metaphor, doubt about the inferences can be relived only by eliminating altogether the problem's metaphor*".

From the analysis of results provided by the modelling and estimation exercises of this chapter and by those of numerous other studies, the main conclusion is that the estimated magnitudes, signs and significance of the variables' coefficients and the general statistical quality of the models, depend crucially on the functional form adopted, on the definition of the variables included, and on the set of the regressors considered. In other words, within a static single equation approach, inference seems to depend "on the choice of the metaphor".

Static single equation models of tourism demand tend to neglect interdependencies among competing destinations, ignore problems arising from nonstationary data, overlook dynamics and lack an explicit theoretical basis within which consumers' preference structure can be appropriately modelled. Hence, these models can only be viewed as specific cases within the wide range of plausible demand behaviours, but not as comprehensive or reliable descriptions of tourist's general conduct. Modelling procedures constrained by these methodological faults are bound to produce inadequate empirical specifications which generate biased and inconsistent estimation results. Perhaps, 'elimination altogether' of the static single equation 'metaphor' is indispensable if research is to proceed in more valuable directions.

The features of consumer behaviour generally omitted from static single equation modelling are, on the one hand, the specification of a dynamic structure which clearly separates short- and long-run effects in the demand functions and, on the other hand, the derivation of demand functions from an explicit economic theoretical basis which permits both the formal testing of utility theory hypotheses and the consideration of cross influences among destinations. In chapter 4, we address the first omission integrating the dynamics of UK tourism demand in a single equation error-correction model. In chapters 5, we deal with the second omission by means of a static system of equations approach. In chapter 6 both features are included in a dynamic system of equations model.

CHAPTER 4

DYNAMIC SINGLE EQUATION MODELLING OF THE UK TOURISM DEMAND

4.1. INTRODUCTION

The first objective of econometric modelling is to provide a coherent explanation of the observed behaviour of economic variables. Understanding relationships between economic phenomena involves a process which relates theoretical ideas and empirical data within a quantitative framework. In this process a number of decisions have to be made starting with the choice of which idea, among alternative sensible ones, is going to be tested against a set of data using quantitative empirical models. Reliable statistical information is crucial in this process. Yet, however exact this information, by itself, it cannot explain economic phenomena. Nor can a theoretical causal structure, however clever and creative, be 'realistic' without the support of empirical evidence. Quantitative analysis of economic phenomena comprises a systematic search for the matching of theory and observation based on empirical econometric models. However "*all models are not born equal and we seek for those which are useful in practice*" (Hendry, 1995, p. 3).

Among the investigators concerned with understanding the how's and why's of economic behaviour, there appears to be some consensus about the abstract concept which defines what a model is, and which attributes are inherent in a 'good' model. Economic researchers seem to agree that economies are too large and too complex for the development of 'true' models. A model is "a free creation of the mind", "a metaphor", "a creative process", "a simplified representation", "an art". These qualifications establish the separation between the "*truth of reality*" and its caricature. That models are "*inevitably false*" seem to be uncontroversial. Indeed, even if the construction of true models were possible their useful handling would be doubtful. For example, a road map is a simplified representation of a complex network of motorways, railroads, countryside lanes, rivers, and mountains. If accurate enough and correctly interpreted, a road map can take one from A to B. The "real road", however, would not even fit inside the car.

Nonetheless, it is important to distinguish between theoretical and empirical models. Despite having the common objective of seeking for the "best" approximation of reality and despite their interdependence in practice (theory underlines the structure of empirical models and empirical findings can change theoretical postulates), theoretical models assume theoretical relationships among latent variables, while empirical models establish 'correspondence relationships' linking latent to observed and measured variables.

In research work, it is generally assumed that a true data generating mechanism exists within the complexity of the economy, and that the objective of the modelling process is to make statements about this mechanism as accurately as possible. The data generating process determines the properties of the data set which, in turn, determine the results obtained from estimating an empirical model. Hence, "*the use of observed data creates a fundamental distinction between theory and empirical models since empirical models must, by default, be simply a recombination of whatever process generated the data*" (Hendry and Richard, 1982, p. 6).

There also appears to exist a consensus among researchers in setting the characteristics and attributes of 'good' empirical models. For example, relevance, simplicity, theoretical plausibility, explanatory ability, accuracy of coefficients and forecasting ability are desirable properties of empirical models that Christ (1966, 1975) underlines. In more recent work of many other investigators such as Charemza and Deadman (1997), Granger (1990), Hendry (1987, 1995), Hendry and Wallis (1984), Leamer (1985) and Mizon and Richard (1986), we can find similar main and auxiliary criteria for qualifying an empirical model as useful and reliable. It seems generally accepted that a good empirical model should exhibit a good fit, absence of residual autocorrelation or heteroscedasticity (data coherency), valid exogeneity assumptions, parameter constancy, theory consistency, data admissibility and encompassing.[1] According to Hendry and Wallis (1984), an adequate empirical model has the ability to "*describe historical data without*

[1]The general to specific model-building approach, seeks not only to characterize data in a parsimonious way within a general theoretical framework, but also to provide a statistical basis against which other models can be evaluated. Encompassing is usually seen as the quality of econometric models which allows the investigator to see how well a given model accounts for the findings of rival studies. Hence, encompassing requires any given 'good' model to explain the results obtained by other models.

producing systematic misfit, to fit equally well to the future, to be consistent with the underlying theory and measurement system and to encompass alternative explanations of the same set of endogenous variables" (p. 6). In Hendry (1995), a broader, less technical and perhaps more ambitious description, states that good empirical models allow for the interpretation of reality within a simplified context and for the evaluation of the explanatory power of competing theories, permit the accumulation and consolidation of empirical knowledge and give a scientific approach to the understanding of human conduct. Summing-up, an adequate model seems to rest upon its ability to portray the past, explain the present, predict the future, encompass rival models and judge competing theories.

How to develop such econometric models from a given theoretical structure and a measurement system, and how to recognise and evaluate the desirable properties of a 'good' model, are neither obvious nor settled among investigators. A unique and consensual path to good empirical modelling does not exist. On the contrary, there are deep differences of opinion among researchers concerning the building, interpretation and evaluation processes of empirical models. In fact, several different modelling strategies co-exist in contemporary applied economics literature (see for example, Granger, 1990, for a survey). These questions seem to be linked to specific methodological issues engaged in the building, estimation and testing of empirical models.

Contemporary discussion of economic problems is heavily influenced by the results of empirical econometric analysis produced since the early 1940's within the so-called traditional approach. According to its critics, traditional econometric methodology appears to lead, in many cases, to models with poor forecasting ability, shaky inference procedures, dissociation between theory and empirical evidence, questionable assumptions and overall unreliable estimation results. Frequently, attempts to fit theoretical models to economic time series led to a number of statistical problems such as auto-correlated residuals (despite the assumption of independently distributed disturbances), 'wrong' signs, insignificance or doubtful magnitudes of coefficients and high collinearity among explanatory variables. Moreover, vital parameters in some models seem to be very unstable as the model specification changes. These problems are diagnosed and 'eliminated' without further consideration about possible reasons for their appearance in the first place. An example of this practice is the detection of residual auto-correlation with Durbin-Watson (DW) statistic and its subsequent 'elimination' by the application of the Cochrane-Orcutt estimation method. However, "*camouflaging the disease by 'removing' the symptoms seems an* *unlikely route to success* " (Hendry, 1983, p.197). The search for the best fit based on such criteria as the highest R^2 or 'significant' t-values is far from being an acceptable process of model building. For instance, Granger and Newbold (1974) show that spurious regressions involving independent random walks tend to present high R^2 and 'significant' t-values.

Unstructured "specification-search", as Leamer (1978) adequately renamed the harsh designation of "data-mining", may produce apparently good results but inappropriate statistical inference. A typical example is the fit of a polynomial in t of order N-1 to a time series of N observations. There are, of course, more subtle ways of specification-search but with similar practical results. R^2 and t-values must be viewed with caution and used as guidelines rather than precise ways of stating the quality of a model. Further discussion of the use and performance of models built within the traditional econometric methodology can be found in, for example, Hendry (1980), Hendry and Morgan (1995), Leamer (1983), Lovell (1983), Sims (1987) and Wallis (1989).

According to the critics of this methodology, the ongoing research seemed more concerned with the question of how best to estimate a model than with the process of its specification. However, even among the critics, there is some disagreement. While some point out that the credibility ascribed to any reported econometric estimate is related to the process by which it is obtained (Leamer, 1983b), others point out that a model's credibility does not dependent on its 'mode of discovery' but on whether it will survive latter evaluation (Hendry, 1987) and that "*the validity of any outcome is intrinsic to the product, not to its method of discovery or construction, methodology can at best reveal the benefits and drawbacks of alternative research strategies*" (Hendry, 1995, p. 10).

However, most critics seem to agree that the practical problems of model specification and selection, generally ignored in traditional literature, ought to be addressed. Hence, confronted with increasing specification uncertainty, modern methodological approaches change the emphasis from estimation to modelling. This change is apparent in recent applied work, where considerable attention is paid to the process of modelling economic time series using alternative strategies, and to the qualitative evaluation of econometric models. Little attention is currently paid to the issue of estimation methods which predominated in the literature for half a century.

The failure of traditional econometrics to produce satisfactory forecasts or resolve divergences between competing economic theories seems to be linked not to the use, but to the abuse of its methodological principles and to a widespread growth of senseless data-mining. Yet, it is possible to conduct a structured and purposeful search for the 'best' model avoiding the worst aspects of data mining. Acceptable specification-search is the process of moving from one model to a better one, using sensible evaluation criteria. However, if changes associated with this process are based on some extremes of data-mining, more often than not they lead the search away from the 'best' model's path. Nevertheless, it has to be recognised that, in practice, some specification-search is unavoidable. As pointed out in Charemza and Deadman (1997), the matter is not whether data-mining is involved in the modelling process, but how its sensible use may contribute to achieve satisfactory specifications. While model-search processes which deliberately ignore or conceal conflicting results are unacceptable, the purist's case of a judicious economic argument leading to a well-specified model, estimated and tested just once, is a nonsense. In the middle of these extremes lies appropriate econometric modelling.

The origins of modern methodology in model building processes are believed to be linked with the work of Davidson, Hendry, Srba and Yeo (1978), known in the literature as the DHSY paper. Since its publication, this work has been generally seen as an important influence on the way econometricians use time series data to model economic relationships. Indeed, some of the issues addressed in the DHSY paper have received considerable attention, leading to the development of new methodological approaches for the building and evaluating of econometric models, which include general-to-specific modelling, error-correction and vector autoregressive specifications, integration and cointegration analysis.

According to Hendry (1995), a 'data-based revision strategy' in model-building which can claim some empirical success, is the general-to-specific approach. Given a measurement system and a theoretical framework, a general model can be specified taking into consideration the sample size (which, *a priori*, constrains its generality), previous empirical findings (for example, nested special cases) and special data features (for example, lagged reactions or rapid adjustment processes). The estimation of the general model provides unrestricted parameters' estimates which can be tested against various pre-defined null hypotheses. Specific theoretical models are usually nested cases within the more general model and can be tested against it. Alternatively, the general model can be simplified, until a parsimonious consistent form results. This should then be tested for all the desirable attributes of a good model.

The existence of a potentially large number of theoretically plausible models which also satisfy the quality criteria renders model-choice a non-trivial problem as the criteria by which the desirable features of a good model are judged are necessary but not sufficient requirements. In addition, the fact that one model encompasses another under some criteria (forecasting, for example) does not necessarily imply that the latter should be retained and the former discarded, as they may serve different purposes. Nevertheless, if the next rival model does encompass an existing adequate model and still satisfy the mentioned criteria it can be qualified, for the time being, as a better model. Yet, as Hendry (1983, p.199) stresses "*until a model has been rigorously tested against new evidence it would seem hazardous to place much weight on its implications, no matter how 'pleasing' these seem*".

Although there are not sufficient conditions to ensure the finding of the 'perfect' empirical model, there are a number of necessary conditions which can be used to rule out inadequate models. This allows us to concentrate on the best remaining candidates. Given the unsatisfactory empirical results obtained in chapter 3, a change of model-building methodology is practically self-imposed. Hence, this chapter investigates the UK demand for tourism using some of the more recent econometric methods in the building, testing and evaluating of empirical models.

The structure of the chapter is as follows. Section 2, presents a succinct literature review of cointegration analysis in time series economic models. The order of integration of the variables included in the models is also addressed. In section 3, a general-to-specific approach is implemented to build econometric models with error-correction mechanisms for the UK tourism demand in France, Spain and Portugal. The estimation results obtained with these models are also provided. Section 4 presents a critical examination of estimation results obtained in different studies using the same methodology. Section 5 concludes.

4.2. ORDER OF INTEGRATION AND COINTEGRATION OF TIME SERIES IN UK TOURISM DEMAND FOR FRANCE, SPAIN AND PORTUGAL

4.2.1. BASIC CONCEPTS ON NONSTATIONARY STOCHASTIC PROCESSES[2]

The basic ideas of applied cointegration analysis are simple to understand and to use, although the underlying theory is not so straightforward. Thus, it may be useful to start with some basic concepts of stochastic processes and time series analysis.

[2] For a clear and simple explanation of the basic concepts of cointegration analysis see, for example, Harris (1995) or Charemza and Deadman (1997).

A stochastic process denoted by $\{Y_t\}$, where t represents time, is a family of real valued random variables Y_1, Y_2, ..., Y_t. For simplicity, let $\{Y_t\}$ be referred as Y_t. A stochastic process is said to be *stationary* (in a strict or strong sense) if the joint and conditional probability distributions of the process are unchanged if displaced in time. In most practical situations a weaker concept of stationary process is used, restricting the scope of the stronger definition to the means, variances and covariances of the process. Hence, a stochastic process Y_t is said to be (weakly) stationary if

$$E(Y_t)=\mu$$

$$E(Y_t-\mu)^2=\sigma^2$$

$$\gamma_k=[(Y_t-\mu)(Y_{t+k}-\mu)]$$

where γ_k is the covariance at lag k between Y_t and Y_{t+k}. If $k = 0$ then $\gamma_k = \sigma^2$. Consequently, if a stochastic process is stationary, its means, variances and covariances (at various lags) remain constant over time. If one or more of these conditions is not fulfilled, the process is said to be *nonstationary*.

An important special case of a nonstationary stochastic process is a *random walk*. A random walk stochastic process Y_t, can be described by the following equation:

$$Y_t = Y_{t-1} + Z_t \qquad \text{(i)}$$

where Z_t represents a series of identical and independent random variables. Another important special case is the nonstationary stochastic process denominated *random walk with a drift* which can be described as

$$Y_t = \mu + Y_{t-1} + Z_t, \quad \mu \neq 0 \qquad \text{(ii)}$$

where μ is a constant.

In the literature, the concept of a time series is often used alongside the concept of a stochastic process. Any time series can be viewed as being generated by a stochastic process, and a specific set of data can be regarded as a particular realisation of this stochastic process.

To make the notation compatible with that of most econometric textbooks, let us use y_t instead of Y_t to denote a time series stochastic process, and ε_t instead of Z_t to denote a series of identically and independently distributed continuous random error variables with zero mean, constant variance and no serial correlation. The stochastic process ε_t is called a *white noise*. With this new notation the equations $y_t = y_{t-1} + \varepsilon_t$ and $y_t = \mu + y_{t-1} + \varepsilon_t$, describe a random walk and a random walk with a drift, respectively.

The nonstationarity of an economic time series may be apparent when it is plotted against time by its propensity to move in an upward or downward direction. This tendency is called a *trend.* A time series can trend up or downwards as a result of random shocks. In this case, the series is called a time series with a *stochastic trend.*

However, a trend in a nonstationary time series can occur if its mean is a specific function (linear, for example) of time. Suppose that a time series can be described as $y_t = \mu_t + \varepsilon_t$ where $\mu_t = \beta_0 + \beta_1 t$, so that $y_t = \beta_0 + \beta_1 t + \varepsilon_t$. In this case the time series is said to have a *deterministic trend.* A mixed stochastic-deterministic trend process is also possible and can be described as $y_t = \beta_0 + \beta_1 t + y_{t-1} + \varepsilon_t$.[3]

The processes discussed above can be viewed as special cases of a broader class of nonstationary processes which can be described as

$$y_t = \theta y_{t-1} + \varepsilon_t \qquad \text{(iii)}$$

where θ=1. Therefore, the processes we have been discussing are called *unit root stochastic processes.*

4.2.2. COINTEGRATION AND ORDER OF INTEGRATION OF TIME SERIES VARIABLES

Nonstationarity of time series has always been regarded as a problem in econometric studies since it can give rise to spurious relationships among the levels of economic variables. It has been shown by, for example, Granger and Newbold (1974), Dickey and Fuller (1979) and, more recently, Phillips (1987), that the statistical properties of regressions using nonstationary time series are dubious. Yet, an implicit assumption in much of the literature concerning regression analysis of time series is that such data are stationary. If this is not the case, the statistical inference and forecasting procedures may not be valid. In fact, regressing one nonstationary time series on another may present apparently satisfactory estimation results even when the regressions are meaningless. An illustrative example would be the rather pointless regression of a linear trend on a quadratic trend. Another such example would be the case where the variables are subject to a stochastic rather than a deterministic trend. These regressions often provide apparently good estimation results, such as high R^2 and 'significant' t-statistics, although they

[3] It is assumed that the stochastic process ε_t is a white noise (zero mean, constant variance and no autocorrelation). However, these conditions may be relaxed to permit, for example, serial correlation. If the errors are autocorrelated, the processes (i) and (ii) can no longer be called random walks but y_t is still nonstationary.

represent spurious relationships. This shows the 'danger' of interpreting regression results involving deterministic or stochastic trended variables: the apparently robust but invalid estimation results can make it difficult to determine whether an economic relationship suggested by the theory has, in fact, any support from the data. As Charemza and Deadman (1997) point out, regression analysis makes sense only for data not subject to a trend, that is, only if the variables involved are either individually stationary or a linear combination of them is stationary. Since most economic data series contain trends (are nonstationary), these have to be purged before any sensible regression analysis can be performed.

A convenient way of purging a trend from an economic time series is by differentiation. In other words, the successive differentiation of a nonstationary time series will produce, sooner or later, a stationary series.[4] The number of times that a time series needs to be differentiated in order to achieve stationarity gives the order of integration of that series. Hence, if a nonstationary time series y_t needs to be differentiated d times before it achieves stationarity is called an integrated process of order d and denoted by $y_t \sim I(d)$. For example, the first difference of a random walk, with or without a drift, is a stationary series. Therefore, random walks are integrated series of order one, or I(1), and their first differences are integrated series of order zero, or I(0). Still, it is not necessary for a series to be a random walk for achieving stationarity by differentiation. Stationarity in a time series can be achieved by differentiation, even if its errors are autocorrelated. In this case, however, the nonstationary series is not a random walk and its errors are not white noise. A white noise series such as ε_t is a stationary process or an integrated series of order zero, that is, a I(0) variable.

From the preceding discussion, it seems that nonstationary time series variables can be a major problem for applied economics. Nonstationary or trended variables, either stochastic or deterministic, may give rise to spurious regressions, invalid inference and forecasting procedures and, generally, make regression results difficult to interpret. Unfortunately in economics, most time series are subject to some kind of trend, that is, are nonstationary, and this problem must be addressed if meaningful relationships are to be obtained with econometric regressions.

[4] Provided, of course, that the nonstationary series can be transformed into a stationary series through differentiation. It may be that the series is not integrated at all, so that no matter how many times it is differentiated it will never be transformed into a stationary series.

The remedy of differentiating nonstationary series until stationarity is attained, may not be an ideal solution as this procedure usually leads to the loss of the model's long-run properties. Indeed, regressing first difference forms of I(1) variables instead of their levels, may imply losing valuable information on the variables' long-run relationships which is given by their levels and not by their first differences. Models in first differences generally reflect short-run behaviour, whereas what is needed are models which can reproduce both the structural relationship and the short-run dynamics underlying the adjustment process to equilibrium.

Most of the theories underlying the relationships between economic variables are established considering their levels and not their differences. Those are static, steady-state or long-run equilibrium theories which assume that the variables' levels adjust fully to their long-run equilibrium in the current period. However, if this assumption does not hold, the econometric specification of the variables' relationship should reflect the behaviour of the adjustment process. The importance of such matters for statistical inference, economic analysis and policy evaluation, implied the reconsideration of the modelling problem using variables in levels to obtain short- and long-run information.

Cointegration theory states that if there is a stable long-run relationship among the levels of economic variables, they cannot diverge much from each other over time, implying that the variables are cointegrated. Cointegration means that one or more linear combinations of these variables is stationary, although individually they are not. In other words, if two or more nonstationary series can be linearly combined into a single time series which is itself stationary, the original variables are said to be cointegrated.[5] If these variables are cointegrated they cannot move far apart from each other and from the "attractor" which is their long-run equilibrium relationship. In contrast, if they are not cointegrated it is possible (but not necessarily so) that such variables have no long-run relationship and can drift arbitrarily apart from each other. "*The power of economic equilibrium as an attractor should force different variables to move together in the long-run even if not in the short-run and even if they are individually nonstationary*" (Engle and Granger 1991, p. 8). Therefore, "synchrony" of movements between time series variables is the intuitive idea underlying their cointegrated long-run relationship.

[5] The stationary series resulting from such linear combination of non-stationary variables may be the residual series of an estimated regression. A regression in which residuals are stationary or I(0) is a *cointegrating regression* and the vector of parameters linking the variables within this regression is known as the *cointegrating vector*.

Cointegration analysis is concerned with the development of ideas, concepts and methods for the investigation of meaningful long-run relationships among economic time series. The problems involved in the detection, estimation and testing of cointegrating regressions are complex and have been the focus of much recent research work. Fortunately, there are some generally accepted simple rules that can guide us through the process of modelling, estimating and testing cointegrated relationships.

Let us assume that economic theory postulates an equilibrium relationship between a pair of nonstationary I(1) series (y_t, x_t) such that $y_t = \beta x_t$. If y_t follows an equilibrium path at any instant in time, then we can rewrite the equilibrium relationship as $y = \beta x$ or $y - \beta x = 0$. The line $y = \beta x$ corresponds to an 'attractor' in Engle and Granger's (1991) nomenclature. However, the attractor equation is not expected to hold at all instances as disturbances can drive the variables' levels away from the equilibrium path. Therefore, out of equilibrium, their relationship is better described by $y_t - \beta x_t = \varepsilon_t$, where ε_t represents the extent to which the relationship is out of equilibrium and may be called an "*equilibrium error*". As ε_t is a I(0) variable with zero mean "*there will be a tendency for the points (y_t, x_t) to be around the line, and thus for the line to act as an attractor*" (Engel and Granger, op. cit.). Since the linear combination of nonstationary variables gives rise to a single stationary series ε_t ~I(0), the variables are cointegrated. Therefore, "*cointegration is a sufficient condition for the existence of an attractor*" (op. cit.). The reverse is also true, as the existence of a long-run relationship between a set of variables implies that they are cointegrated. Furthermore, there is a correspondence between cointegrated systems and the error correction mechanism (ECM). An ECM constitutes one case of a systematic adjustment process through which cointegrated variables are prevented from drifting apart from the 'attractor' line. Therefore, cointegrated variables can always be viewed as being generated by error-correction equations. The converse is also true.

From the previous discussion we conclude that any meaningful econometric analysis of an equilibrium relationship between economic time series levels must include tests for cointegration of the variables involved. Cointegration among a set of economic variables depends, among other things, on the order of integration of those variables. Although there is a similarity between cointegration and order of integration tests (commonly denoted 'unit root' tests), the latter are performed on univariate time series while the former deal with the relationships among a group of variables where (unconditionally) each has a unit root. Consequently, before any sensible regression analysis can be performed, it is essential to identify the order of integration of each of the relevant variables. In the next section, all the relevant time series variables used in the modelling of the UK tourism demand for France, Spain and Portugal are examined and a set of tests to determine their order of integration is performed.

4.2.3. ORDER OF INTEGRATION OF VARIABLES IN UK TOURISM DEMAND MODELS.

The variables included in the econometric models of this chapter are the same as those used in the previous chapter. Hence, we begin this section by recalling the definition of the variables described in chapter 3. Next, we present plots of the variables' levels and first differences, for an easier inspection of their trend features. Finally, we test for their order of integration using the Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) unit root tests.

The UK demand for tourism in each destination i, [i=F (France), S (Spain) and P (Portugal)], is measured by the logarithm of the per capita UK tourism expenditure allocated to destination i, deflated by its consumer price index (CPI_i). YF, YS and YP denote the UK tourism demand for France, Spain and Portugal, respectively. Tourism prices in destination i and competitive destinations j are measured by the logarithm of the ratio of consumer price indexes in destination i (j) and in the UK, adjusted by the relevant exchange rate. Tourism effective prices in France, Spain and Portugal are denoted by, respectively, PF, PS and PP. The UK real per capita income is measured by the logarithm of the per capita disposable income, deflated by the UK consumer price index and denoted by I.

Figures 4.1 to 4.10 present a set of graphs showing how the variables' levels and their first differences evolved over the sample period. Figures 4.1 to 4.3 relate to the levels of the variables, and Figures 4.4 to 4.10 relate to their first differences. Figure 4.1 presents the UK real per capita expenditure in France, Spain and Portugal. Figure 4.2 presents tourism effective prices in these three destinations and Figure 4.3 the UK per capita real income. Figures 4.4 to 4.10 present the individual plots of all the first differenced variables.

Figure 4.1: UK per capita real expenditure in France, Spain and Portugal

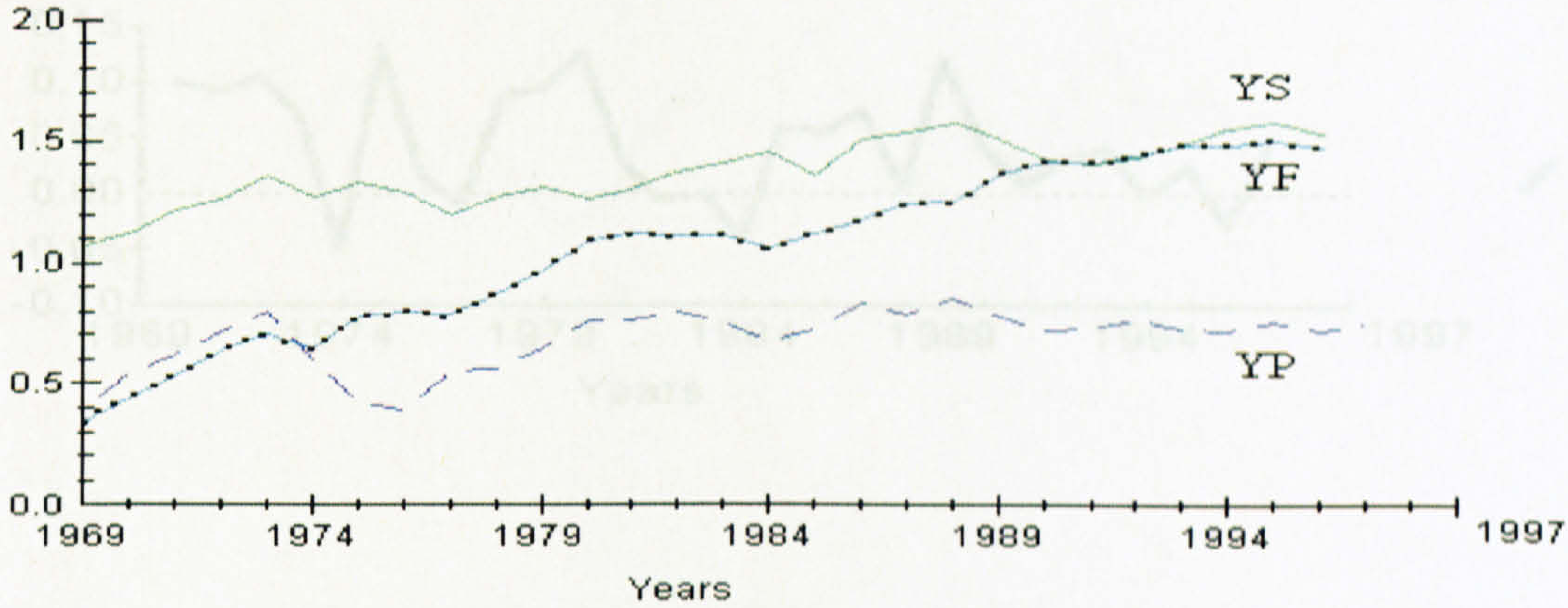


Figure 4.2: Effective prices of tourism in France, Spain and Portugal

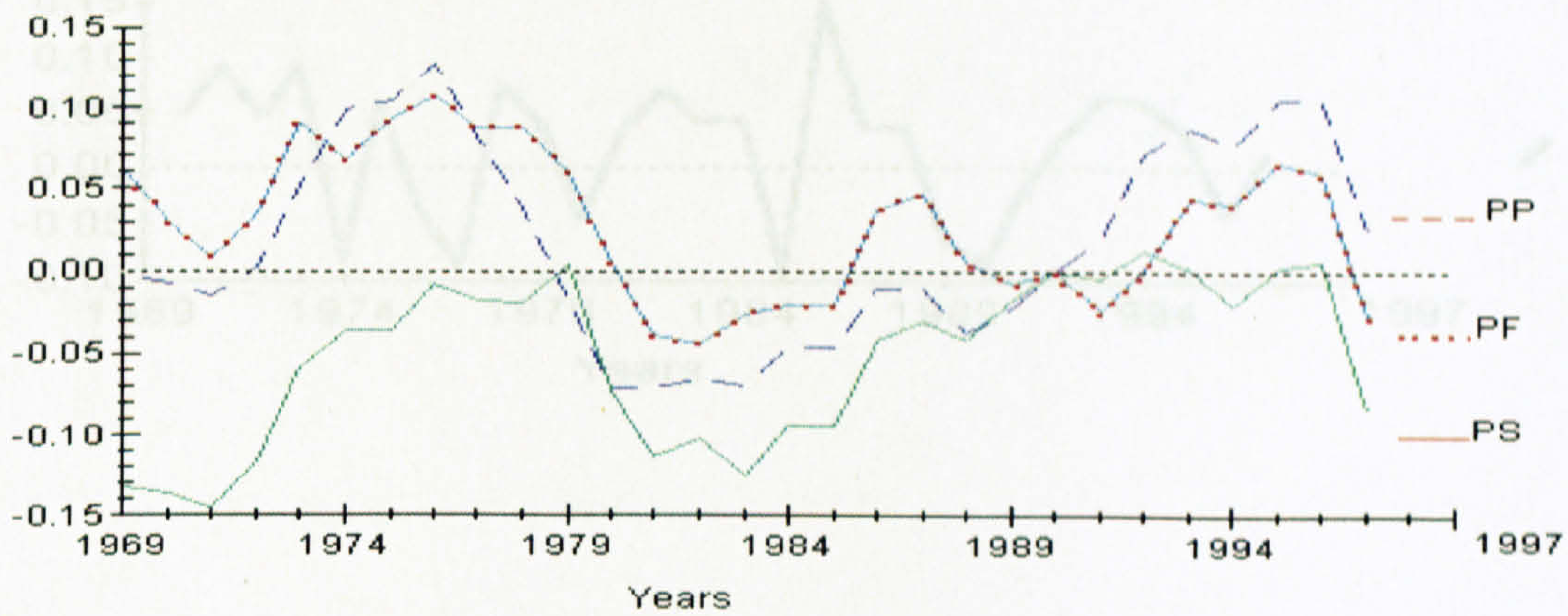


Figure 4.3: UK real per capita Income

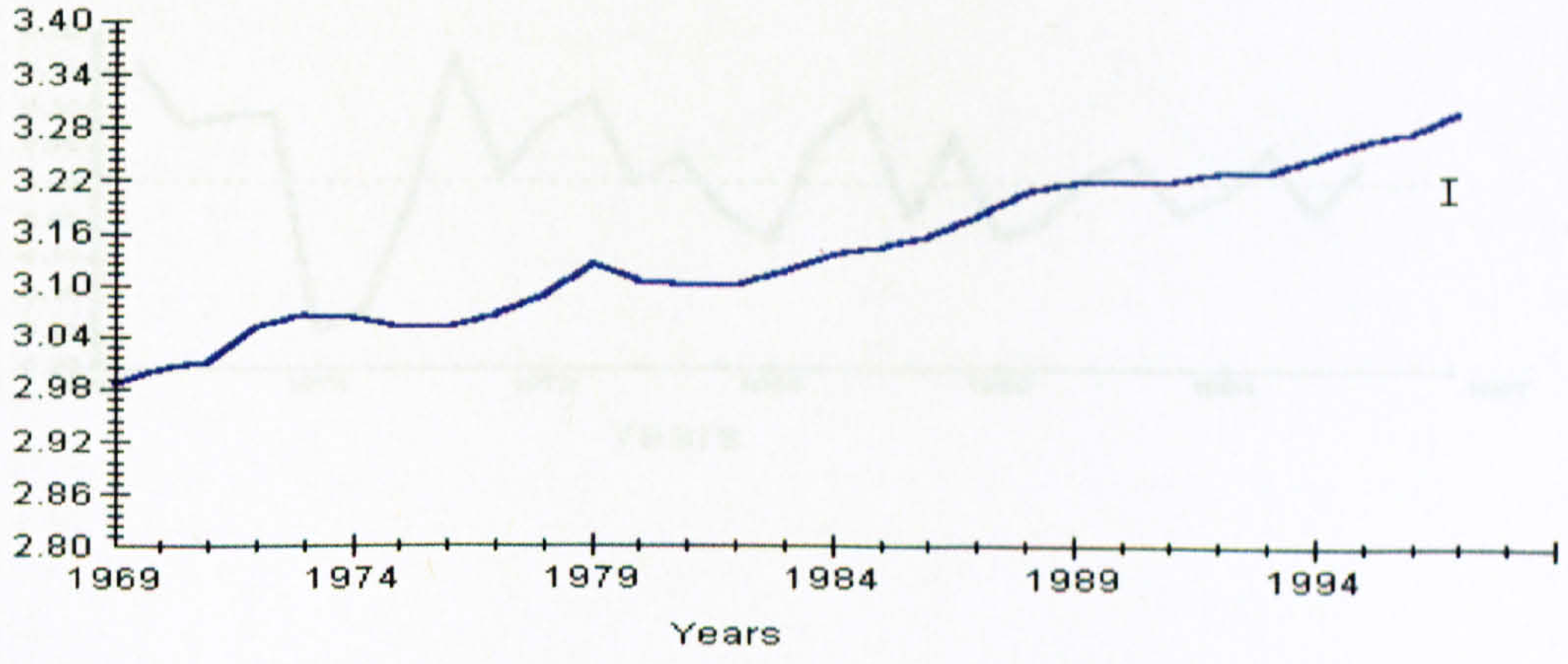


Figure 4.4: First difference of the UK real expenditure in France

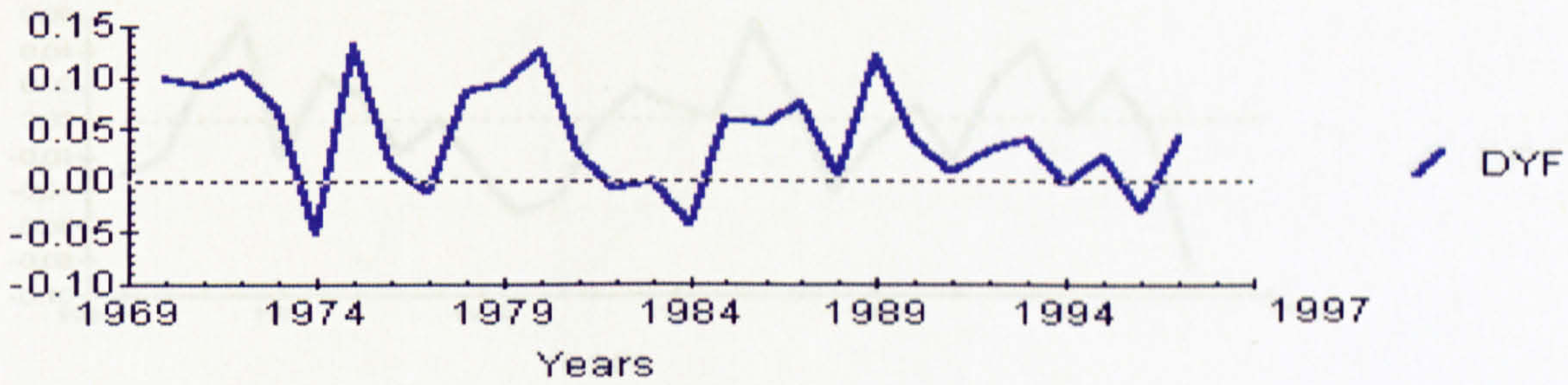


Figure 4.5: First difference of the UK real expenditure in Spain

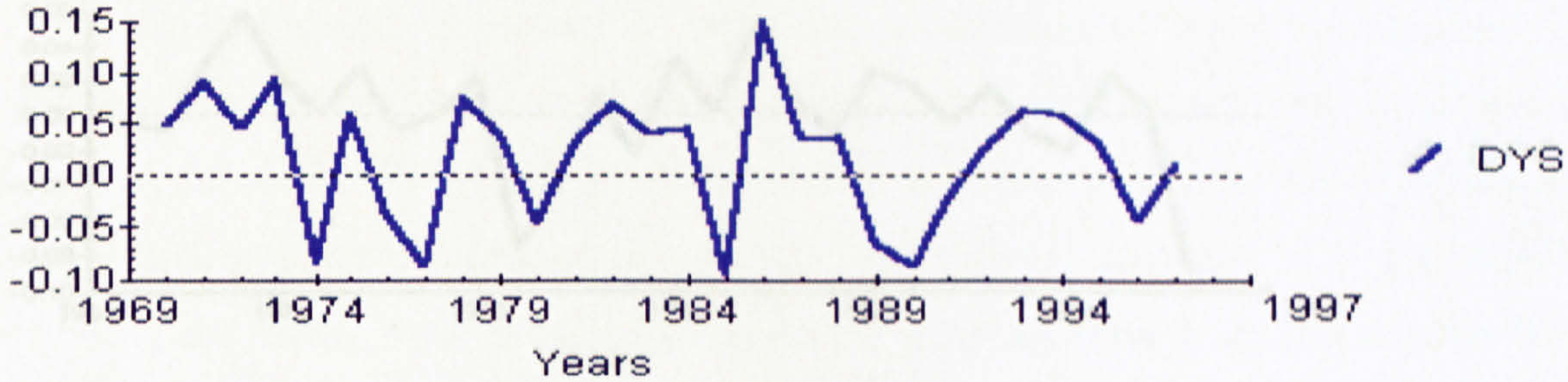


Figure 4.6: First Difference of the UK real expenditure in Portugal

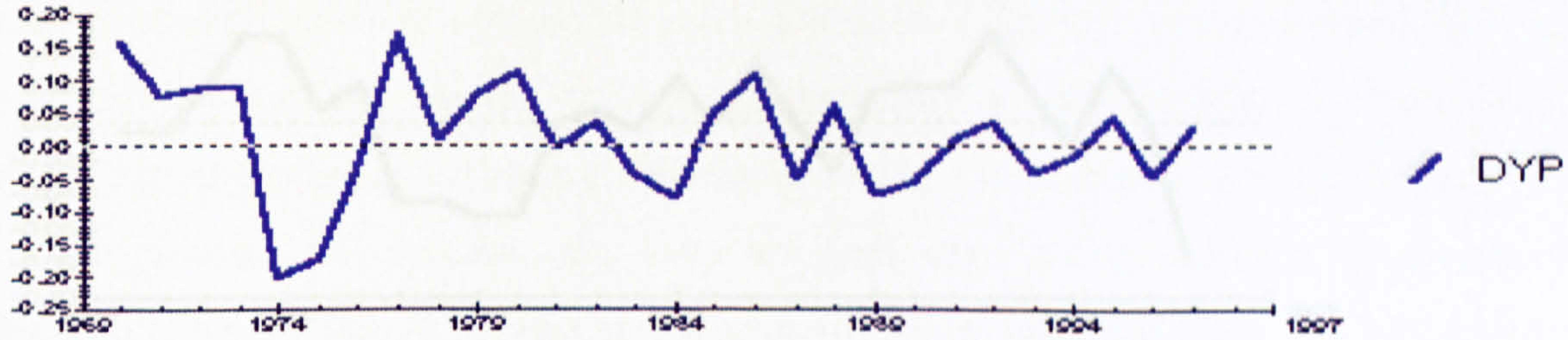


Figure 4.7: First differences of the effective price of tourism in France

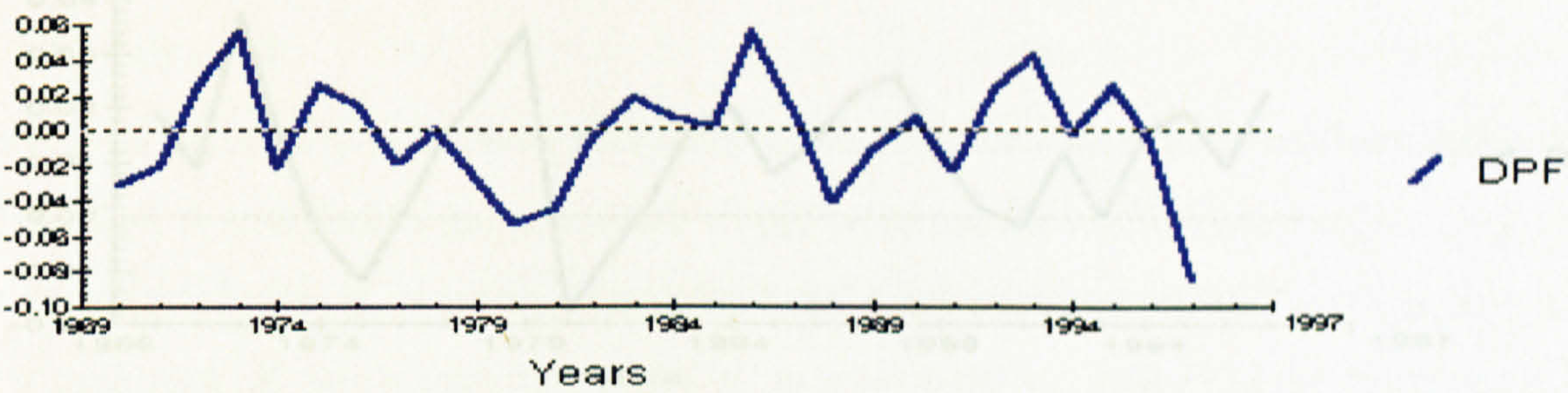


Figure 4.8: First differences of the effective price of tourism in Spain

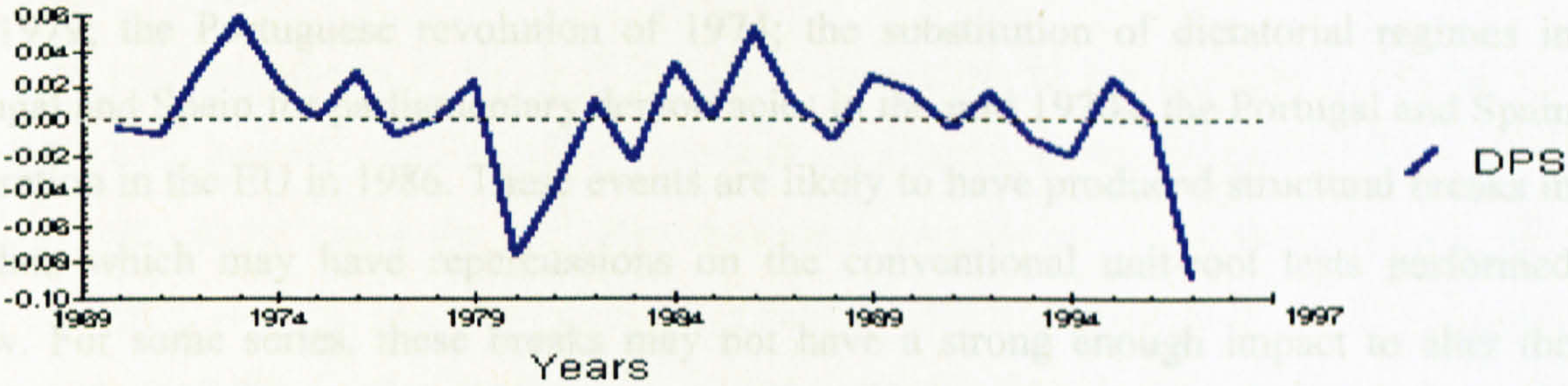


Figure 4.9: First differences of the effective price of tourism in Portugal

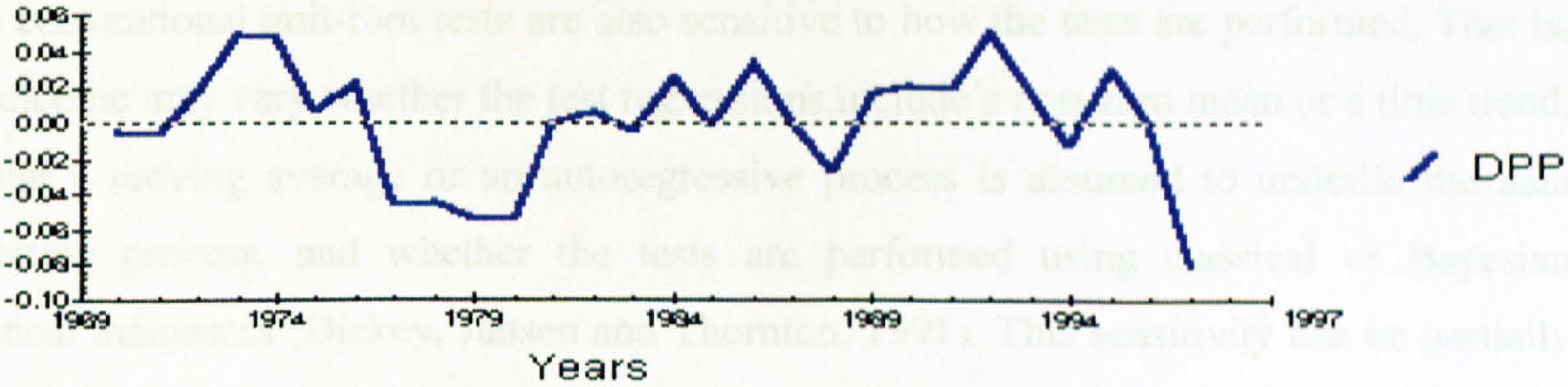
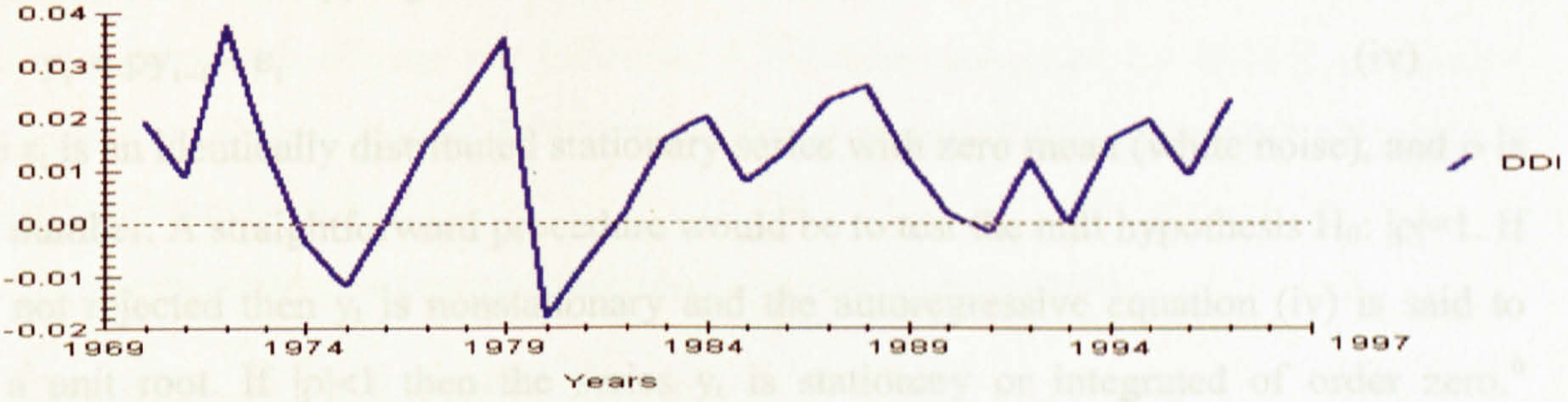


Figure 4.10: First differences of the UK real per capita income



Some of the variables' features can readily be spotted from the graphs above. For instance, the plots show that both the level and the first difference variables behave peculiarly in a sub-period of the sample that can roughly be placed at 1973-1989. This peculiar behaviour is more obvious for the variables related to Portugal and Spain. Several events, which can justify this behaviour, took place within this period: the oil crises of 1973 and 1979; the Portuguese revolution of 1974; the substitution of dictatorial regimes in Portugal and Spain for parliamentary democracies in the mid 1970s; the Portugal and Spain integration in the EU in 1986. These events are likely to have produced structural breaks in the data which may have repercussions on the conventional unit-root tests performed below. For some series, these breaks may not have a strong enough impact to alter the conclusions of the tests. However, for the tourism price series of Portugal and Spain, a clear conclusion is not to be expected and different tests may have to be performed to establish their order of integration. Apart from this potential problem, a fairly 'normal' behaviour is observed for all series, as the oscillating movements of their first differences seem to indicate stationarity and, hence, the presence of a unit root in the variables' levels.

Besides the potential problem of existing structural breaks, it should be noted that some conventional unit-root tests are also sensitive to how the tests are performed. That is, the outcome may vary whether the test regressions include a non-zero mean or a time trend, whether a moving average or an autoregressive process is assumed to underlie the data generating process, and whether the tests are performed using classical or Bayesian statistical inferences (Dickey, Jansen and Thornton, 1991). This sensitivity can be partially justified by the lack of power of these tests against an alternative hypothesis of a stationary but larger than unit root. Consequently, the DF and ADF tests are not necessarily the best way to search for stationarity in some time series and whenever these tests present dubious results, alternative approaches may have to be implemented.

Suppose we want to test the hypothesis that a time series variable y_t is integrated of order one. Assume that y_t is generated by

$$y_t = \rho y_{t-1} + \varepsilon_t \qquad \text{(iv)}$$

where ε_t is an identically distributed stationary series with zero mean (white noise), and ρ is a real number. A straightforward procedure would be to test the null hypothesis H_0: $|\rho|=1$. If H_0 is not rejected then y_t is nonstationary and the autoregressive equation (iv) is said to have a unit root. If $|\rho|<1$ then the series y_t is stationary or integrated of order zero.[6] However, as Newbold and Davis (1978) show, tests based on OLS estimations of nonstationary series might result in spurious significance. Therefore, unit root tests should be performed using a stationary series as dependent variable. An appropriate method for testing the order of integration of y_t is the Dickey-Fuller test (DF) proposed in Dickey and Fuller (1979) which is based on the estimation of an equivalent regression to (iv) such that[7]

$$\Delta y_t = \delta y_{t-1} + \varepsilon_t \qquad \text{(v)}$$

where the null hypothesis is H_0: $\delta=0$ (which is equivalent to the null $|\rho|=1$) against the alternative $\delta<0$. Rejection of the null in favour of the alternative implies that y_t is integrated of order zero. Since under the null, the dependent variable in (v) is stationary [is I(0), implying that y_t is I(1)], equation (v) is a regression of an I(0) variable on an I(1) variable and in such cases the usual critical values for the t-statistic are not valid. In these cases, the t statistic is known as the τ statistic for which the critical values can be found in, for example, Fuller (1976), MacKinnon (1991) and Charemza and Deadman (1997). In Charemza and Deadman's (1997) tables, a critical band is provided for a given number of observations and level of significance. If the t statistic for the δ coefficient in regression (v) is smaller (more negative) than the lower value of the critical band, the null hypothesis of a unit root has to be rejected in favour of the alternative of stationarity for y_t. If the computed t statistic is greater than the upper value of the critical band, the null hypothesis cannot be rejected. If the computed statistic falls within the critical band the test is inconclusive.

[6] It is conventionally assumed that "explosive" processes implying $\rho > 1$ are implausible in economics and, in practice, it is unusual for economic series to be integrated of an order greater than two.

[7] The DF test can also be used for testing the order of integration of a variable generated as a stochastic process with drift. Additionally, the Dickey-Fuller equation (v) can be modified to accommodate tests for the simultaneous presence of stochastic and deterministic trends.

A weakness of the DF test is that it does not take into consideration the possibility of serial correlation of the disturbances ε_t. If the error terms are not a white noise, the OLS estimates of equation (v) are not efficient. A solution suggested by Dickey and Fuller (1981) consists of using lagged dependent variables as additional explanatory variables to 'whiten' the error terms.[8]

In the presence of autocorrelation the regression to be tested is

$$\Delta y_t = \delta y_{t-1} + \sum_{i=1}^{m} \beta_i \Delta y_{t-i} + \varepsilon_t \qquad \text{(vi)}$$

The test is now called Augmented Dickey-Fuller test and denoted by ADF(p); p is the number of lags of the dependent variable in the right-hand side of equation (vi). The procedure to implement the test is the same as before, with the same null and alternative hypothesis and the same decision criteria The ADF test is generally regarded as the most efficient among the simpler integration tests. Yet, it is not very powerful "*in finite samples for alternatives H_1: $\delta=\delta^*<1$, when δ^* is near unity. There is a size-power trade-off depending on the order of augmentation used to deal with the residual serial correlation.* (Pesaran and Pesaran, 1997, p.217). ADF tests with a large p relative to the sample size have very low power.

The order of the ADF regression is selected by using the two step procedure for model selection. First, the augmentation order is selected, based on such criteria as the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC). Then, using the selected lag length, the ADF test is performed.

The results obtained for the ADF(p) tests are presented in table 4.1. The variables' levels and their first differences (indicated by the operator Δ) are written in the first column; the following columns present the t-statistic for the ADF(0) and ADF(1) regressions with intercept and no deterministic trend for all variables, the AIC and SBC criteria for lag selection and the critical values for the τ statistic at the 5% significance level. These critical values are computed using the response surface estimates in MacKinnon (1991, Table 1). The MacKinnon critical value at the 5% level for the ADF(0) test (27 observations) is –2.975, and for the ADF(1) test (26 observations) is –2.980. As these critical values do not

[8] The length of the lagged variables to be included (the order of augmentation) is the minimum which secures a with noise property for the error terms.

differ much we use –2.980 as the unique MacKinnon critical value for both tests. The bold values indicate the valid statistic for comparison with the critical value, according to the selection criteria.

Table 4.1: ADF(0) and ADF(1) unit root tests for the variables in levels YF, YS, YP, PF, PS, PF, I and their first differences ΔYF, ΔYS, ΔYP, ΔPF ΔPS, ΔPP, ΔI.

Variables	Test	t-statistic	AIC criterion	SBC criterion	5% critical value
YF	ADF(0)	**-1.9706**	42.63*	41.33*	-2.98
	ADF(1)	-1.9518	41.70	39.75	
ΔYF	ADF(0)	**-4.9874**	39.12*	37.86*	
	ADF(1)	-3.4481	38.13	36.24	
YS	ADF(0)	**-1.9431**	36.21*	34.91*	-2.98
	ADF(1)	-1.8724	35.22	33.28	
ΔYS	ADF(0)	**-5.5582**	33.33*	32.07*	
	ADF(1)	-3.8840	32.38	30.49	
YP	ADF(0)	-2.1911	29.19	27.89*	-2.98
	ADF(1)	**-2.6103**	29.82*	27.88	
ΔYP	ADF(0)	**-4.1809**	26.00*	24.74*	
	ADF(1)	-3.9364	25.54	23.66	
PF	ADF(0)	**-1.8686**	56.53*	54.99*	-2.98
	ADF(1)	-2.3113	54.76	52.82	
ΔPF	ADF(0)	**-3.4354**	53.05*	51.75*	
	ADF(1)	-3.4165	50.38	48.49	
PS	ADF(0)	**-2.0832**	56.85*	55.52*	-2.98
	ADF(1)	-2.4279	54.26	52.32	
ΔPS	ADF(0)	**-3.7881**	52.29*	51.00*	
	ADF(1)	-2.9683	48.92	47.04	
PP	ADF(0)	-1.4130	55.12	53.79	-2.98
	ADF(1)	**-2.5996**	56.67*	54.73*	
ΔPP	ADF(0)	**-2.4806**	54.34*	53.04*	
	ADF(1)	-2.2523	50.85	48.96	
I	ADF(0)	-0.1979	79.88*	78.55*	-2.98
	ADF(1)	-0.0670	76.37	74.43	
ΔI	ADF(0)	**-3.9281**	80.70*	79.35*	
	ADF(1)	-4.0379	77.50	75.56	

The tests clearly indicate all variables in levels as nonstationary and, except for ΔPP, all variables in first differences as stationary. This implies that all level variables, except for the price of Portugal (PP), can be considered as integrated of order one or I(1) according to the ADF test. Indeed, given the ADF test results of nonstationarity for both the level variable PP and its first difference (ΔPP), we cannot conclude that PP is an I(1) variable under the rules of this particular unit root test.

These results were expected, given the economic and political events of the mid 1970's which are likely to have affected the structure of all time series in general, and of the series 'price of Portugal' in particular. These events affected Portugal in a particular way due to the political revolution in this country in 1974. The effects of this revolution are believed to have lasted until, at least, 1979. Indeed, during the period 1974-1979, the Portuguese currency devalued sharply, foreign and domestic investment attained their lowest values ever, inflation rose to its highest values ever, and general political and social instability adversely affected all economic sectors in the country. Hence, it is quite possible that a structural break exists in the series PP.

The weakness of the ADF tests related to the presence of structural break in a series is addressed in Perron (1989, 1990). Perron (1989) suggests tests to examine stationarity in the presence of various types of shocks and shows, among other things, that the effect of a "pulse" variable (a dummy taking the value of unity when the shock is 'active' and zero otherwise) on an I(1) variable is permanent while on an I(0) variable is not.

We believe that we can safely assume variable PP to be I(1) and its first difference to be I(0), given the lack of power of the ADF test and given that the variable for the price of tourism in Portugal would be accepted as I(1) with MacKinnon critical values at the 10% level, and is accepted as I(1) with the Charemza and Deadman (1997) critical values at the 5% level.[9] Moreover, the graph of the variable's first difference, except for the period 1974-1979, shows the typical oscillatory movement of a stationary series, Consequently, we conclude that all relevant time series included in the regressions of the UK demand for tourism in France, Spain and Portugal are integrated of order one. The next step is to

[9] In Chapter 7, we use the Phillips-Perron (1988) test to investigate the stationarity of time series PP. Based on this test results, we conclude that the variable in first differences is stationary, and the variable in levels is non-stationary. Therefore, according to the Phillips-Perron test, variable PP is integrated of order one or I(1).

investigate whether these variables are cointegrated in the empirical models built using the 'general to specific' modelling approach. The following section addresses this issue.

4.3. COINTEGRATION ANALYSIS AND ERROR-CORRECTION MECHANISM IN THE UK TOURISM DEMAND FOR FRANCE, SPAIN AND PORTUGAL

The need to build, estimate, test and evaluate models which combine both long-run and short-run properties and which, at the same time, are stationary in the linear combination of the variables they assemble, has led to the reconsideration of regression analysis involving variables in levels. Researchers' attention has recently been focusing on the modelling of economic time series which, although individually nonstationary, can be linearly combined into a single series which is itself stationary. Series with such propriety are known as cointegrated series. Hence, cointegration analysis is concerned with the investigation of long-run relationships among a set of variables in levels, each of which (unconditionally) has a unit root.

In section 4.2.2 we established the link between cointegration and meaningful long-run relationships among variables in levels, using as an example a pair of nonstationary I(1) variables, y_t and x_t,. The "attractor" designated y-βx=0, illustrates the equilibrium path that forces the variables to move together in the long-run, even if not in the short-run or even if they are not individually stationary. If such an "attractor" exists then, in the out of equilibrium relationship $y_t-\beta x_t=\varepsilon_t$, the expected value of the residuals ε_t should be zero. Furthermore, the residuals' series should be stationary, so that the variables' levels may deviate from their equilibrium values in the short-run, but should converge to it in the long-run. Failure to find a stationary relationship among the variables (meaning that they are not cointegrated) implies that the long-run relationship does not exist in any meaningful sense. Summarising, cointegration is a sufficient condition for the existence of a long-run relationship among variables in levels and the existence of a long-run relationship among a set of variables implies that they are cointegrated.

Most economic theories stipulate long-run relationships among a group of variables and assume the existence of a cointegrating vector combining the group of series into a univariate series ε_t. Given this assumption, a test for cointegration could be performed by applying any conventional unit root test to the series of residuals of the long-run regression. However, this procedure requires that the hypothesised cointegrating vector is known.

Since in most cases the cointegrating vector is unknown, some linear combination of the relevant variables must first be assumed to be stationary and the unknown cointegrating vector must then be estimated. Consequently, we first address the building of empirical models describing the relationship between the UK demand for tourism in France, Spain and Portugal with its assumed determinants. Then we estimate the underlying cointegrating vectors. Finally we test for stationarity of the resulting series of residuals.

Demand theory establishes the existence of a steady state relationship between the quantity demanded of a commodity and its own-price, consumers' disposable income and prices of the commodity and its substitute and/or complement commodities. Theoretical models are generally established within a static framework, which assumes that the adjustment process of the economic variables to their long-run equilibrium values occurs instantaneously. Hence, in general, theory does not provide information on the dynamics of the short-run adjustment process. However, it has been recognised that empirical models which can provide information on both the long- and short-run aspects of economic behaviour are important.

An approach to modelling short-run dynamics was proposed in Davidson *et al.* (1978) and has been developed by a number of researchers ever since. In Davidson *et al.* (1978), the short run dynamics is modelled through an ECM which allows the dependent variable to return to its long-run equilibrium path following a disturbance shock. Cointegrated variables can always be viewed as being generated by an error-correction mechanism. Therefore, a sound model-building strategy to analyse the existence of a cointegrated relationship between the UK tourism demand and its determinants, (as well as its short-run dynamic adjustment process), would be to estimate the underlying error-correction model. This model may result from applying a 'general to specific' approach to an autoregressive distributed lag model (ARDL). Through successive hypothesis testing, the ARDL model may be reduced to an interpretable and parsimonious form which is data-coherent and statistically robust. The parsimonious ARDL model can then be used to derive the short-run error-correction mechanism embedded in its specification and to test for cointegration among the variables included.

Assume the following ARDL model of the UK tourism demand for destination i.[10]

[10] The maximum lag-length of the ARDL model is one, as suggested by the AIC and SBC criteria.

$$Yi_t = \alpha_i + a_{i0}PP_t + a_{i1}PP_{t-1} + b_{i0}PS_t + b_{i1}PS_{t-1} + c_{i0}PF_t + c_{i1}PF_{t-1} + + \beta_{i0}I_t + \beta_{i1}I_{t-1} + \delta_i V_t + \theta_i Yi_{t-1} + ui_t \tag{4.1}$$

where i = F (France), S (Spain) and P (Portugal); Yi is the logarithm of the UK per capita real tourism expenditure in destination i; PP, PS and PF are the logarithms of the effective tourism prices in, respectively, Portugal, Spain and France; I is the logarithm of the UK real per capita disposable income and V is a dummy variable representing a disturbing event which assumes the value of unity for a sequence of observations and zero otherwise.

If a steady-state equilibrium relationship exists, the variables in equation (4.1) may not move from their equilibrium values; that is, at any time, any variable x assumes its long-run equilibrium value such that, $x_{t-1} = x_t = x^*$. Therefore, equation (4.1) can be written as representing the steady-state long-run equilibrium relationship such that,

$$Yi^* = \frac{\alpha_i}{1-\theta_i} + \frac{a_{i0}+a_{i1}}{1-\theta_i}PP^* + \frac{b_{i0}+b_{i1}}{1-\theta_i}PS^* + \frac{c_{i0}+c_{i1}}{1-\theta_i}PF^* + \frac{\beta_{i0}+\beta_{i1}}{1-\theta_i}I^* + \frac{\delta_i}{1-\theta_i}V$$

or $$Yi^* - (\alpha_{i0} + \alpha_{i1}PP^* + \alpha_{i2}PS^* + \alpha_{i3}PF^* + \alpha_{i4}I^* + \alpha_{i5}V) = 0 \tag{4.2}$$

where the coefficients, $\alpha_{ij} = \frac{k_{i0}+k_{i1}}{1-\theta_i}$, for all k=a, b, c, β and j=0, ..., 4 and $\alpha_{i5} = \frac{\delta_i}{1-\theta_i}$ represent the long-run impacts of the determinants' changes on the dependent variable and $\alpha_{i0} = \frac{\alpha_i}{1-\theta_i}$ is the intercept.

Since all variables (except V) are measured in logarithms, the α coefficients may be interpreted as long-run elasticities. On the other hand, the coefficients of the variables' current values in equation 4.1 (except V) represent short-run elasticities, and $1-\theta$ represents the adjustment velocity.

If the long-run equilibrium given in (4.2) is disturbed by a shock, its equilibrium form can be restored by an error-correction mechanism which can be described as

$$Yi_t - (\alpha_{i0} + \alpha_{i1}PP_t + \alpha_{i2}PS_t + \alpha_{i3}PF_t + \alpha_{i4}I_t + \alpha_{i5}V_t) = \varepsilon i_t \tag{4.3}$$

where εi_t is the 'equilibrium error', denoted as EQE hereafter.

The ECM model representing the short-run dynamic adjustment process to the long-run equilibrium relationship, can be derived from equation (4.1) in the following way: first, subtract from both sides of equation (4.1) the lagged dependent variable Yi_{t-1} obtaining

$$\Delta Yi_t = \alpha_i + a_{i0}PP_t + a_{i1}PP_{t-1} + b_{i0}PS_t + b_{i1}PS_{t-1} + c_{i0}PF_t + c_{i1}PF_{t-1} +$$
$$+ \beta_{i0}I_t + \beta_{i1}I_{t-1} + \delta_i V_t + (\theta_i - 1)Yi_{t-1} + ui_t$$

then, sum and subtract $a_{i0}PP_{t-1}$, $b_{i0}PS_{t-1}$, $c_{i0}PF_{t-1}$, $\beta_{i0}I_{t-1}$ and $\delta_i V_{t-1}$. Rearranging and collecting terms we obtain

$$\Delta Yi_t = a_{i0}\Delta PP_t + b_{i0}\Delta PS_t + c_{i0}\Delta PF_t + \beta_{i0}\Delta I_t + \delta_i \Delta V_t -$$
$$(1-\theta_i)\left[Yi_{t-1} - \left(\frac{\alpha_i}{1-\theta_i} + \frac{a_{i0}+a_{i1}}{1-\theta_i}PP_{t-1} + \frac{b_{i0}+b_{i1}}{1-\theta_i}PS_{t-1} + \frac{c_{i0}+c_{i1}}{1-\theta_i}PF_{t-1} + \frac{\beta_{i0}+\beta_{i1}}{1-\theta_i}I_{t-1} + \frac{\delta_i}{1-\theta_i}V_{t-1}\right)\right] + ui_t$$

(4.4)

The expression in brackets in equation (4.4) represents the equilibrium error of equation (4.3). Therefore, equation (4.4) can be simplified such that,

$$\Delta Yi_t = a_{i0}\Delta PP_t + b_{i0}\Delta PS_t + c_{i0}\Delta PF_t + \beta_{i0}\Delta I_t + \delta_i \Delta V_t - (1-\theta_i)EQE + ui_t \qquad (4.5)$$

Equation (4.5) represents the error-correction model associated with the long-run equilibrium model (4.2), and both models are derived from the general dynamic ARDL model in (4.1).

In the next sections we define for each destination i, a general dynamic ARDL model and derive the associated error-correction models.

4.3.1. THE UK DEMAND FOR TOURISM IN FRANCE

The ARDL general specification of the UK tourism demand for France is assumed to be given by the following equation:

$$YF_t = \alpha_F + a_{F0}PP_t + a_{F1}PP_{t-1} + b_{F0}PS_t + b_{F1}PS_{t-1} + c_{F0}PF_t + c_{F1}PF_{t-1} + d_{F0}YP_t + d_{F1}YP_{t-1} +$$
$$+ e_{F0}YS_t + e_{F1}YS_{t-1} + \beta_{F0}I_t + \beta_{F1}I_{t-1} + \delta_{F1}F7581_t + \delta_{F2}F8789_t + \delta_{F3}F9497 + \theta_F YF_{t-1} + uF_t$$

(4.6)

Hence, we assume that the current value of the UK tourism demand for France (YF_t) depends on its own lagged value (YF_{t-1}), and on the current and lagged values of own- and competing destinations' prices (Pj_t and Pj_{t-1} for j=F, S, P), UK real per capita income (I_t and I_{t-1}) and real per capita expenditure allocated to France's neighbouring competitors (YS_t YS_{t-1} and YP_t YP_{t-1}). We also assume that the events taking place during the periods 1974-1981 (due to the oil price shocks of 1973 and 1979, Portuguese revolution in 1974, political regime changes in Spain and Portugal during the period 1975-1979 and Spain's joining the

OECD in 1980), 1987-1989 (due to Spain and Portugal's simultaneous joining the EU in 1986) and 1994-1997 (due to the opening of the Channel Tunnel), had a significant impact on the UK demand for tourism in France. Hence, three dummy variables are added in equation (4.6) taking the value of unity during these periods and zero otherwise.

Table 4.2 presents the estimation results (t-values in brackets) for the general ARDL model (denoted by 'regression 1'), and for the intermediate and final parsimonious forms (denoted by 'regression 2' and 'regression 3') which were derived from the former by the systematic testing of plausible restrictions. The last eight rows of table 4.2 present several criteria for evaluating the regressions goodness of fit (adjusted R^2, overall significance F test and the residual sum squares) and statistical information based on the Lagrange Multiplier (LM) version of several diagnostic tests: LM test for residual serial correlation; Ramsey's RESET test using the square of the fitted values for the functional form; residuals normality test based on the skewness and kurtosis of residuals, and residuals heteroscedasticity test based on the regression of squared residuals on squared fitted values. For all these tests the respective p values are shown in brackets.

The t-statistic for individual significance and the LM and F tests for additional variables joint significance, indicate the variables for tourism price in Portugal, UK expenditure in Portugal and UK expenditure in Spain as irrelevant variables. The same statistics indicate tourism prices in Spain and France, UK real income and the one-lag dependent variable as relevant explanatory variables in the parsimonious model 'regression 3'. The three dummies appearing in regressions 1 and 2 have been merged into one single dummy (F) which appears in regression 3. The construction of the 'merging' dummy F takes into consideration the relationship between the coefficients' estimates of F7581, F8789 and F9497. As indicated by regression 2 estimation results, the coefficient estimates for F8789 and F9497 are not significantly different from each other and are not significantly different from 80% of the coefficient estimate for F7581. In fact, the null hypothesis H_0: $\delta_2=\delta_3$ $\Lambda\delta_2=0.8\delta_1$ is strongly supported by the Wald test statistic which is $\chi(2)=0.006(0.997)$. Therefore, the new dummy variable F assumes the value of unity in the period 1975-81, the value 0.8 in the periods 1987-89 and 1994-97 and zero otherwise. Moreover, an F-test for additional variables was performed on the joint significance of PP_t, PP_{t-1}, PP_t, YP_{t-1}, YS_t and YS_{t-1} using regression 1 against regression 2, and the null of no significance was not rejected at the 1% level with the value of F(6, 13)=1.82.

Table 4.2: Estimation results of the ARDL model for the UK demand for tourism in France

	Regression 1	Regression 2	Regression 3
Intercept	-4.9185 (-5.03)	-4.6021 (-4.50)	-4.6108 (-7.50)
PP_t	-0.5394 (-1.78)		
PP_{t-1}	0.0455 (0.12)		
PS_t	0.0154 (0.05)	-0.1563 (-0.55)	-0.1672 (-0.79)
PS_{t-1}	0.8235 (1.71)	1.0450 (3.06)	1.0441 (4.19)
PF_t	0.8283 (2.38)	0.2145 (0.83)	0.2233 (1.10)
PF_{t-1}	-2.0355 (-6.44)	-2.127 (-6.09)	-2.1189 (-8.31)
I_t	0.3103 (0.62)	0.9654 (2.50)	0.9633 (2.80)
I_{t-1}	1.6040 (2.85)	0.7498 (1.65)	0.7539 (1.83)
YP_t	-0.0415 (-0.55)		
YP_{t-1}	-0.0444 (-0.66)		
YS_t	-0.0833 (-0.61)		
YS_{t-1}	-0.1532 (-1.67)		
F7581	0.0641 (3.35)	0.0956 (5.20)	0.0952 (6.75)
F8789	0.0921 (3.91)	0.0758 (3.89)	
F9497	0.0947 (2.91)	0.0776 (2.09)	
YF_{t-1}	0.4269 (6.08)	0.3551 (5.25)	0.3568 (5.97)
Adjusted R^2	0.9971	0.9957	0.9961
Residual Sum Square	0.0350	0.0080	0.0080
F-statistic	574.18	623.26	870.42
Serial Correlation	17.41 (0.00)	0.02 (0.90)	0.01 (0.92)
Functional Form	0.001 (0.97)	0.14 (0.71)	0.12 (0.73)
Normality	0.718 (0.70)	1.50 (0.47)	1.53 (0.47)
Heteroscedasticity	0.835 (0.36)	0.49 (0.48)	0.52 (0.47)

All statistics and diagnostic tests presented in Table 4.2 show regression 3 as a robust model with no evidence of residual serial correlation or heteroscedasticity, with high explanatory power, good fit, acceptable functional form and evidence of a normal distribution of the residuals. Therefore, regression 3 is accepted as a good model.[11]

From the regression 3 estimation results, we can now derive the long-run equilibrium relationship and compute the long-run effects of the determinants' changes on the dependent variable. The long-run estimates derived from the ARDL (1,1,1,1) model of the UK tourism demand for France are:

Intercept: $\hat{\alpha}_0 = \dfrac{\hat{\alpha}}{1-\hat{\theta}} = \dfrac{-4.6108}{1-0.35682} = -7.1688$

Price of Spain: $\hat{\alpha}_1 = \dfrac{\hat{b}_0 + \hat{b}_1}{1-\hat{\theta}} = \dfrac{-0.16721+1.0441}{1-0.35682} = 1.3634$

Price of France: $\hat{\alpha}_2 = \dfrac{\hat{c}_0 + \hat{c}_1}{1-\hat{\theta}} = \dfrac{0.22328-2.1189}{1-0.35682} = -2.9473$

UK Income: $\hat{\alpha}_3 = \dfrac{\hat{\beta}_0 + \hat{\beta}_1}{1-\hat{\theta}} = \dfrac{0.96329+0.7539}{1-0.35682} = 2.6699$

Dummy F: $\hat{\alpha}_4 = \dfrac{\hat{\delta}}{1-\hat{\theta}} = \dfrac{0.095218}{1-0.35682} = 0.14804$

Velocity of adjustment: $1-\hat{\theta} = 1-0.35682 = 0.64318$

Therefore, the estimated long-run equilibrium relationship for the UK tourism demand in France can be written as (t ratios in brackets):

$$YF_t = \underset{(-13.46)}{-7.1688} + \underset{(4.60)}{1.3634}PS_t - \underset{(-10.34)}{2.9473}PF_t + \underset{(16.25)}{2.6699}I_t + \underset{(7.08)}{0.14804}F_t + \hat{u}_t \tag{4.7}$$

The residuals of regression (4.7) constitute the EQE series which is used to construct the ECM associated with the long-run relationship derived from the ARDL model. The EQE series is given by

$$EQE_t = YF_t - \left(-7.1688 + 1.3634PS_t - 2.9473PF_t + 2.6699I_t + 0.14804F_t\right) \tag{4.8}$$

and the ECM for the UK demand for tourism in France is

$$\Delta YF_t = b_0\Delta PS_t + c_0\Delta PF_t + \beta_0\Delta I_t + d_0\Delta PS_t + \delta\Delta F_t - (1-\theta)EQE_{t-1} + v_t \tag{4.9}$$

[11]We address the issue of coefficients' stability in section 4.4. The models' forecasting ability is examined in chapter 8.

The estimation results of the error-correction representation for the selected ARDL model, are the following (t-ratios in brackets):

$$\Delta\hat{Y}F_t = \underset{(-0.79)}{-0.1672}\Delta PS_t + \underset{(1.10)}{0.2233}\Delta PF_t + \underset{(2.80)}{0.9633}\Delta I_t + \underset{(6.75)}{0.09522}\Delta F_t - \underset{(-10.75)}{0.64318}EQE_{t-1}$$

Adjusted R^2= 0.839 RSS=0.008 F(5, 22)=29.79 DW=1.95

It should be noted that the coefficients and respective t-ratios for all variables (except the error correction term) are the same as the coefficients and corresponding t-ratios of the current value variables (the short-run impacts) obtained from regression 3 estimation. As the EQE coefficient is (1-0), although its standard deviation is the same (0.059809) as the coefficient of the lagged dependent variable in regression 3, its t-value is now given by -(1-0.35682)/0.059809=–10.75.

The next step is to apply the conventional unit root tests to the EQE series. If the series is stationary then the long-run equilibrium relationship expressed in equation (4.7) is meaningful and the variables involved are cointegrated. The results of unit root tests for EQE using an intercept and no trend are as follow:[12]

Test	EQE series	MacKinnon critical values[13]	
		5%	10%
ADF(0) = DF	-3.1898	-4.1315	-3.7286
ADF(1)	-3.0332		
ADF(2)	-2.5659		
ADF(3)	-3.5544		
ADF(4)	-3.6796		
ADF(5)	-4.2366		

The results indicate cointegration at the 5% level for a lag-length of 5 in the test regression. Yet, it should be noted that in some situations, the conventional unit root tests may not be powerful enough to enable a clear conclusion about the stationarity of the series under analysis. Hence, in the cointegration analysis of the relationship between the UK

[12] The AIC and SBC criteria indicate five lags as the appropriate lag-length to eliminate autocorrelation in the EQE series. Therefore, the number of observations available is now 23.
[13] The MacKinnon critical values are computed for 23 observations and three I(1) regressors for which the null of non-cointegration is being tested. The lag-length dimension is likely to weaken the power of the test.

tourism demand and its determinants we apply, alongside the conventional unit root tests, the estimation and testing procedure advanced in Pesaran and Shin (1995) and Pesaran *et al.* (1996), denoted as the PP procedure hereafter. In Pesaran and Pesaran (1997) it is claimed that "*the main advantage of this estimation and testing strategy lies in the fact that it can be applied irrespective of whether the regressors are* I(0) *or* I(1)" (p.308).

The PP procedure involves two stages. In the first stage, the existence of a long-run relationship between the variables is tested by computing the F-statistic for the joint significance test of the lagged variables in the error correction model. The Pesaran and Pesaran (1997) error correction version of the ARDL model corresponding to the UK demand for tourism in France is:

$$\begin{aligned}\Delta YF_t &= \gamma_0 + \gamma_1 \Delta PS_t + \gamma_2 \Delta PF_t + \gamma_3 \Delta I_t + \gamma_4 \Delta F_t \\ &\quad + \lambda_1 PS_{t-1} + \lambda_2 PF_{t-1} + \lambda_3 I_{t-1} + \lambda_4 F_{t-1} + \lambda_5 YF_{t-1} + w_t\end{aligned} \tag{4.10}$$

and the hypothesis to be tested is whether a long-run relationship exists between the variables YF, PS, PF, I and F. Hence, the null of non-existence of a long-run relationship is defined by H_0: $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = 0$ against the alternative H_1: $\lambda_i \neq 0$, for any i =1, ...,5. The relevant statistic for this test is the F-statistic. However, the (asymptotic) distribution of this F-statistic is non-standard, irrespective of the integration order of the regressors. The appropriate critical values are included in Pesaran *et al.*'s (1996) tables.[14] If the computed F-statistic exceeds the upper value of the critical band then we can reject the null of non-existence of a long-run relationship between the variables.[15]

Another feature of this procedure is the possibility of establishing whether the regressors are "*long-run forcing*" variables.[16] This procedure is as follows. Once the existence of a long-run relationship between the variables has been (statistically) recognized by the rejection of the null, we repeat the test in exactly the same terms, using as the dependent variable the explanatory variables in first difference, one by one. In the case of the UK demand for France, we test the same null hypothesis considering, successively,

[14] The tables provide a 'band' covering all possible classifications of the regressors into I(0), I(1) or even fractionally integration orders. If the computed F-statistic falls out of the band, a conclusive decision (either rejecting or not rejecting the null) can be made. If the computed F-statistic falls within the band, the result is inconclusive.

[15] Since the EQE term can be viewed as expressing the relationship between the variables' lagged values in equation (4.8), the test proposed by Pesaran *et al* (1997) may be seen as equivalent to a test for individual significance of the EQE variable itself, in regression (4.9). This procedure for testing for cointegration is referred in Benerjee *et al.* (1998).

[16] In Pesaran and Pesaran (1997), the term 'long-run forcing' variables is used interchangeably with the concept of exogeneity. Therefore, a long-run forcing variable can be interpreted as an exogenous variable.

the variable ΔPS, ΔPF, ΔI and ΔYS as the dependent variable, while ΔYF now appears as one of the regressors. If, for all the additional regressions tested, the null of non-existence of a long-run relationship cannot be rejected, we may conclude that there is only one long-run relationship linking the variables which indicates ΔYF as the 'true' dependent variable, and all the other first differenced variables as explanatory variables.[17] The results obtained for this set of tests within the second stage of the PP procedure are presented in Table 4.3.

Table 4.3: Significance test for additional variables YF_{t-1}, PS_{t-1}, PF_{t-1}, I_{t-1} and F_{t-1}

Dependent variable	Explanatory variables	F-statistic (5, 18)	Critical band at the 1 % level	Test result
ΔYF	ΔPS, ΔPF, ΔI and ΔF	28.2135	(3.219; 4.378)	Rejection of the null
ΔPS	ΔYF, ΔPF, ΔI and ΔF	1.2927		Non-rejection of the null
ΔPF	ΔPS, ΔYF, ΔI and ΔF	2.3165		Non-rejection of the null
ΔI	ΔPS, ΔPF, ΔYF and ΔF	2.6287		Non-rejection of the null

The null hypothesis of non-existence of a long-run relationship between the UK tourism demand in France and its determinants is strongly rejected by the data, reinforcing the conclusion obtained with the conventional unit root tests above. The PP procedure also indicates that there is no other significant long-run relationship among the variables. This indicates that the explanatory variables in the UK demand for France equation can be viewed as 'long-run forcing' variables.

Regressors' exogeneity is one relevant aspect of empirical models' evaluation which has been addressed in many studies, for example, Richard (1980), Engle, Hendry and Richard (1983), Hendry (1983), Johansen (1988) and Johansen and Juselius (1990). The validity of the estimation results of static models depends, among other requisites, on the assumption that all regressors are exogenous. Hence, to sanction the statistical inference based on the estimation results of such models, exogeneity of the explanatory variables has either to be assumed or statistically proved. The concept of exogeneity is a complex one, ranging from 'weak' to 'super' exogeneity. The definitions of its various levels involve

[17] In some circumstances, the results of these additional tests for the existence of 'long-run forcing' variables may not be very powerful. A more appropriate procedure to test for exogeneity of a set of variables is the "fully maximised likelihood systems procedure" proposed in Johansen (1988) and in Johansen and Juselius (1990). This procedure is used in chapter 8 within a system of equations' approach.

mathematical derivations and statistical representations which are outside the scope of this chapter. Therefore, we address this issue in a summarised way, based on Hendry's (1995, pp. 181-191) exposition of "weak exogeneity and unit roots".

Let the long-run cointegrating relationship between the UK tourism demand for destination i and its determinants be

$$Yi_t = \alpha_0 + \alpha_1 Pi_t + \alpha_2 Pj_t + \alpha_3 I_t + \varepsilon 1_t \qquad (4.11)$$

where Pi and Pj are the price of tourism in destination i and competitive destination j, respectively, $\varepsilon 1$ is the disturbance term and the other variables are defined as before. Suppose that the PP procedure indicates, alongside the long-run relationship expressed by (4.11), another long-run relationship between Pi and all the other variables; that is, it indicates Pi as a potential endogenous variable. The relationship between Pi and all the other variables can be written as

$$\Delta Pi_t = \varphi \Delta Pi_{t-1} + \rho\left[Yi_{t-1} - \left(\alpha_0 + \alpha_1 Pi_{t-1} + \alpha_2 Pj_{t-1} + \alpha_3 I_{t-1}\right)\right] + \varepsilon 2_t \qquad (4.12)$$

where the expression in brackets represents the lagged error term of equation (4.11), $\varepsilon 1$.

According to Hendry (1997), there are three configurations of parameter values where exogeneity holds:

a) When $\varphi = \rho = 0$ and there is no contemporary correlation between $\varepsilon 1$ and $\varepsilon 2$, that is, $cov(\varepsilon 1_t, \varepsilon 2_t)=0$, then (4.11) is a valid regression equation between I(1) variables and Pi_t is both weakly and strongly exogenous for the parameters of interest α_i.

b) When $\varphi = \rho = 0$ but $cov(\varepsilon 1_t, \varepsilon 2_t) \neq 0$, then equation (4.11) suffers from 'simultaneity bias' but a valid regression equation can be defined if the variable ΔPi_t is included in the long-run relationship and Pi_t is both weakly and strongly exogenous for the parameters of interest. The addition of the impact variable ΔPi_t 'corrects' for the contemporaneous correlation between $\varepsilon 1$ and $\varepsilon 2$ and restores valid single-equation inference.

c) When $\rho = 0$ and $cov(\varepsilon 1_t, \varepsilon 2_t)=0$ but $\varphi \neq 0$, Pi_t cannot be strongly exogenous for the parameters of interest, but could be weakly exogenous. In this case, equation (4.12) is uninformative about the α parameters of interest, allowing single equation inference to be valid without loss of information.

The ARDL model is a dynamic specification which includes current as well as lagged values of the explanatory variables. Moreover, if an ARDL model includes the appropriate lag length for all variables, that is, if the model is correctly specified, the possible contemporaneous correlation between the error terms is already accounted for. Hence, there is no need to include any 'impact variables' to correct for contemporaneous correlation in well-specified ARDL models. In fact, Pesaran and Shin (1995) have shown that "*valid asymptotic inferences on the short-run and long-run parameters can be made, using the least squares estimates of the ARDL model, once the order of the ARDL model is appropriately augmented, to allow for possible contemporaneous correlation between the error terms. Therefore, the ARDL method continues to be applicable even if the explanatory variables are endogenous and irrespective of whether they are I(1) or not*", (Pesaran and Pesaran 1997, pp.183-4). Consequently, if the existence of a long-run cointegrated relationship between the dependent variable and its regressors cannot be rejected, the existence of potentially endogenous regressors of the appropriately augmented ARDL model should not affect the validity of the statistical inference based on its estimation results.

Given these considerations and since the tests performed indicate the existence of a long-run relationship between the UK demand for France and its determinants, we conclude that this relationship comprises a steady-state equilibrium path acting as an 'attractor' which, through an error-correction mechanism, prevents the variables from wandering too far apart, even if disturbed by shocks in the short-run.

We now turn to the cointegration analysis of the relationship between the relevant variables within the equation for the UK tourism demand in Spain.

4.3.2 The UK demand for tourism in Spain

The starting point for explaining the UK demand for tourism in Spain is the definition of an ARDL general specification similar to the one used previously. Hence, it is assumed that the UK tourism demand in Spain can be described by the following equation:

$$\begin{aligned} YS_t = \alpha + a_0PP_t + a_1PP_{t-1} + b_0PS_t + b_1PS_{t-1} + c_0PF_t + c_1PF_{t-1} + d_0YP_t + d_1YP_{t-1} + \\ + e_0YF_t + e_1YF_{t-1} + \beta_0 I_t + \beta_1 I_{t-1} + \delta_1 S7576_t + \delta_2 S8689_t + 0YS_{t-1} + u_t \end{aligned} \tag{4.15}$$

where S7576 and S8689 represent dummy variables. The former accounts for the events which affected Spain's tourism following the Portuguese revolution of 1974; the latter reflects the consequences for Spain of the Portugal and Spain joining the EU in 1986.

The estimation results for equation (4.15) are presented in table 4.4.

Table 4.4: Estimation results of the ARDL model for the UK tourism demand in Spain

	Regression 1	Regression 2
Intercept	-2.1633 (-1.25)	-2.7561 (-4.78)
PP_t	-0.5765 (-1.28)	
PP_{t-1}	0.6273 (1.11)	
PS_t	-0.0573 (-0.12)	-0.7219 (-2.90)
PS_{t-1}	-1.0731 (-1.73)	
PF_t	0.8412 (1.63)	1.0444 (3.56)
PF_{t-1}	0.0294 (0.06)	-0.5367 (-2.23)
I_t	1.3607 (1.72)	1.1209 (5.11)
I_{t-1}	-0.4067 (-0.48)	
YP_t	0.0883 (0.63)	0.2687 (2.88)
YP_{t-1}	0.2970 (3.28)	
YF_t	0.0060 (0.02)	
YF_{t-1}	0.1544 (0.92)	
S7576	0.1202 (2.54)	0.1117 (3.95)
S8689	0.0609 (2.48)	
YS_{t-1}	0.0129 (0.10)	0.2713 (2.33)
Adjusted R^2	0.9552	0.9294
Residual Sum Square	0.009	0.023
F-statistic	39.37	51.78
Serial Correlation	3.54 (0.06)	1.14 (0.29)
Functional Form	2.60 (0.11)	0.03 (0.87)
Normality	0.54 (0.76)	0.24 (0.89)
Heteroscedasticity	1.64 (0.20)	0.54 (0.46)

The structure of table 4.4 is similar to that of table 4.2. The procedure used to reduce the general regression 1, to specific regression 2 was as follows. Considering the estimation results of regression 1 we tested, for the dummy variables' coefficients, the hypothesis H_0: $\delta_1=2\delta_2=\delta$ which was not rejected by the Wald test with the statistic value $\chi^2(1)=0.0004(0.99)$. Therefore the dummy variables S7576 and S8689 are merged into a new dummy (S) which takes value 1 in the period 1975-1976, 0.5 in the period 1986-1989 and zero otherwise. High collinearity between YS_{t-1}, YP_{t-1} and the new dummy variable S was detected, indicating that the contribution of S and YS_{t-1} to the explanation of the dependent variable almost overlaps that of the variable YP_{t-1}. Consequently, this variable was omitted. We also performed an F-test for the joint significance of additional variables PP_t, PP_{t-1}, PS_{t-1}, I_{t-1}, YF_t and YF_{t-1} against regression 2 specification. The null of no added significance was not rejected with an F statistic value of 0.886(0.53).

All statistics and diagnostic tests presented in table 4.4 indicate regression 2 as a good specification. There is no statistical evidence of heteroscedasticity or residual serial correlation, the residuals are normally distributed, it presents a high explanatory power, good fit and acceptable functional form. Therefore, we take regression 2 as a good representation of the relationship between the UK tourism demand in Spain and its determinants. From regression 2 estimation results, we can derive the long-run equilibrium model of UK tourism demand in Spain and compute the long-run coefficients' estimates. These estimates supplied by the ARDL (1,0,1,0,0) are:

Intercept: $\hat{\alpha}_0 = \dfrac{\hat{\alpha}}{1-\hat{\theta}} = \dfrac{-2.7561}{1-0.27126} = -3.7820$

Price of Spain: $\hat{\alpha}_1 = \dfrac{\hat{b}_0}{1-\hat{\theta}} = \dfrac{-0.72186}{1-0.27126} = -0.99055$

Price of France: $\hat{\alpha}_2 = \dfrac{\hat{c}_0+\hat{c}_1}{1-\hat{\theta}} = \dfrac{1.0444-0.5367}{1-0.27126} = 0.69659$

UK Income: $\hat{\alpha}_3 = \dfrac{\hat{\beta}_0}{1-\hat{\theta}} = \dfrac{1.1209}{1-0.27126} = 1.5381$

UK Expenditure in Portugal: $\hat{\alpha}_4 = \dfrac{\hat{d}_0}{1-\hat{\theta}} = \dfrac{0.26873}{1-0.27126} = 0.36876$

Dummy S: $\hat{\alpha}_5 = \dfrac{\hat{\delta}}{1-\hat{\theta}} = \dfrac{0.11171}{1-0.27126} = 0.15329$

Velocity of adjustment: $1-\hat{\theta} = 1-0.27126 = 0.72874$

Therefore, the long-run equilibrium relationship for the UK demand for Spain is (t-ratios in brackets):

$$YS_t = \underset{-(5.96)}{-3.782} - \underset{(-2.58)}{0.991}\,PS_t + \underset{(1.85)}{0.697}\,PF_t + \underset{(7.47)}{1.538}\,I_t + \underset{(2.75)}{0.369}\,YP_t + \underset{(4.16)}{0.153}\,S_t + \hat{u}_t \qquad (4.16)$$

The residuals of regression (4.16) constitute the EQE series which is used to construct the ECM associated with the long-run relationship derived from the ARDL model. The EQE series is given by

$$EQE_t = YS_t - \left(-3.782 - 0.991PS_t + 0.697PF_t + 1.538I_t + 0.369YP_t + 0.153S_t\right) \qquad (4.17)$$

The error-correction model associated with the long-run relationship is given by the following equation (t-ratios in brackets):

$$\Delta\hat{YS}_t = \underset{(-2.90)}{-0.7219}\,\Delta PS_t + \underset{(3.56)}{1.044}\,\Delta PF_t + \underset{(5.11)}{1.121}\,\Delta I_t + \underset{(2.88)}{0.269}\,\Delta YP_{t-1} + \underset{(3.95)}{0.112}\,\Delta S_t - \underset{(-6.25)}{0.729}\,EQE_{t-1}$$

Adjusted R^2=0.712 RSS=0.0231 F(6, 21)=12.28 DW=2.30

The next step is to apply the conventional unit root tests to the EQE series. If the series is stationary then the long-run equilibrium relationship expressed in equation (4.16) is meaningful and the variables involved are cointegrated. The results of unit root tests for EQE produced the following results:[18]

	EQE_t series	MacKinnon critical values[19]	
		5%	10%
ADF(0)	-4.1948	-4.5643	-4.1510
ADF(1)	-4.4975		
ADF(2)	-4.9520		
ADF(3)	-5.1480		

The results show that at the 5% significance level the EQE series is stationary and, hence, the variables in the long-run relationship are cointegrated.

As before, we apply the PP procedure to confirm the existence of the long-run relationship indicated by the conventional unit root tests and to establish whether the

[18] The SBC and AIC criteria indicate 3 lags as the appropriate lag length for the test regression to eliminate autocorrelation in the EQE series. The regression is estimated with intercept, no trend and 25 observations.
[19] The MacKinnon critical values were computed considering 25 observations, an intercept, no trend and four I(1) regressors for which cointegration is being tested.

regressors are "long-run forcing" variables. The PP error-correction version of the ARDL model for the UK demand for tourism in Spain is

$$\Delta YS_t = \gamma_0 + \gamma_1 \Delta PS_t + \gamma_2 \Delta PF_t + \gamma_3 \Delta I_t + \gamma_4 \Delta YP_t + \gamma_5 \Delta S_t + \lambda_1 PS_{t-1} + \lambda_2 PF_{t-1} + \lambda_3 I_{t-1} + \lambda_4 YP_{t-1} + \lambda_5 S_{t-1} + \lambda_6 YS_{t-1} + u_t \quad (4.18)$$

and the hypothesis to be tested is whether a long-run relationship exists between the variables YS, PS, PF, I, YP and S. The hypothesis under test is the null of non-existence of a long-run relationship, defined by H_0: $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = \lambda_6 = 0$ against the alternative H_1: $\lambda_i \neq 0$, for any i =1, ...,6. As before, the relevant statistic for this test is the F-statistic and the appropriate critical values are included in Pesaran *et al.* (1996). The results obtained using this procedure are presented in Table 4.5.

Table 4.5: Additional variables significance test for YS_{t-1}, PS_{t-1}, PF_{t-1}, I_{t-1}, YP_{t-1} and S_{t-1}

Dependent variable	Explanatory variables	F-statistic (6, 16)	Critical band at the 5% level	Test result
ΔYS	ΔYS, ΔYF, ΔI, ΔYP and ΔS	5.0172	(2.850; 4.049)	Rejection of the null
ΔPS	ΔYS, ΔPF, ΔI, ΔYP and ΔS	1.1216		Non-rejection of the null
ΔPF	ΔPS, ΔYS, ΔI, ΔYP and ΔS	1.6770		Non-rejection of the null
ΔI	ΔPS, ΔPF, ΔYS, ΔYP and ΔS	2.1567		Non-rejection of the null
ΔYP	ΔPS, ΔPF, ΔYS, ΔYP and ΔS	1.5938		Non-rejection of the null

The null hypothesis of non-existence of a long-run relationship between the UK demand for Spain and its determinants is rejected by the data, supporting the conclusion already obtained with the conventional unit root test above. The PP procedure also indicates that there is no statistical evidence of other long-run relations among the variables except for that linking the dependent variable (YS) and its explanatory variables. Hence, we can infer that the UK tourism demand for Spain comprises a meaningful long-run equilibrium path acting as an 'attractor' which prevents the variables from drifting apart, even if disturbed in the short-run. We now turn to the cointegration analysis of the UK demand for tourism in Portugal.

4.3.3 THE UK DEMAND FOR TOURISM IN PORTUGAL

To explain the UK demand for tourism in Portugal we start with an ARDL general specification similar to the ones used previously. Hence, it is assumed that the UK demand for tourism in Portugal can be described by the following equation:

$$\begin{aligned} YP_t = \alpha + a_0 PP_t + a_1 PP_{t-1} + b_0 PS_t + b_1 PS_{t-1} + c_0 PF_t + c_1 PF_{t-1} + \beta_0 I_t + \beta_1 I_{t-1} + \\ + d_0 YS_t + d_1 YS_{t-1} + e_0 YF_t + c_1 YF_{t-1} + \delta_1 P7479_t + \delta_2 P8391 + 0 YP_{t-1} + u_t \end{aligned} \tag{4.19}$$

The dummy variable P7479 corresponds to the Portuguese revolution period. The dummy P8391 related to the period 1983-1991 deserves further explanation. This period corresponds to the pre-joining (1983-85) and the post-joining periods (1986-91) of the Portuguese economy's integration in the EU. Just emerging from the ordeals of a revolution and still enduring the "straight jacket" of restrictive economic policies imposed by the IMF, the traditionally closed, weak and unstable Portuguese economy had to undertake major reforms necessary to achieve the openness and flexibility required for its full integration among the stronger EU economies. The Portuguese economy lagged behind the other European economies to such a high degree that the European authorities allowed an extended adaptation period which, officially, lasted until 31 December 1989 but 'unofficially' extended beyond that date. Therefore, the intercept dummy variable P8391 is added to the equation (4.19) to account for these events.

The estimated results of equation (4.19) are presented in table 4.6.

Table 4.6: Estimation results of the ARDL model for the UK tourism demand in Portugal

	Regression 1	Regression 2	Regression 3
Intercept	-2.5674 (-0.64)	0.9416 (1.37)	0.9407 (1.45)
PP_t	-1.6415 (-1.70)	-0.8424 (-3.57)	-0.8430 (-4.20)
PP_{t-1}	-0.0600 (-0.04)		
PS_t	0.6904 (0.56)	1.0381 (2.74)	1.0381 (2.81)
PS_{t-1}	0.8940 (0.61)		
PF_t	1.1783 (1.05)		
PF_{t-1}	-1.1255 (-0.97)		
I_t	1.5390 (1.08)	1.6300 (2.35)	1.6305 (2.43)
I_{t-1}	-0.4445 (-0.26)	-1.7633 (-2.43)	-1.7636 (-2.50)
YS_t	0.1351 (0.37)		
YS_{t-1}	0.0439 (0.15)		
YF_t	-0.8014 (-1.45)		
YF_{t-1}	0.4158 (1.47)		
P7479	-0.1557 (-1.90)	-0.1624 (4.27)	-0.1623 (-4.71)
P8391	-0.0821 (-1.51)	-0.0534 (-2.10)	
YP_{t-1}	0.2667 (1.01)	0.3992 (3.74)	0.3994 (4.02)
Adjusted R^2	0.7797	0.8360	0.8438
Residual Sum Square	0.035	0.044	0.044
F-statistic	7.37	20.66	25.31
Serial Correlation	2.06 (0.15)	1.68 (0.20)	1.58 (0.21)
Functional Form	1.34 (0.25)	0.26 (0.61)	0.24 (0.63)
Normality	0.146 (0.93)	1.67 (0.43)	1.68 (0.43)
Heteroscedasticity	0.00 (0.97)	1.16 (0.28)	1.16 (0.28)

The search for the model expressed by regression 3 obeys the same rules used before; that is, we start from the general ARDL model including all variables and, by means of successive tested restrictions, reduce it to the specific form of regression 3. We performed the joint significance F-test for additional variables PP_{t-1}, PS_{t-1}, PF_t, PF_{t-1}, YS_t, YS_{t-1}, YF_t and YF_{t-1} against the regression 2 specification. This test does not reject the null of no significance of the additional variables, with an F statistic value of 0.357(0.925). After imposing the restriction suggested by the F test, we established that the coefficient estimate of variable P7479 was approximately three times the coefficient estimate of variable P8391. Therefore, we tested, for the dummy variables' coefficients, the hypothesis $H_0{:}\delta_1=3\delta_2=\delta$ which was not rejected by the Wald test with the statistic value of $\chi^2(1)=0.011(0.92)$. Therefore, the dummy variables P7479 and P8391 were merged into a new dummy P which assumes the value of unity in the period 1974-1979, 0.33 in the period 1983-1991 and zero otherwise. This new variable P is included in the regression 3 equation.

From the regression 3 estimation results we can derive the long-run relationship between the UK demand for Portugal and its determinants, and compute the long-run coefficients' estimates. The long-run estimates given by the ARDL (1,0,0,1) model are:

$$\text{Intercept: } \hat{\alpha}_0 = \frac{\hat{\alpha}}{1-\hat{\theta}} = \frac{0.94069}{1-0.60065} = 1.5661$$

$$\text{Price of Portugal: } \hat{\alpha}_1 = \frac{\hat{a}_0}{1-\hat{\theta}} = \frac{-0.84304}{1-0.60065} = -1.4035$$

$$\text{Price of Spain: } \hat{\alpha}_2 = \frac{\hat{b}_0}{1-\hat{\theta}} = \frac{1.0381}{1-0.60065} = 1.7283$$

$$\text{UK Income: } \hat{\alpha}_3 = \frac{\hat{\beta}_0+\hat{\beta}_1}{1-\hat{\theta}} = \frac{1.6305-1.7636}{1-0.60065} = -0.2215$$

$$\text{Dummy P: } \hat{\alpha}_4 = \frac{\hat{\delta}_1}{1-\hat{\theta}} = \frac{-0.16232}{1-0.60065} = -0.27024$$

$$\text{Velocity of adjustment: } 1-\hat{\theta} = 1-0.39935 = 0.60065$$

Hence, the equilibrium relationship for the UK demand for tourism in Portugal is given by:

$$YP_t = \underset{(1.29)}{1.566} - \underset{(-4.14)}{1.404}\,PP_t + \underset{(2.53)}{1.728}\,PS_t - \underset{(-0.60)}{0.222}\,I_t - \underset{(-3.78)}{0.270}\,P_t + \hat{u}_t \qquad (4.20)$$

The residuals of regression (4.20) constitute the EQE series which is used to construct the ECM associated with the long-run ARDL model. The EQE series is given by

$$EQE_t = YP_t - \left[1.566 - 1.404\,PP_t + 1.728PS_t - 0.222I_t - 0.270P_t\right]$$

The estimated results of the error-correction representation for the selected ARDL model are as follow (t-ratios in brackets):

$$\Delta\hat{Y}P_t = -\underset{(-4.20)}{0.843}\,\Delta PP_t + \underset{(2.81)}{1.038}\,\Delta PS_t + \underset{(2.43)}{1.631}\,\Delta I_t - \underset{(-4.71)}{0.162}\,\Delta P_t - \underset{(-6.05)}{0.601}\,EQE_{t-1}$$

Adjusted R^2=0.730 RSS=0.044 F(5, 22)=15.84 DW=2.31

The conventional unit root test for this EQE provides the following results:[20]

Test	EQE_t series	MacKinnon critical values[21]	
		5%	10%
ADF(0)=DF	-4.9110	-4.0583	-3.6785

The results show that, at the 5% significance level, the EQE series is stationary and the variables in the long-run relationship are cointegrated.

As before, we apply the PP procedure to confirm the existence of the long-run relationship indicated by the conventional unit root tests, and to establish whether the regressors are "long-run forcing" variables. The PP error-correction version of the ARDL model for the UK tourism demand in Portugal is

$$\begin{aligned}\Delta yP_t &= \gamma_0 + \gamma_1 \Delta pP_t + \gamma_2 \Delta pS_t + \gamma_3 \Delta I_t + \gamma_4 \Delta P_t \\ &\quad + \lambda_1 pP_{t-1} + \lambda_2 pS_{t-1} + \lambda_3 I_{t-1} + \lambda_4 P_{t-1} + \lambda_5 yP_{t-1} + w_t\end{aligned} \tag{4.21}$$

and the hypothesis to be tested is whether a long-run relationship exists between the variables YP, PS, PP, I and P. Hence, the null of non-existence of a long-run relationship is defined by H_0: $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = 0$ against the alternative H_1: $\lambda_i \neq 0$, for any i =1, ...,5. As before, the relevant statistic for this test is the F-statistic and the appropriate critical values are found in Pesaran *et al.*'s (1996) tables. The results obtained using this procedure are presented in table 4.7.

[20] The SBC and AIC criteria indicate that the appropriate lag-length for the test regression with intercept and no trend is zero. Therefore, we use the ADF(0) test with 28 observations available.
[21] The MacKinnon critical values were computed considering 28 observations, an intercept, no trend and three I(1) regressors for which the null of non-cointegration is being tested.

Table 4.7: Joint significance F-test for additional variables YP_{t-1}, PS_{t-1}, PPt_{t-1}, I_{t-1} and P_{t-1}

Dependent variable	Explanatory variables	F-statistic (5, 18)	Critical band at the 5% level	Test result
ΔYP	ΔPS, ΔPP, ΔI, and ΔP	6.0411	[3.219; 4.378]	Rejection of the null
ΔPP	ΔPS, ΔYP, ΔI, and ΔP	6.1219		Rejection of the null
ΔPS	ΔPP, ΔYP, ΔI, and ΔP	6.3743		Rejection of the null
ΔPI	ΔPS, ΔPP, ΔYP, and ΔP	2.1197		Non-rejection of the null

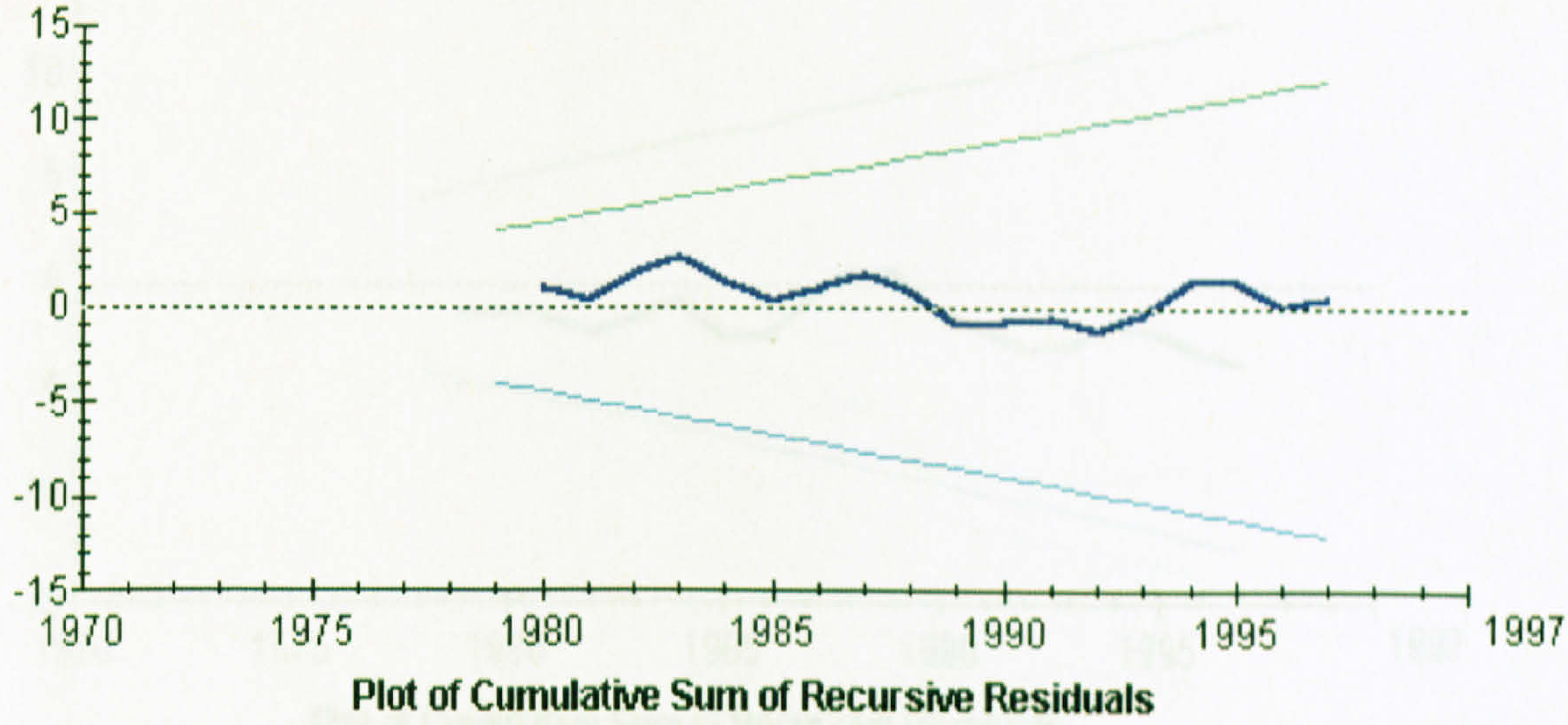
The above results suggest that a long-run relationship exists between the UK tourism demand for Portugal and its determinants, supporting the indication obtained with the conventional unit root tests above. The PP test also suggests that Spain and Portugal price variables cannot be treated as 'long-run forcing' for the explanation of YP that is, the test indicates the possibility of PS and PP being endogenous.[22]

4.3.4. Analysis of the estimated results for the UK tourism demand in France, Spain and Portugal

In this section, we start by establishing whether the regressions' coefficients are structurally stable. Then, we focus on the economic interpretation of the estimation results obtained in the previous section. The structural stability of the regressions' coefficients is tested by using the cumulative sum of residuals (CUSUM) and the cumulative sum of squared residuals (CUSUMSQ) tests proposed by Brown *et al.* (1975). The CUSUM test is particularly useful for detecting systematic changes in the regression coefficients, whereas the CUSUMSQ test is useful for detecting random or sudden departures from the constancy of the regression coefficients. The graphs below show the plot of the CUSUM and CUSUMSQ statistics for each of the three ARDL equations. The pair of straight lines bordering the plot of the statistics represents the confidence interval at the 5% significance level. If either of the lines is crossed, the null hypothesis of correct specification for the regression equation must be rejected at the 5% level.

[22] See section 4.3.1. on the subject of endogenous regressors within a long-run relationship derived from an appropriately specified ARDL model.

Figure 4.11: CUSUM and CUSUMSQ statistics for the UK demand equation for France



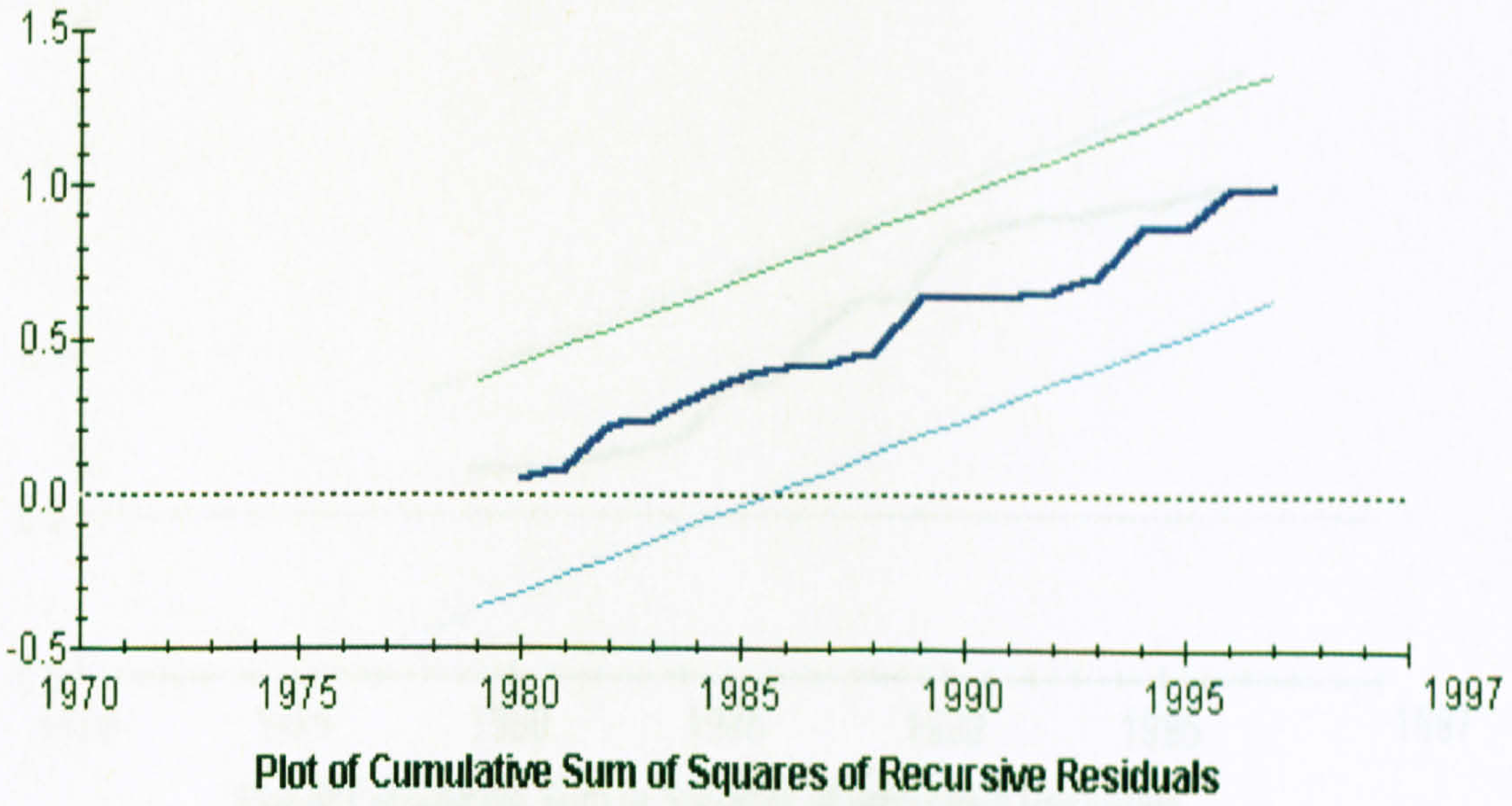
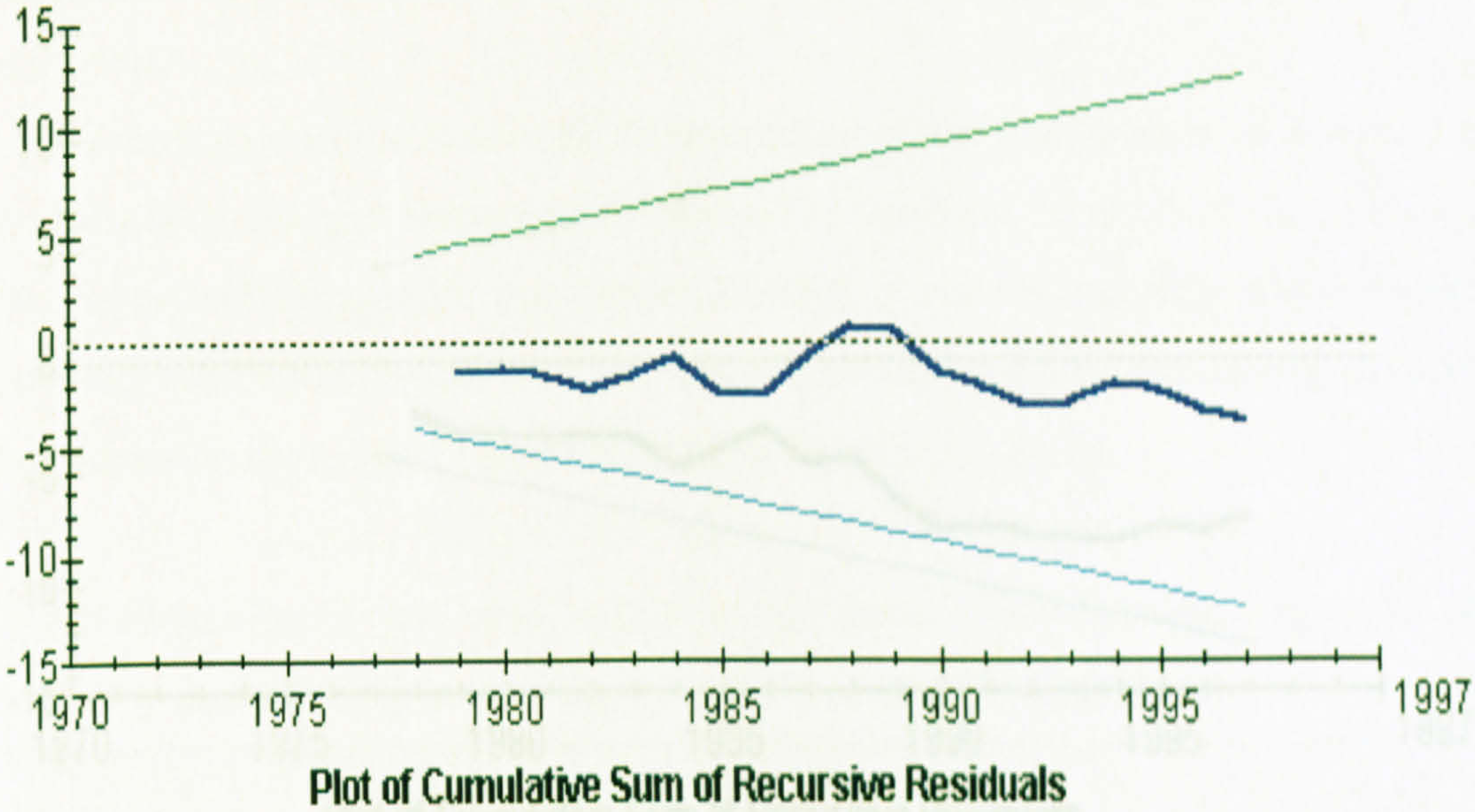


Figure 4.12: CUSUM and CUSUMSQ statistics for the UK demand equation for Spain



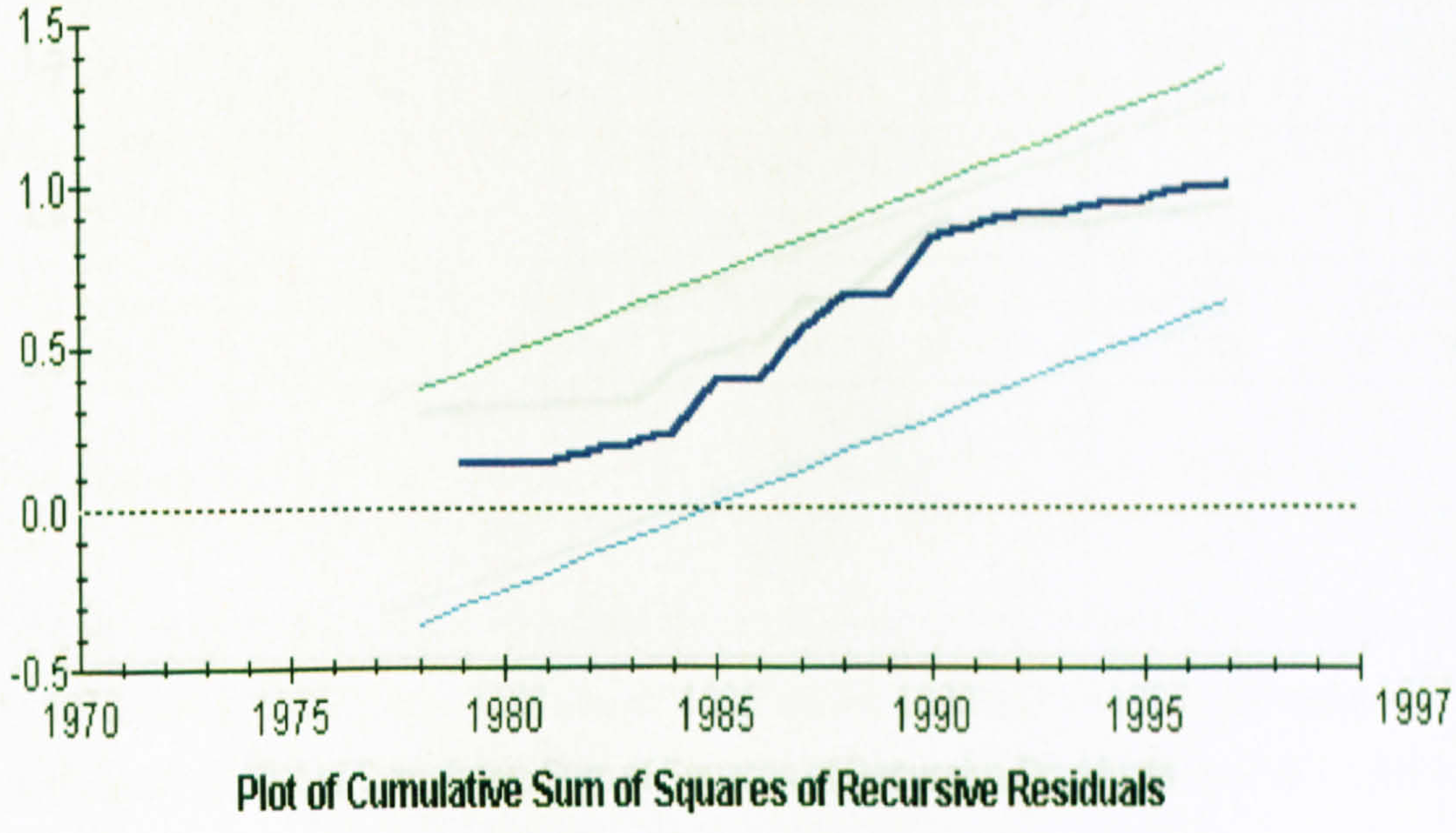
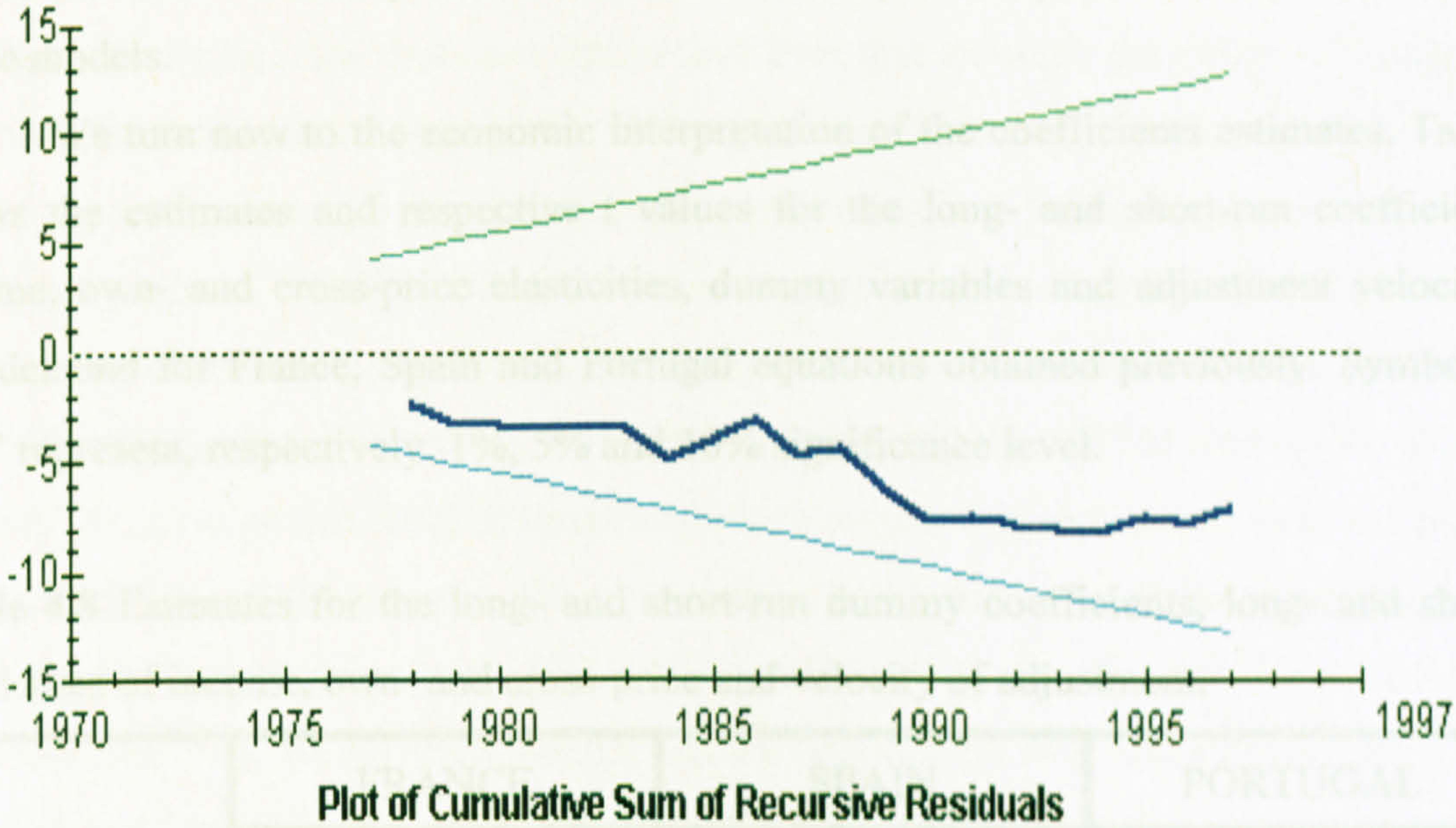
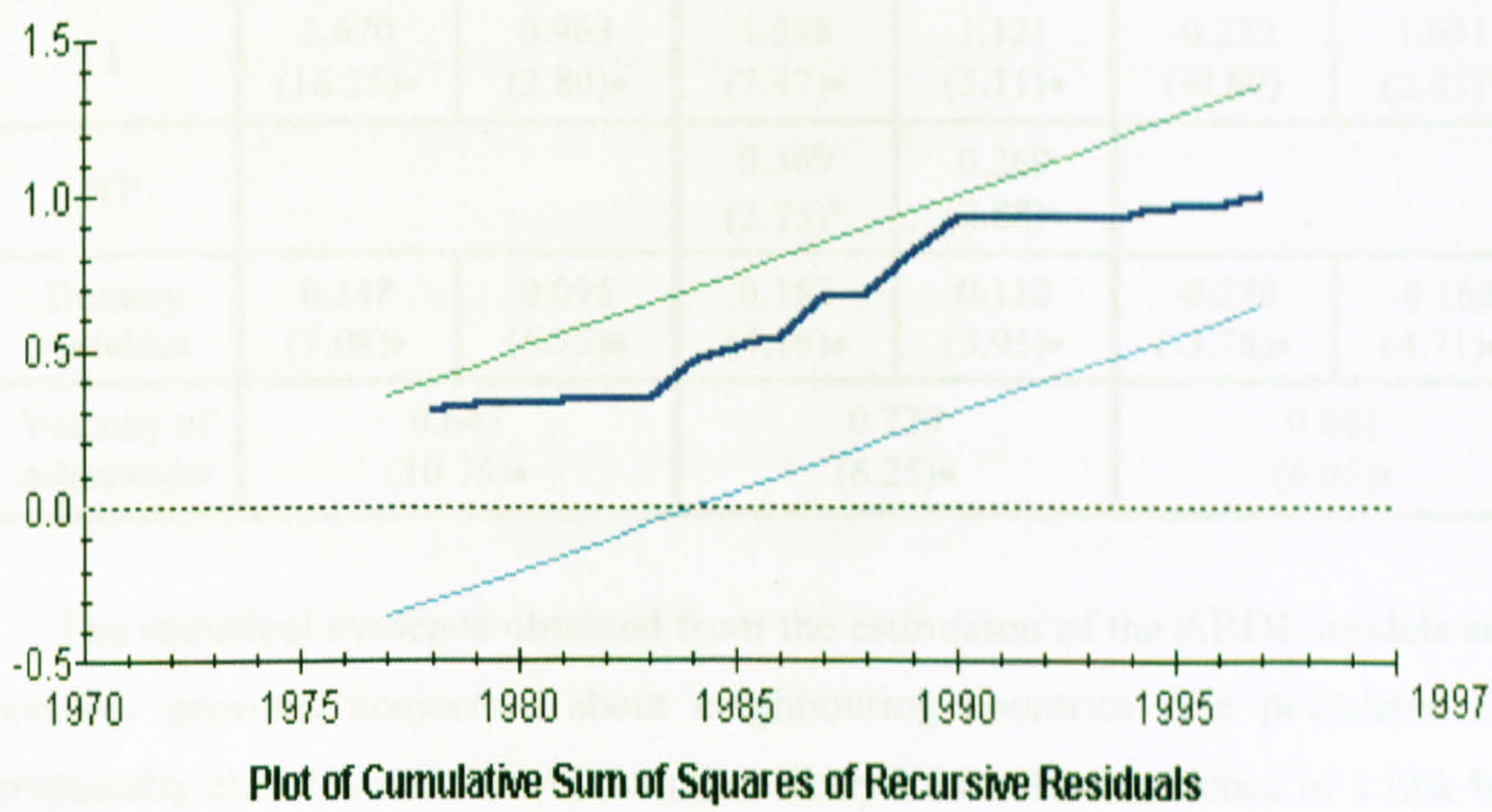


Figure 4.13: CUSUM and CUSUMSQ statistics for the UK demand equation for Portugal





The tests indicate that the null hypothesis of correctly specified equations, cannot be rejected at the 5% level of significance[24]. Therefore, we can accept that the regressions' coefficients are structurally stable and add one more positive point to the quality evaluation of the models.

We turn now to the economic interpretation of the coefficients estimates. Table 4.8 shows the estimates and respective t values for the long- and short-run coefficients of income, own- and cross-price elasticities, dummy variables and adjustment velocities of UK demand for France, Spain and Portugal equations obtained previously. Symbols •, ° and ' represent, respectively, 1%, 5% and 10% significance level.

Table 4.8 Estimates for the long- and short-run dummy coefficients, long- and short-run elasticities of income, own- and cross-price and velocity of adjustment.

	FRANCE		SPAIN		PORTUGAL	
	Long-run	Short-run	Long-run	Short-run	Long-run	Short-run
PP					-1.404 (-4.14)•	-0.843 (-4.20)•
PS	1.363 (4.60)•	-0.167 (-0.79)	-0.991 (-2.58)•	-0.722 (-2.90)•	1.728 (2.53)°	1.038 (2.81)•
PF	-2.917 (-10.34)•	0.223 (1.10)	0.697 (1.85)'	1.044 (3.56)•		
I	2.670 (16.25)•	0.963 (2.80)•	1.538 (7.47)•	1.121 (5.11)•	-0.222 (-0.60)	1.631 (2.43)°
YP			0.369 (2.75)°	0.269 (2.88)•		
Dummy variables	0.148 (7.08)•	0.095 (6.75)•	0.153 (4.16)•	0.112 (3.95)•	-0.270 (-3.78)•	-0.162 (4.71)•
Velocity of Adjustment	0.643 (10.75)•		0.729 (6.25)•		0.601 (6.05)•	

The statistical evidence obtained from the estimation of the ARDL models seems to support our previous conjecture about neighbouring countries. We postulated that the geographically closer countries are, the more likely it is to find evidence of a link between

[24] We also applied the recursive least squares regression and the rolling least squares regression estimation methods for testing the structural stability of the equations. These tests confirm the coefficients' stability indicated by the CUSUM and CUSUMSQ tests.

them, which indicates close neighbouring destinations as competitors rather than complements. In fact, the estimation results suggest a significant link between France and Spain through the price of Spain in the equation for France and the price of France in the equation for Spain, and between Spain and Portugal through the price of Spain in the equation for Portugal and through the UK expenditure in Portugal in the equation for Spain. However, no significant link is found between the equations for France and Portugal.

In general, the elasticities estimates have plausible magnitudes and the expected signs. When statistically significant, the income and own-price elasticities are, respectively, positive and negative as expected from normal commodities, and the cross-price elasticities are positive as expected from destinations which are competitors rather than complements. The velocities of adjustment are reasonably high and of similar magnitude across destinations, suggesting that UK tourists adjust fast to changes in their demand determinants. Yet, for all destinations, UK tourists need more than one period (15 to 16 months) to adjust fully their demand to its long-run equilibrium path.

In the equation for France, the estimate for the long-run own-price elasticity is negative, as expected, implying that an increase in tourism prices in France causes a fall in the UK tourism demand for this country. Tourism in France is, therefore, a normal commodity. However, while in the long-run the UK demand for tourism in France responds significantly and highly negatively to price changes in this country, in the short-run its response is positive although not statistically significant. The reason for these results may rest on the velocity of adjustment estimate (0.64), which indicates that UK tourists need more than one period (around 16 months) for fully adjusting to changes in the determinants of their demand for France. Therefore, short-run changes will not be perceived immediately (within one period) and plans already made for crossing the Channel will remain so, no matter what changes occur in prices.[25]

The positive sign of the long-run estimate for the Spain's price elasticity in the equation for France, indicates Spain as a competitor destination of France. If prices rise by 1% in Spain, the UK tourism demand for France increases by 1.4%. However, this elasticity estimate is not significant in the short-run, indicating that the UK tourism demand for France is not affected by short-term price changes in Spain. On the other hand, the UK

[25] Note that increases in demand are measured by increases in expenditure and although the number of tourists crossing the Channel may remain the same (or even diminish) in response to an increase in prices, those that have already decided to do so are likely to maintain their decision, which will increase expenditure (demand) due to the increase in prices but not necessarily the number of tourists.

demand for tourism in Spain is sensitive to price changes in France both in the long- and in the short-run, and more so in the short-run. The demand for France is income elastic in the long-run and inelastic in the short-run, while that for Spain is income elastic both in the long- and in the short-run. A possible reason might be the fact that Spain is a "primary" destination relative to its neighbours. Thus, short-lived increases in UK tourists' budgets (a lucky bet in a wining horse, for example), if directed to tourism, will preferably be spent in Spain rather than France. The demand for Portugal is income elastic only in the short-run, as increases in the UK tourism budget in the long-run do not affect Portugal significantly.

The long- and short-run estimates for the dummy variable coefficient indicate positive impacts on the UK demand for France in the periods 1975-1981, 1987-1989 and 1994-1997, positive impacts on the UK demand for Spain in the periods 1975-1976 and 1986-1989 and negative impacts on the UK demand for Portugal in the periods 1974-1979 and 1983-1991. It is our hypothesis that, because of their geographical location, these destination countries share a related faith in UK tourists' preferences. The fact that, for all destinations' equations, the dummy variables included are defined over similar periods may be interpreted as an extra link among these neighbouring countries concerning the UK demand for tourism. We believe that the change of political regimes to democracy in Portugal and Spain in the mid 1970's, their subsequent simultaneous joining the EU on the 1st of January 1986, the substantial appreciation of the Spanish peseta between 1986 and 1991,[26] and the opening of the Channel Tunnel in 1994, brought about, although in different magnitudes and signs, consequences which affect the UK tourism demand for these destinations both in the short- and long-run. France seems to have profited in the short-run, from the upheavals of the Portuguese revolution (1974) and political regime change in Spain (1976) and, in the long-run, from the stability of the democratic regimes established in both these countries, their joining the EU and from the opening of the Channel Tunnel, compounded with the decreasing popularity of some traditional tourism resorts in Spain since the mid 1990's. The long- and short-run estimates of the dummy coefficient in the equation for Spain indicate similar effects based on reasons similar to those corresponding to the equation for France, meaning that UK tourism in Spain may

[26] Following Spain's entry into the EU, consumption and investment boomed and the real value of the Spanish peseta increased substantially against the currencies of Spain's major European trading partners. For example, the peseta/deutsche mark real exchange rate appreciated almost 30% by 1991. However, the process was reversed starting in 1992: investment fell sharply and the real exchange rate depreciated. A similar account of facts can be given for the case of Portugal, bearing in mind the differences of size and strength between the two economies.

have benefited in the short-run from the ordeals of the Portuguese revolution, particularly in the period 1975-76, and in the long-run from Spain and Portugal's change of political regimes and entrance in the EU. The political and social upheavals in Portugal during the mid 1970's, the vulnerability of its economy then and during the 1980's, and its slow and uneven recovery ever since, seem to have 'long-memory' effects on UK tourists' preferences, which account for the negative impacts of the dummy coefficients on the UK demand for tourism in Portugal. Furthermore, these effects seem to bear some responsibility for the non-significance of the long-run income elasticity estimate.

The positive signs of the elasticity estimates of the UK demand for Spain relating to price changes in France indicate this country as a competitor destination of Spain and more so in the short- than in the long-run. Price increases in France have a positive, elastic and significant (1% level) effect in the UK tourism demand for Spain in the long-run, while in the short-run, this effect is also positive but inelastic and only significant at the 10% level.

In the equation for Spain, the results indicate the own-price elasticity of the UK demand as negative and below unity both in the long- and short-run. Therefore, tourism in Spain is a normal good. The estimates for the sensitivity of UK tourism demand to price changes in Spain suggest that a permanent rise in prices in Spain has a negative impact on the UK demand for Spain, but a positive effect on the UK demand for France and Portugal. This seems to support further the hypothesised existence of a strong link between the UK tourism demand for Spain and France on the one hand, and that for Spain and Portugal on the other. The association between the UK tourism demand for Spain and France is supported by the significant price effects between these two countries, qualifying them as competitors in UK tourists' preferences. The association between the UK tourism demand for Spain and Portugal is supported, in the equation for Spain, by the significant and positive effects that increases in the UK tourism expenditure in Portugal have on tourism demand for Spain and, in the equation for Portugal by the positive and significant effects that price changes in Spain have in the UK demand for Portugal. Price changes in Portugal do not affect UK tourism in Spain, while price changes in Spain do affect, positively and significantly, UK tourism in Portugal, both in the short- and in the long-run. As for France, UK income increases also have positive effects on the UK tourism demand in Spain, both in the short- and in the long-run. However, these effects are greater for Spain in the short-run, and for France in the long-run. Therefore, France tends to benefit more than Spain from permanent increases in UK tourists' wealth, while Spain tends to benefit more than France from short-run positive impacts on the UK per capita income.

In the equation for Portugal, the estimates indicate the own-price elasticity as negative and inelastic in the short-run and negative and elastic in the long-run. They also indicate that the sensitivity of UK demand for tourism in Portugal towards price changes in neighbouring Spain is significant, positive and elastic both in the short- and in the long-run. This shows Portugal as a destination competing with Spain for the UK tourists' preferences.

Apparently, in the short-run, own-price increases in Portugal affect the UK demand for this destination more negatively than own-price increases in Spain or France. In the long-run, the negative effects of own-price increases are greater for France and Portugal than for Spain. The high relative sensitivity of the UK demand for Portugal towards own-price changes in Portugal and price changes in Spain, compounded with the positive effect that increases in the UK demand for Portugal have in that for Spain, should make prices a priority and Spain the key competitive destination for targeting in tourism policies implemented by Portuguese authorities. The UK income elasticity and adjustment velocity estimates further suggest that tourism policy changes should concentrate on the long- rather than on the short-run. The short-run income elasticity estimate indicates Portugal as a top preference among the three neighbouring destinations for UK tourists whilst, in the long-run, increases in UK tourists' income do not have a significant impact on their demand for Portugal. In contrast, positive, significant and high long-run effects on tourism demand are expected for Spain and, particularly, for France from UK income increases. Therefore, tourism policy in Portugal should focus on reverting some of these effects in its favour.

The above considerations seem to be consistent with the underlying economic theory, permitting the specifications and the econometric methodology applied in their derivation to be seen as reliable means to analyse the UK tourism demand for France, Spain and Portugal. Other interesting aspects may arise from comparisons between these results and those of other studies using similar methodologies and identical origin-destinations.

4.4. EMPIRICAL STUDIES USING A COMPARABLE METHODOLOGY FOR THE ANALYSIS OF UK TOURISM DEMAND

Studies of tourism demand using single equation models which are derived from economic theory and apply appropriate methodological tools for their specification, estimation and testing are few in number. These include Syriopoulos (1995), Kulendran (1996), Kim and Song (1998), Vogt and Wittayakorn (1998) and Song, Romilly and Liu (2000). The studies by Syriopoulos (1995) and Song *et al.* (2000) are particularly suitable for the comparison purpose we have set, because they apply the general to specific approach to construct empirical error-correction models of the UK demand for tourism in the same destinations we are concerned with. Compared with most previous literature in tourism demand analysis, these two studies show considerable advances in aspects both theoretical and methodological. Syriopoulos (1995) includes a dynamic dimension that was previously lacking from almost every empirical study on the subject. Song *et al.* (2000) apply the most recent cointegration analysis techniques for the evaluation of long-run relationships between tourism demand and its determinants in models derived with the Engel-Granger (1987) two-steps method.

Syriopoulos' (1995) study analyses the UK demand for tourism in several Mediterranean destinations including Portugal and Spain, during the period 1960-1987. This study considers a general to specific approach for identifying the long-run and short-run effects of determinants' changes on the dependent variable using an error-correction model. The model contains, simultaneously, long- and short-run information about the impacts of the regressors' changes on the UK demand for tourism. However, the error-correction mechanism specified in this model is imposed as a restricted or partial form of the 'normal' ECM that would emerge from an ARDL model as suggested by, for example, Pesaran and Pesaran (1997), or from an assumed long-run equilibrium relationship, as suggested by Engel and Granger (1987) two-steps method. However, Syriopoulos' study does not test for cointegration of the long-run relationship which is derived from his parsimonious final model. Therefore, we do not know if the long-run coefficients' estimates hold as steady-state equilibrium values or not. The UK demand for tourism in destination i is given by the following equation:[27]

$$\begin{aligned}\Delta y_t &= \beta_1\Delta I_t + \beta_2\Delta^2 I_t + \beta_3\Delta rp_t + \beta_4\Delta^2 rp_t + \beta_5\Delta e_t + \beta_6\Delta^2 e_t + \beta_7\Delta crp_t + \beta_8\Delta^2 crp_t + \beta_9\Delta ce_t \\ &\quad + \beta_{10}\Delta^2 ce_t + \delta\Delta y_{t-1} + \phi(y - I)_{t-2} + \gamma_1 I_{t-2} + \gamma_2 ep_{t-2} + \gamma_3 cep_{t-2} + \beta_{11}DV + u_t\end{aligned} \tag{a}$$

where 'y' is the UK demand for tourism measured as the origin's tourism expenditure in the relevant destination, 'I' is UK per capita real income, 'rp' and 'e' are, respectively, the relative price and exchange rate of a destination relative to the origin, 'crp' and 'ce' are,

[27] We changed the notation to facilitate interpretation and comparison.

respectively, relative price and exchange rate of a destination relative to competitors, 'ep' is the own-effective price of a destination and 'cep' are effective prices of its competitors. DV is a dummy variable. The variables' first and second differentials are denoted by Δ and Δ^2, respectively. The variables are measured in logs so their coefficients represent elasticities. The short-run effects are given by the coefficients attached to the variables in differences (β coefficients), while the long-run effects are given by the coefficients attached to the variables in levels (γ coefficients) when appropriately weighted by the velocity of adjustment ϕ. Tables 4.9 and 4.10 summarise Syriopoulos' estimation results for the regressions of Spain and Portugal.

Table4.9: ECM for the UK tourism demand in Spain and Portugal.

	SPAIN	PORTUGAL
ΔI_t	0.89 (3.67)	1.72 (3.67)
Δrp_t	-0.98 (-2.22)	
$\Delta^2 rp_t$	0.93 (1.24)	2.98 (3.54)
Δe_t	0.94 (1.86)	0.94 (1.94)
$\Delta^2 e_t$	-1.47 (-3.83)	-0.27 (-1.73)
Δcrp_t		-3.71 (-2.43)
Δy_{t-1}	-0.18 (-2.94)	-0.51 (-3.45)
$(y-I)_{t-2}$	-0.60 (-2.93)	-0.25 (-2.99)
I_{t-2}		0.45 (5.52)
ep_{t-2}	-1.26 (-3.68)	-0.46 (-2.96)
cep_{t-2}		-0.74 (-2.90)
DV		-0.38 (-3.59)

Source: Syriopoulos (1995) Table 1a, p.324.

Table 4.10: Long-run income, effective own-price and substitute-price elasticities for the UK demand for tourism in Spain and Portugal

	SPAIN	PORTUGAL
Income elasticity	1.00	2.80
Own-price elasticity	-2.10	-1.84
Substitute-price elasticity		-2.96[28]

Source: Syriopoulos (1995), tables 6, 7 and 8, p.330-332

Some of the estimation results presented are peculiar and seem difficult to interpret in a coherent way both in theoretical and prior empirical knowledge terms. For instance, it seems difficult to justify that the UK demand for Portugal does not react significantly to short-run own-price changes in Portugal, but reacts significantly and positively to the growth rate of short-run own-price changes. In fact, the coefficient estimate for Δ^2rp variable, which represents the growth rate of own-price changes in Portugal, is 2.98. This means that if the growth rate of own-price changes in Portugal increases by 1%, the UK demand for Portugal rises by approximately 3%. The magnitudes and signs of the estimates for the short- and long-run effects of changes in competitors' prices (cep_{t-2}) in the equation for Portugal present this country as a strong complement of all the other Mediterranean destinations considered, while in the equation for Spain, the absence of significance of this variable indicates that Spain has no competitors or complements among the Mediterranean destinations. Another puzzling matter is the following: how can one relate the negative sign of the coefficient estimate for the first difference of the dependent variable in the equation for Portugal (Δy_{t-1}), which can be interpreted as a negative growth rate of the UK demand for tourism in Portugal, with the highly positive long-run income elasticity of the UK demand for this destination? In addition, considering the origins UK, Germany and USA, Syriopoulos' study presents Portugal as the most income elastic destination (2.80, 2.85 and 3.32, respectively) among all Mediterranean countries. This is hardly credible, particularly in the case of the USA, given the relative obscurity of Portugal as a tourism destination in the 1960's and the political, social and economic instability of this country in the 1970's which covers more than two-thirds of Syriopoulos' data sample.

[28] In table 8 of Syriopoulos (1995) work this elasticity estimate is incorrectly written as 2.56.

The results for Spain seem to be more plausible than those for Portugal, particularly in the short-run. The estimates for the relative and effective own-price elasticities in the short-run, seem to be of comparable magnitudes and signs to corresponding elasticities found in other studies. However, the estimate for the long-run elasticity of UK demand for Spain to own-price changes (-2.10) seems to be overvalued. For the long-run income elasticity for Spain, a value of unity was tested and could not be rejected.

Some of the above questions could be partially answered if we take into consideration the definition of the error correction term in Syriopoulos' (1995) model. This definition implies that deviations of the UK tourism expenditure from its long-run equilibrium path are 'corrected' by a mechanism that only takes into account adjustments in the income variable. Hence, the error-correction mechanism adjusting the short-run values to the long-run equilibrium may be mis-represented, leading to the under-estimation of the adjustment velocities and hence, to the over-estimation of the long-run effects.

Song *et al.*'s (2000) study focuses on the examination of the UK tourism demand for twelve destinations including France and Spain, during the period 1965-1994, using pre-modelling time series integration analysis and the general-to-specific approach to identify the long-run equilibrium and associated short-run error-correction mechanism with the Engel-Granger (1987) two-stage method. Besides the usual right-hand side variables, this study constructs and includes an additional explanatory variable, "*destination preference index*", which measures "*non-economic preferences towards the destinations*".

The dependent variable in these models is total holiday visits per capita which prevents us from formally comparing these estimation results with those obtained in the previous section. Nevertheless, the analysis of Song *et al.*'s study provides increased knowledge about both the UK demand for tourism behaviour and the use of cointegration analysis techniques applied to tourism demand models. Furthermore, some comparison can be established if, although bearing in mind that Song *et al.*'s elasticity estimates are related to the number of tourists while our models' elasticity estimates are related to tourism expenditure, we consider number of tourists as a rough proxy for tourism expenditure.[29]

[29] Tourism expenditure, rather than tourist numbers, is generally considered as a more appropriate measure of tourism demand since arrivals may increase or decrease at a different pace from that of expenditure or may even increase (decrease) while real tourism expenditure decreases (increases) due to lower (higher) effective prices, lower (higher) spending propensity, or lower (higher) average length of stay of the incoming tourists. Therefore, the interpretation of tourist arrivals as a proxy for tourism expenditure may be misleading.

According to the authors, the ADF test suggests that all time-series variables included in the UK demand for tourism equations are I(1). The cointegration tests performed on the 12 equations indicate cointegration in eleven of them but fail to do so in the equation for France. This means that the statistical inference based on the estimation results of this equation may not be valid since the long-run relationship between the UK demand for tourism in France and its determinants may be spurious. For a group of six (including Spain) out of the twelve equations, both long- and short-run models pass all the diagnostic tests for residuals normality, heteroscedasticity and autocorrelation, functional form and structural stability. Not surprisingly, the equation for France fails several of them.

In table 4.11 we present the estimation results obtained by Song *et al.* (2000) for the long- and short-run models of the UK demand for France and Spain.

Table 4.11: Estimation of the long-run and short-run models for the UK tourism demand in France and Spain

	FRANCE		SPAIN	
	Long-run	Short-run ECM	Long-run	Short-run ECM
$PF_t(\Delta)$	-1.079 (-3.14)	-0.778 (-2.32)		
$PS_t(\Delta)$			-0.496 (-2.54)	-0.491 (-2.44)
$PIR_t(\Delta)$	0.951 (4.13)	0.785 (2.39)	1.372 (4.73)	1.047 (3.39)
$PIT_t(\Delta)$			-0.67 (-3.16)	-0.329 (-1.25)
$I_t(\Delta)$	2.123 (11.23)	1.665 (2.12)	2.199 (19.81)	2.770 (3.68)
PREF (Δ)	0.818 (8.80)	0.904 (10.89)	0.791 (11.63)	0.919 (10.00)
EQE_{t-1}		-0.286 (-1.26)		-0.899 (-4.16)
R^2	0.985	0.881	0.981	0.855
DW	1.08	1.82	1.86	1.86
ADF(0)	-3.03		-4.87	

Source: Song et al. (2000), appendix A, pp. 621-24

The variables' notation is modified to facilitate interpretation and comparison of results. Therefore, PF, PS, PIR, and PIT represent, respectively, the price of tourism in France, Spain, Ireland and Italy; I stands for UK real per capita income, PREF is the preference index and EQE represents the residuals series obtained from the OLS estimation of the long-run equilibrium model. In the first column of table 4.11 we present the names of the explanatory variables which, when in levels, represent the relevant variables for the long-run equations and, when in first differences (Δ), represent the relevant variables for the short-run/error-correction (ECM) equations. The last three rows of the table present the R^2 and DW statistics as well as the t-statistic values for the cointegration test run as the ADF(0) unit root test for the EQE_{t-1} series of residuals. The values between brackets beneath the coefficients' estimates represent the respective t-ratios.[30]

The authors claim that the results of the ADF(0) test indicate absence of cointegration in the equation for France, which means that the long-run relationship between the UK demand for tourism and its assumed determinants may not exist in any meaningful way and statistical inference and forecasting procedures may be invalid for this equation. Therefore, we will focus on the estimation results of the equation for Spain.

Besides passing all diagnostic tests as well as being cointegrated, a postulated long-run specification must be data-coherent and theoretically-consistent to be a reliable representation of the underlying long-run relationship between the dependent variable and its assumed determinants. However, some of the estimation results reported in Song *et al.* seem difficult to associate with a plausible economic interpretation. For instance, it is hardly believable that Ireland should be such a 'strong' substitute for tourism in Spain, both in the short- and in the long-run, whereas Italy is a 'weak' complement and only in the long-run. Moreover, the estimation results reported indicate Ireland as a significant substitute for all 12 destinations considered, while Spain is presented as a significant complement of such destinations as Austria, Germany, Nederland and the USA in UK tourists' preferences. Otherwise, the estimation results for Spain seem consistent with those obtained in our study, particularly relating to the high speed of adjustment of UK tourists to changes in their demand, the inelastic own-price elasticity and the elastic income elasticity, both in the long- and in the short-run.

[30] As pointed out in Engel and Granger (1987), the t-ratio of individual significance, as any other testing procedure for the first-step equation coefficients obtained with the Engel-Granger method, are not valid means for statistical inference. Indeed, because the long-run (first-step) equation is assumed to be a steady-state equilibrium, a dynamic mis-specification may exist, invalidating inference.

4.5. CONCLUSION

In chapter 3, we concluded that the static single equation approach for modelling tourism demand was not the appropriate means for undertaking UK tourism demand analysis. This approach lacks both the theoretical basis upon which plausible behavioural features can be explicitly modelled and interpreted, and the empirical methodology with which a sound specification, and valid estimation and testing procedures can be implemented. In static single tourism demand equations, the essential features usually omitted are, on the one hand, a dynamic structure explicitly incorporating the process by which short-run consumer behaviour adjusts to its long-run equilibrium and, on the other hand, the formal testing of utility theory restrictions. Furthermore, within a single equation framework, the existence of interdependencies among competing destinations may be indicated but not formally identified or tested, since such interdependencies need to be specified within a system of equations configuration. Hence, following the inconsistent results obtained in chapter 3 (a wide range of elasticities estimates, obtained by introducing minor changes in the models, associated with unexpected signs and implausible magnitudes for some of the coefficients' estimates and poor statistical performance of the models), our task for the remainder of this study was the construction of alternative empirical models which could overcome the omissions and flaws inherent in static single equation models.

In this chapter we addressed the modelling of the UK tourism demand within a dynamic single equation approach, based on the principles of economic theory and using appropriate econometric methods for the specification, estimation and evaluation of the models adopted. The construction of UK tourism demand models in this chapter followed a "data-based revision" strategy known as the general-to-specific approach. The general model was defined as an ARDL structure taking into consideration the dynamic nature of the UK tourism demand through the insertion of an error-correction mechanism.

The dynamic specification of the ARDL models assumed that, in order to maintain the steady state relationship, tourists adjust the current values of their expenditure partly in response to current changes in the explanatory variables, and partly in response to the desequilibrium observed in the previous period. A parsimonious form of the model was then derived from the general ARDL specification through the systematic testing of plausible restrictions and comparative quality evaluation of alternative models. From the parsimonious final form, we derived the equations for the long-run equilibrium relationship and the short-run error-correction model which provide information about the short- and long-run impacts on the UK tourism demand originating from changes in its determinants.

Prior to the estimation of the parsimonious ARDL models we tested the order of integration of all the time series included in the equations. The unit root tests indicated all the variables included to be I(1). Nonstationarity of time series has always been regarded as a major problem in empirical studies since it can generate spurious relationships among the levels of economic variables, giving rise to invalid estimation, inference and forecasting procedures. However, theory suggests that if a stable long-run relationship exists among the levels of economic variables, they cannot diverge too far apart from each other and from an 'attractor' which defines their steady-state equilibrium path. In this case the variables are said to be cointegrated. Cointegration means that a linear combination of variables is stationary although individually they are not. If two or more nonstationary time series can be linearly combined into a single series which is itself stationary, the original variables are said to be cointegrated. In this case, a genuine long-run relationship exists among the variables levels, and this relationship defines the long- run equilibrium path or attractor.

From the parsimonious ARDL models we derived the long-run equilibrium regressions and tested for stationarity in their residuals series using conventional unit root tests. These tests indicate stationarity of the residuals and hence, the existence of a cointegrated relationship. However, in a single equation specification, the cointegrated relationship found may not be unique. The uniqueness of a long-run equilibrium relationship can be tested using the methodology suggested by Pesaran and Shin (1995) and Pesaran *et al.* (1996). We implemented these testing procedures and the results obtained support the existence of a genuine and meaningful long-run relationship between the dependent variable (the UK demand for tourism) and its determinants. The statistical validity of the estimation results obtained from both the long-run and the short-run specifications is further confirmed by a battery of diagnostic tests which provided sufficient statistical evidence to classify the ARDL specifications as robust, structurally stable and well-defined models.

The results indicate a significant link between Spain and France, on the one hand, and Spain and Portugal on the other hand, but no significant link between France and Portugal. Price changes in Spain affect the UK tourism demand for France in the long-run but not in the short-run, while price changes in France affect the UK demand for Spain both in the long- and in the short-run. The magnitudes and signs of this interrelationship qualify

France as a stronger long-run price-competitor relative to Spain, while Spain seems to be a stronger short-run price-competitor relative to France. Therefore, we can conclude that France and Spain share a strong link of dependence concerning UK tourism demand and, further, that this dependence is bi-directional.

In the long-run, the UK demand for tourism is more sensitive to own-price variations in France than in Spain or Portugal. In fact, increases in the own-price of France induce decreases in the UK tourism demand for this country three and two times greater, respectively, than those for Spain and Portugal. In the long-run, France's gains from increases in UK per capita income are larger than those expected for Spain or Portugal while in the short-run, Portugal benefits more from these increases.

Price changes in Portugal do not significantly affect the UK demand for tourism in Spain while price changes in Spain do have a significant impact on the UK demand for Portugal. This means that Portugal is not a significant price-competitor of Spain while Spain is a significant price-competitor of Portugal. In addition, increases in the UK tourism expenditure in Portugal affect the UK tourism demand for Spain significantly and positively, although in a modest magnitude, both in the long- and in the short-run. This means that the more UK tourists spend in Portugal the more they tend to spend in Spain. Therefore, the UK tourists' preference for Spain relative to Portugal, is apparent and it seems that Portugal, as a tourism destination, depends more on what happens to the UK demand for Spain, than Spain depends on Portugal. This reinforces our previous expectations about the existence of a strong link between Spain and Portugal and, further, indicates its unidirectional nature.

The above considerations permit us to qualify the empirical specifications used in this chapter and the methodologies employed in their derivation, as sound and reliable means for analysing the UK demand for tourism in France, Spain and Portugal. The estimation results obtained allow for a theoretically consistent and data-plausible interpretation of reality within a simplified context, permitting the consolidation of empirical knowledge about the UK tourists' behaviour both in the short- and in the long-run. The models estimated in this chapter fulfil the quality criteria of good empirical models: they are consistent with the underlying economic theory, compatible with the empirical data, simple, relevant, plausible, accurate, reliable, and useful. Their forecasting ability will be addressed in chapter 8.

Nevertheless, the equation-by-equation nature of the models derived in this chapter disallows the formal testing of utility theory restrictions involving cross-equation coefficients. Indeed, although they suggest the nature and direction of the destinations' competing conduct, the ARDL specifications do not allow for the formal testing of these hypotheses. Therefore, the structure of the model must change to embrace a system of equations within which these features can be included.

CHAPTER 5

MODELLING THE UK TOURISM DEMAND WITHIN A STATIC SYSTEM OF EQUATIONS FRAMEWORK

5.1 INTRODUCTION

In chapter 3, we followed the traditional approach of tourism demand analysis using a static single equation framework to specify empirical models. In chapter 4, we adopted a more recent methodology in the estimation of dynamic ARDL models and associated error-correction equations for the UK tourism demand in France, Spain and Portugal. These models have the advantage of explicit treatment of the dynamics of demand behaviour, and allow for improved econometric methods in the specification and testing of single equations. However, a comprehensive analysis of tourists' demand behaviour requires the consideration of theoretical economic principles within which tourism demand functions should be specified. Indeed, as we move on to consider complete systems of equations, theory becomes directly relevant at all stages of the econometric analysis: specification, estimation, and statistical inference. As Deaton (1986, p.1768) points out "*it is not possible to study applied demand analysis without keeping statistics and economic theory simultaneously in view*". This view is reflected in a number of empirical studies which explicitly base the derivation of their models on the principles of consumers' behaviour, using what has become known as 'flexible' functional forms to specify a theory-consistent system of demand equations.

Since Stone's (1954) first estimation of a demand system explicitly derived from consumer theory, there has been a continuing search for alternative specifications which can be considered good approximations to consumers' preference structure. Many models have been proposed, but theoretically consistent models which allow for general income and price effects and also nest simpler specifications are few. Indeed, the choice of an appropriate functional form to mimic the true data generating process has been the centre of intense debate among investigators for the last three decades.

One of the earlier specifications frequently used to test consumer theory hypothesis, was first proposed by Theil (1965) and Barten (1968, 1969) and is known as the Rotterdam model. This model is similar to that of Stone but uses differenced variables instead of their levels. As mentioned in chapter 4, it is generally recognised that models using variables in differences are not the best way to estimate long-run equilibrium relationships, which is one of the main interests of this study. In addition, applications of the Rotterdam-type model tend to supply results which contradict the basic assumptions of consumer behaviour theory due, perhaps, to being based on "*an unhappy approximation or choice of functional form*" (Deaton and Muellbauer, 1980b, p. 72).

In more recent studies many other 'flexible' functional forms have been used to approximate consumers' indirect utility or cost functions. Although Lewbel (1987, 1990, 1991, 1995) groups these forms in terms of fractional demand systems, full rank demand systems and Engel curve approximations, there are other classification possibilities that appear to describe their different properties in a much clearer way. For instance, Fisher *et al.* (2001) divides the flexible functional forms into three major sub-groups: locally flexible functional forms, which include the translog models of Christensen et al. (1975), and Jorgensen et al. (1982), the AIDS specification of Deaton and Muellbauer (1980a, 1980b) and the generalised Leontief system of Caves and Christensen (1980); globally regular functional forms, which include the miniflex Laurent models discussed in Barnett et al. (1985, 1987), the general exponential model of Cooper and McLaren (1996) and the quadratic AIDS model of Banks et al. (1997); and asymptotically globally flexible forms, which include the Fourier flexible model discussed in Chalfant and Gallant (1985) and the asymptotically ideal model of Barnett and Yue (1998).

The locally flexible functional forms have been recognised to have some limitations associated with the regularity regions they are able to define. For instance, Caves and Christensen (1980) show that when preferences are not homothetic and substitution increases among goods, the generalised Leontief (GL) system has a rather small regularity region, and the Monte Carlo analysis of Guilkey and Lovell (1980) shows that the GL and the translog (TL) models fail to provide a satisfactory approximation of the true data generating process.

Another problem with the TL model is that it can classify goods as complements when they are actually substitutes (Fisher *et al.* 2001). The AIDS model is a widely used flexible form derived from a PIGLOG (price independent generalised logarithm) expenditure function which may have a small regularity region. However, Chalfant (1987) using Fourier series and Ramajo (1994) using Laurent series show that the regularity region of the AIDS model can be considerably increased.

A partial solution to the problem of small regularity regions has been provided by the development of flexible forms with larger theoretical regularity regions and higher rank models that can approximate better more general Engel curves. These are globally regular functional forms as discussed, for example, in Cooper and McLaren (1996). The rank of a demand system has implications for non-linear Engel curves. Some Engel curves can be more non-linear than the rank 2 AIDS and translog models account for. In these cases, higher rank models, such as the quadratic AIDS (QUAIDS) of Banks et al. (1997) which is a rank 3 demand system by extension of the AIDS model, can approximate better these non-linear Engel curves. However, as noted in Fisher *et al.* (2001), these models may fail to provide an effective approximation of the curvature (derivatives) of the true utility or cost functions. Asymptotic globally flexible forms such as the Fourier flexible form (FFF) and the asymptotic ideal model (AIM), also known as semi-non-parametric functions, can provide asymptotically global approximations for more complex data generating processes as well as for their partial derivatives.

It should be noted, however, that no matter what degree of flexibility is adopted, there is always the possibility of regularity conditions being rejected, and more flexible form than the locally flexible are not immune to this problem. This is so not only because the data generally bear a "heavy load of assumptions" independent of the technique used, but also because violation of regularity can be due to such factors as aggregation bias and mis-specified dynamics which no functional form overcomes by itself. In addition, given a specific data set, the flexible functional forms may differ substantially in their approximation properties. Indeed, as the true data generating process underlying specific economic relationships is unknown, we agree with Deaton and Muellbauer (1980b) in that "*there is no reason to believe that utility functions are exactly translogarithmic* (or any other pre-defined functional form, for that matter) *either at the individual or aggregate levels*". At most, functional forms are approximations and, like all approximations, can be more or less accurate. Therefore, it should be emphasised that while the semi-non-parametric forms may have more interesting asymptotic properties than the parametric functions, it does not follow that one is to be preferred to the other given specific data sets.

Recent findings in the econometrics literature, state that estimation and statistical inference crucially depend on the integration and cointegration properties of the time series involved. For example, Attfield (1997) shows that with the linear share equations of an AIDS system, homogeneity cannot be rejected once the cointegrating properties of the time series are imposed in the estimation. This implies that, in the spirit of Engel and Granger (1987), testing for linear cointegration and deriving linear forms for the underlying error-correction mechanism is essential. Except for the AIDS model, all flexible functional forms referred above have non-linear share equations and, as pointed out by Granger (1995), the non-linear modelling of non-stationary variables is a new, complex and largely tentative subject. There is a trade-off between estimating demand with linear or non-linear systems: if the data generating process is non-linear, the estimates obtained with a linear approximation may be biased; if the variables are cointegrated, the results obtained from non-linear models may be imprecise (Fisher et al., 2001, p.67). As will be shown in chapter 7, the variables involved in the estimation of the UK tourism demand for France, Spain and Portugal are cointegrated. Therefore, the AIDS model appears to be a good choice to approximate the UK tourist's preference structure.

The AIDS model is seen as a particularly convenient specification with attractive features which, with appropriate transformations and restrictions, can be made to nest a variety of models (see, for example, Anderson and Blundell 1983, 1984 and chapter 6). The AIDS model gives an arbitrary first-order approximation to any demand system, satisfies the axioms of preferences exactly, permits perfect aggregation over consumers, allows for simple linear estimation methods and the imposition of linear restrictions to test homogeneity and symmetry. Furthermore, early results obtained with the AIDS specification showing serial correlation problems and rejecting the hypothesis of demand theory, have been qualified as possible shortcomings due to the static nature of the orthodox model, non-stationarity of the variables or the result of income effects arising from aggregation. Indeed, the results of more recent empirical applications with aggregate models, capturing previously omitted characteristics of consumers' behaviour by the introduction of dynamic adjustments or trend terms, tend to produce results consistent with all consumer demand postulates.

The demand theory conceives an individual consumer endowed with a fixed income and facing a market of commodities with given prices. The consumer's objective is to allocate his/her total expenditure to a specific bundle of goods in order to maximise the utility function that reflects his/her preferences' order. The utility maximisation hypothesis allows for the derivation of demand functions relating quantities purchased to given prices and expenditure.

Deriving demand equations systems from consumer behaviour theory has advantages. In particular, it permits the inclusion of a number of theoretical constraints on the parameters of the equations, leading to parsimonious and efficient models as more *a priori* knowledge can be incorporated via the imposition of such restrictions. It also assures consistency between each equation and the total expenditure providing models that can be used in aggregate terms.[1] In addition, an essential feature of consumer demand theory – the interdependence of related commodities – can also be embodied within these systems through the imposition of cross-equation restrictions. Indeed, "*the essence of the consumer demand systems approach consists in providing empirical demand analysis with a conceptual framework to deal with the interdependencies of demand for various commodities*" (Barten, 1977, p. 57).

Although most investigators concerned with tourism demand analysis have chosen a single equation approach for specifying their empirical models others, such as White (1982), O'Hagan and Harrison (1984), Syriopoulos and Sinclair (1993) Papatheodorou (1999) and De Mello *et al* (2001), linked their empirical specifications to theoretical principles integrated in consumer utility theory, using system of equations based on the AIDS model. Rather than the *ad hoc* reasoning that tends to underlie the static single equation approach, the AIDS model gives a first-order approximation to any demand system, satisfies the axioms of choice, permits perfect aggregation over consumers without imposing parallel Engel curves, is data-consistent, simple to specify and easy to estimate. Therefore, the AIDS system permits not only the estimation of the complete set of relevant

[1] Aggregate data have relatively less income variation than disaggregated data. Low income variations and relatively large price variations, often found in aggregate data, may favour functions with simple Engel curves and/or low rank such as the AIDS model. In contrast, functions having more income flexibility relative to price flexibility such as the QUAIDS model, may be preferred when using household-level data.

elasticities which supply crucial information on the interdependencies of competing products, but also the formal testing of the consumer theory assumptions.

In this chapter, a version of the 'orthodox' static AIDS model is applied to the analysis of the UK demand for tourism in France, Spain and Portugal with the introduction of two innovations: the concept of neighbourhood between the origin and the destinations and the inclusion of a possible non-constancy of the real expenditure coefficient during the period under analysis. The former is believed to be a relevant factor in the explanation of the competitive behaviour of the destination countries, as complementarity or substitutability, shown by the signs of the cross-price elasticities, is of particular relevance in this context. The latter is based on the belief that tourists' allocation of expenditure may change over time due to factors that modify the political and economic relationships between the origin and the destinations and among the destinations themselves. This is particularly interesting in the cases of Portugal and Spain, as these countries experienced a transition from 'developing' economies at the beginning of the sample period (1969), to a 'developed' status at its end (1997). The consideration of France as a neighbouring destination allows for the comparison of tourism demand behaviour between a developed country and its poorer neighbours and the examination of the extent to which this behaviour becomes more or less similar over time, given changes in its determinants.

This chapter proceeds as follows. Section 5.2 provides a summarised derivation of the AIDS model, its assumptions and main features. Section 5.3 presents the static AIDS model of the UK demand for tourism in France, Spain and Portugal and a description of the variables included. Section 5.4 reports the estimation results obtained and their interpretation. A comparison between these results and those obtained in other studies applying the AIDS model to tourism demand analysis is also addressed. Section 5.5 concludes.

5.2. THEORETICAL FEATURES OF THE AIDS MODEL[2]

The AIDS model can be viewed as an extension (including price effects) of the Working-Leser model which relates the budget shares w_i to the logarithm of the total

[2] The derivation of the AIDS model in this section follows that of Deaton and Muellbauer (1980a, 1980b)

expenditure x, such that $w_i = \alpha_i + \beta_i \log x$. The model rests upon a specific class of preferences – the PIGLOG class – which are represented via a cost or expenditure function defining the minimum expenditure necessary to attain a specific level of utility at given prices. As Muellbauer (1976) shows, the PIGLOG class permits an exact aggregation over consumers without imposing identical preferences.

Let x be the exogenous budget or total expenditure which is to be spent within a given period on some or all of n products. These products can be bought in non-negative quantities q_i at given prices p_i, i=1, ..., n. Let $\mathbf{q} = (q_1, q_2, ..., q_n)$ be the quantities vector of the n products purchased, and $\mathbf{p} = (p_1, p_2, ..., p_n)$ the prices vector.

The budget constraint of the representative consumer is $\sum_{i=1}^{n} p_i q_i = x$. Defining the utility function as $u(\mathbf{q})$, the consumer's aim is to maximise the utility subject to the budget constraint:

$$\max u(\mathbf{q}) \text{ subject to } \sum_{i=1}^{n} p_i q_i = x \tag{5.1}$$

The solution of this maximisation problem leads to the Marshallian (uncompensated) demand functions $q_i = g_i(\mathbf{p}, x)$.

Alternatively, the consumer's problem can be defined as the minimum total expenditure necessary to attain a specific level of utility u*, at given prices:

$$\min \sum_{i=1}^{n} p_i q_i = x \text{ subject to } u(\mathbf{q}) = u^* \tag{5.2}$$

The solution of this minimisation problem leads to the Hicksian (compensated) demand functions $q_i = h_i(\mathbf{p}, u)$. Therefore, a cost function can be defined as

$$C(\mathbf{p}, u) = \sum_{i=1}^{n} p_i h_i(\mathbf{p}, u) = x \tag{5.3}$$

Given the total expenditure x and prices **p**, the utility level u* is derived from the solution of the problem stated in equation (5.1). Solving (5.3) for u, an indirect utility function is obtained such that:

$$u = v(\mathbf{p}, x) \tag{5.4}$$

Therefore, the AIDS model specifies a cost function which is used to derive the demand functions for the commodities under analysis. The process of derivation can be summarised in the following three steps:

First, $\frac{\partial C(p,u)}{\partial p_i} = h_i(p,u)$ is derived establishing the Hicksian demand functions.

Second, solving (5.3) for u, the indirect utility function is obtained, such that $u = v(p,x)$.

Third, $h_i[p, v(p,x)] = g_i(p,x)$ is retrieved stating the Hicksian and the Marshallian demand functions as equivalent.

The Hicksian and Marshallian demand functions have the following properties:

1. Adding-up: $\sum_i p_i h_i(p,u) = \sum_i p_i g_i(p,x) = x$, meaning that all budget shares sum to unity;

2. Homogeneity: $h_i(p,u) = h_i(\theta p,u) = g_i(p,x) = g_i(\theta p,\theta x) \quad \forall \theta > 0$, meaning a proportional change in all prices and expenditure has no effect on the quantities purchased;

3. Symmetry: $\frac{\partial h_i(p,u)}{\partial p_j} = \frac{\partial h_j(p,u)}{\partial p_i}$, $\forall i \neq j$, meaning that consumer's choices are consistent;

4. Negativity: The n-by-n matrix of elements $\frac{\partial h_i(p,u)}{\partial p_j}$ is negative semidefinite, that is, for any n vector ξ, the quadratic form $\sum_i \sum_j \xi_i \xi_j \frac{\partial h_i(p,u)}{\partial p_j} \leq 0$. This means that a rise in prices results in a fall in demand as required when the commodities under analysis are normal goods.

The AIDS model specify the following cost function:

$$\ln C(p,u) = a(p) + ub(p) \tag{5.5}$$

where $a(p) = \alpha_0 + \sum_i \alpha_i \ln p_i + \frac{1}{2}\sum_i \sum_j \gamma_{ij} \ln p_i \ln p_j$ and $b(p) = \beta_0 \prod_i p_i^{\beta_i}$

The derivative of (5.5) with respect to ln p_i is:

$$\frac{\partial \ln C(p,u)}{\partial \ln p_i} = \alpha_i + \sum_j \gamma_{ij} \ln p_j + u\beta_i \beta_0 \prod_i p_i^{\beta_i} \tag{5.6}$$

As $C(p,u) = x \Leftrightarrow \ln C(p,u) = \ln x$, then

$$\ln x = a(p) + u\,b(p) \tag{5.7}$$

Solving (5.7) for u we obtain

$$u = \frac{\ln x - a(p)}{b(p)} \tag{5.8}$$

Substituting (5.8) in (5.6) we have

$$\frac{\partial \ln C(\bullet)}{\partial \ln p_i} = \frac{\partial C(\bullet)}{\partial p_i}\frac{p_i}{C(\bullet)} = h_i(\bullet)\frac{p_i}{C(\bullet)} = \frac{p_i q_i}{x} = w_i = \alpha_i + \sum_j \gamma_{ij} \ln p_j + \beta_i[\ln x - a(p)]$$

If we define a price index P such that ln P = a(p), then

$$\frac{\partial \ln C(p,u)}{\partial \ln p_i} = \alpha_i + \sum_j \gamma_{ij} \ln p_j + \beta_i[\ln x - \ln P]$$

or $$w_i = \alpha_i + \sum_j \gamma_{ij} \ln p_j + \beta_i \ln\left(\frac{x}{P}\right) \tag{5.9}$$

where $$\ln P = \alpha_0 + \sum_k \alpha_k \ln p_k + \frac{1}{2}\sum_k \sum_\ell \gamma^*_{k\ell} \ln p_k \ln p_\ell \tag{5.10}$$

equations (5.9) and (5.10) are the basic equations of the AIDS model.

In a tourism analysis context, i is a destination country among a group of n alternative destinations demanded by tourists of a specific origin. The dependent variable w_i, represents destination i share of the origin's tourism budget allocated to the set of n destinations. This share's variability is explained by the price of tourism (p) in i and in alternative destinations j and by the per capita expenditure (x) allocated to the group of n destinations, deflated by price index P. The model is based on the following assumptions:

1. All budget shares sum to unity, the adding-up restriction, requiring:

$$\sum_i \alpha_i = 1, \quad \sum_i \beta_i = 0, \quad \sum_i \gamma_{ij} = 0, \text{ for all j;}$$

2. A proportional change in all prices and expenditure has no effect on the quantities purchased, the homogeneity restriction, requiring:

$$\sum_j \gamma_{ij} = 0, \text{ for all i;}$$

3. Consumers' choices are consistent, the symmetry restriction, requiring:

$$\gamma_{ij} = \gamma_{ji}, \text{ for all i, j;}$$

4. A rise in prices results in a fall in demand the negativity restriction, which requires the condition of negative own-price elasticities for all destinations.

The restrictions imposed on α and γ comply with these assumptions and ensure that equation (5.10) defines P as a linear homogeneous function of individual prices. If prices are relatively collinear, then "P *will be approximately proportional to any appropriately defined price index, for example, the one used by Stone, the logarithm of which is* $\sum w_k \ln p_k = \ln P^*$" (Deaton & Muellbauer, 1980a, p.76)[3]. Hence, the deflator P in equation (5.10) can be substituted by the Stone price index ln P* such that,

$$\ln P^* = \sum_i w_i^B \ln p_i \tag{5.11}$$

where w_i^B is the budget share of destination i in the base year. With this simplification for the expression of P, equation (5.9) can be rewritten and estimated in the following form:

$$w_i = \alpha_i^* + \sum_j \gamma_{ij} \ln p_j + \beta_i \ln\left(\frac{x}{P^*}\right) \tag{5.12}$$

Equation (5.12) specifies a model in the linear-log form which prevents the direct interpretation of its coefficients as elasticities. However, the interpretation of the signs attached to the coefficients of the model in this form can give a preliminary indication of how the dependent variable reacts to changes in its determinants. The coefficients of the price variables (γ_{ij}) represent the absolute change in the expenditure share allocated to commodity i due to a 1% change in the price of good j, *ceteris paribus*. For i=j, the coefficient's sign of the own-price variable is expected to be negative according to the theoretical rule of negativity which qualifies i as a normal good in economic theory terminology. For i≠j, the sign of the coefficient is expected to be positive if i and j are substitutes, and negative if they are complements. The β_i coefficients represent the absolute change in the i^{th} expenditure share i, given a 1% change in real per capita expenditure, prices being held constant. Generally, in an AIDS model framework, a coefficient $\beta_i > 0$ gives rise to an expenditure elasticity above unity, and a coefficient $\beta_i < 0$ gives rise to an expenditure elasticity in the interval (0, 1). In the first case, the demand for commodity i would be expenditure (income) elastic and in the second case, expenditure (income) inelastic. In economic theory terminology, commodities with income-elastic demands are qualified as "luxuries" and commodities with income-inelastic demands

[3] If prices are not collinear, the linear approximation of the AIDS model obtained through the use of the Stone price index can bias the parameters' estimates of the budget share equations. However, the bias is likely to be more important in micro rather than in aggregate data as showed by Pashardes (1993).

are qualified as "necessities". This qualification seems to indicate that while a "luxury" good can always be given up when restrictions on income are imposed and its budget allocation can always be redistributed to other purposes, a "necessary" good is essential to the consumer and budget restrictions affect only slightly the demand for this type of good. In this sense, an income-inelastic ($\beta_i < 0$) good is likely to have few competitors while an income-elastic ($\beta_i > 0$) good is likely to have many. This qualification seems to be more adequate in a tourism demand context than the theory terminology, which would qualify a tourism destination, in the first case, as a "necessity" and in the second, as a "luxury".

The model assumes separability between consumption and labour supply, and excludes quantity constraints. It also assumes that consumers allocate their budget to groups of commodities in a multi-stage budgeting process, implying that each group of preferences is independent from others. Thus, it is assumed that the expenditure allocated by UK tourists to France, Spain and Portugal is separable from expenditure allocated to other destinations and that the decision to spend money in those countries is made in several stages. First, UK tourists allocate their consumption budget to tourism and other goods; then to tourism in their southern neighbouring countries and other parts of the world; finally they decide between France, Spain and Portugal. The AIDS model is applied to this last stage of expenditure allocation.

5.3. THE STATIC AIDS MODEL OF THE UK DEMAND FOR TOURISM

The UK tourism demand for France, Spain and Portugal is estimated using an AIDS model allowing for possible non-constancy of the real expenditure coefficient over the sample period. Changes in this coefficient can occur if tourists' allocation of expenditure is affected, over time, by factors that modify the political and economic relationships between origin and destinations and among the destinations themselves. The model is as follows:

$$\begin{cases} WP = \alpha_P + \gamma_{PP}PP + \gamma_{PS}PS + \gamma_{PF}PF + \beta_P E + \beta'_P\,[SE] + \delta_P D + 0_P T + u_P \\ WS = \alpha_S + \gamma_{SP}PP + \gamma_{SS}PS + \gamma_{SF}PF + \beta_S E + \beta'_S\,[SE] + \delta_S D + 0_S T + u_S \\ WF = \alpha_F + \gamma_{FP}PP + \gamma_{FS}PS + \gamma_{FF}PF + \beta_F E + \beta'_F\,[SE] + \delta_F D + 0_F T + u_F \end{cases} \quad (5.13)$$

The dependent variables are each country's share of the UK tourism budget allocated to these three countries, denoted as Wi, where i = F (France), S (Spain) and P (Portugal). The independent variables are the effective prices of tourism in each country (Pi), the UK real expenditure allocated to all destinations per capita of the UK population (E), a dummy variable D and a trend variable T. The effective price of tourism in country i and the UK real per capita expenditure are defined as follows:

$$Pi = \ln\left(\frac{CPI_i/CPI_{UK}}{R_i}\right) \quad \text{and} \quad E = \ln\left(\frac{\sum_i E_i / UKP}{P^*}\right)$$

where CPI_i is the consumer price index of destination i, CPI_{UK} is the consumer price index of the UK, R_i is the exchange rate between country i and the UK, E_i is the nominal UK tourism expenditure in country i, UKP is the UK population and $\ln P^*$ is the Stone price index defined in equation (5.11). The data for UK tourism expenditure, disaggregated by destinations and measured in million pounds sterling, were obtained from one common source, the *Business Monitor MA6* (1970-1993), continued as *Travel Trends* (1995-1998). Data on the UK population, price indexes and exchange rates were obtained from the *International Financial Statistics* (IMF) *Yearbooks* (1984, 1990 and 1998).

The dummy variable D accounts for several events that appear to have influenced the UK tourism demand in France, Spain and Portugal, during the period 1974-1981. The first was the political turmoil that followed the Portuguese revolution in April 1974. This event had a substantial negative effect on the UK tourism flows to Portugal, which is believed to have lasted, at least, until 1979. Second, Spain was affected by both the events in Portugal and its own political changes, involving the substitution of a dictatorial regime of forty years by a parliamentary democracy. Third, additional events which had adverse effects on the demand for tourism all over the world and which particularly affected weaker and unstable economies like those of Portugal and Spain were the oil crises in 1973 and 1979.

Ideally, the model would include three different dummy variables portraying the events that affected demand in each of the destination countries. However, this would mean the estimation of additional parameters, reducing further the already few degrees of freedom available. Tests performed on an equation-by-equation basis show that the different periods to take into consideration are very close to each other (1974-1979 for Portugal, 1975-1980 for

Spain and 1975-1981 for France). Hence, one single dummy covering the period 1974-1981 is included in the model. The coefficients of the dummy variable are expected to be negative in the equations for Portugal and Spain and positive in the equation for France.

A structural break in the influence of the UK real per capita expenditure on the dependent variables separates the sample period into two sub-periods, 1969-79 and 1980-97. Important events that contributed to the structural break are Spain's membership of EFTA in 1980 and Spain and Portugal's negotiations for EC membership which started, for Portugal, in October 1978 and for Spain in mid 1979, marking the turning point from isolation to partnership in one of the biggest potential markets in the world.

This view can be illustrated by the statement of José da Silva Lopes (1996, p.136): "*One of the most important consequences of these negotiations was the ending of a secular separation between the Spanish and the Portuguese economies. From an economic point of view the two countries had, traditionally, lived back to back. Before membership to the EC, both countries allowed more trading concessions to any other country then to its neighbour. In 1973 Spain tried to negotiate with Portugal the creation of an Iberian common market, but Portugal declined. The Portuguese authorities believed that the differences in dimension and strength between the two economies were such that their integration would only be admissible in a vast multilateral framework where other influences and powers could also be present.*" Events of such importance affecting Spain and Portugal inevitably affected France and, hence, the distribution of expenditure among the countries. Thus, a dummy variable S, assuming the value 1 in the years 1980-1997 and 0 otherwise, is multiplied by the variable E to account for the structural break detected. This new variable is denoted as SE in system (5.13).

A proxy for possible omitted dynamics, demographic shifts and deterministic non-stationarity was added in the form of a trend variable. A trend variable may also account for changes in tastes of UK tourists - patterns of behaviour such as 'country i-addicted' (giving rise to a positive trend coefficient) or 'cultural curiosity' (resulting in a negative trend coefficient) - not taken into account by other explanatory variables. Tourists may adhere to the already known or may display more adventurous behaviour by visiting other countries, even when confronted with unchanged expenditure and prices in destination countries already visited. If this is the case and a trend variable is not included, the estimates of prices and expenditure coefficients may be spurious and the estimation results may present signs of autocorrelation.

The individual and joint statistical significance of the variables D, SE and T where tested on an equation-by-equation and a system basis. These hypotheses were tested by using the likelihood ratio (LR) and the F statistics for the single equations, and the Wald statistic for the system of equations. The LR and Wald statistics are asymptotically distributed as a Qui-square with degrees of freedom equal to the number of restrictions imposed. The results of these tests are presented in table 5.1 (p values in brackets). In all tests, the explanatory variables included in the unrestricted (U) and restricted (R) models are displayed.

Table 5.1: Tests for the individual and joint significance of variables D, SE and T

Equations	Hypothesis in test	Single equation: LR statistic	Single equation: F statistic	System of equations: Wald statistic
France	**Insignificance of D**	$\chi^2(1)$=13.09 (0.00)	F(1, 23)=13.12(0.00)	$\chi^2(2)$=42.47 (0.00)
Spain	U: PP, PS, PF, E, D	$\chi^2(1)$=4.92 (0.03)	F(1, 23)=4.25 (0.05)	
Portugal	R: PP, PS, PF, E	$\chi^2(1)$=18.74 (0.00)	F(1, 23)=20.90(0.00)	
France	**Insignificance of SE**	$\chi^2(1)$=5.59 (0.02)	F(1, 22)=4.68 (0.04)	$\chi^2(2)$=5.36 (0.07)
Spain	U: PP, PS, PF, E, D, SE	$\chi^2(1)$=6.30 (0.01)	F(1, 22)=5.34 (0.03)	
Portugal	R: PP, PS, PF, E, D	$\chi^2(1)$=1.51 (0.22)	F(1, 22)=1.18 (0.29)	
France	**Insignificance of T**	$\chi^2(1)$=5.71 (0.02)	F(1, 21)=4.57 (0.04)	$\chi^2(2)$=5.36 (0.07)
Spain	U: PP, PS, PF, E, D, SE, T	$\chi^2(1)$=6.39 (0.01)	F(1, 21)=5.18 (0.03)	
Portugal	R: PP, PS, PF, E, D, SE	$\chi^2(1)$=1.06 (0.30)	F(1, 21)=0.78 (0.39)	
France	**Insignificance of D, SE,T**	$\chi^2(1)$=24.39 (0.00)	F(3, 21)=9.23 (0.00)	$\chi^2(6)$=52.01 (0.00)
Spain	U: PP, PS, PF, E, D, SE, T	$\chi^2(1)$=17.61 (0.00)	F(3, 21)=5.85 (0.00)	
Portugal	R: PP, PS, PF, E	$\chi^2(1)$=21.31(0.00)	F(3, 21)=7.60 (0.00)	

The results of table 5.1 show that the variables SE and T are highly significant in the equations for France and Spain but not significant in the equation for Portugal. Considering all the equations as a system, the orthodox AIDS model, involving only prices and per capita expenditure as regressors, is rejected against the 'unorthodox' model adopted, involving the additional explanatory variables D, SE and T.

The adding-up restriction is incorporated in the system by suppressing one of the equations from the estimation. The estimation results are invariant irrespective of which equation is excluded, and the coefficients of the deleted equation can be recovered from the other equations' estimates by applying the adding-up property. The homogeneity and symmetry restrictions are tested by imposing the required linear constraints on the appropriate parameters. The homogeneity- and symmetry-restricted system is estimated using Zellner's (1962) seemingly unrelated regressions (SUR) method.[4]

An implicit assumption in most of the literature concerning time series regression analysis is that such data are stationary. Nonstationarity of time series has always been regarded as a problem in econometric studies since it can give rise to spurious relationships among the levels of economic variables. A common practice for avoiding the problem of spurious association between variables in most regression analysis involving time-series is the introduction of a trend variable.[5] Although this procedure is not free of criticism under the scrutiny of recent theoretical work on time-series analysis, in some cases it can make the regression coefficients reflect the true association between the dependent and the explanatory variables of an econometric model. As shown in the previous chapter, the appropriate way of approaching the problem of nonstationarity in time series data is cointegration analysis, which investigates the existence of meaningful long-run relationships among economic variables.

The stationarity of the time-series variables representing tourism shares of the three destinations (WF, WS, WP), tourism prices in France, Spain and Portugal (PP, PS, PF) and the UK per capita real expenditure (E) was analysed using the Dickey-Fuller and augmented Dickey-Fuller tests. We found all the variables in levels to be integrated of order one, or I(1), and their first differences to be integrated of order zero, or I(0). Therefore, the variables' levels are nonstationary and their first differences are stationary. In this chapter, we assume that the variables involved in the AIDS model are cointegrated and that the equations in system (5.13)

[4] In the case of the unrestricted model, the OLS and Zellner's estimators are equivalent since the vector of independent variables is identical in all equations. In the case of the restricted model, Zellner's estimator is more efficient than the OLS estimator. The efficiency gain is directly related to the correlation between the different equations' disturbance terms and between the sets of independent variables.
[5] See, for example, Gujarati (1995), pp. 240-1.

represent meaningful long-run equilibrium relationship between the dependent variables and their determinants.[6] Although we do not yet have statistical proof of cointegrated equations in the AIDS model, the estimation results presented in the next section already indicate system (5.13) as a theoretically consistent, data-plausible and statistically robust econometric model, endorsing the presence of a genuine steady-state relationships.

5.4. EMPIRICAL RESULTS AND THEIR INTERPRETATION

The unrestricted system is estimated equation-by-equation using OLS. The symmetry and homogeneity restrictions are tested by imposing the required linear constraints on to the appropriate parameters. The restricted system is estimated with Zellner's (1962) SUR method for consistency with economic theory assumptions upon which the model is based. The Wald test of homogeneity, symmetry and homogeneity and symmetry simultaneously, provides χ^2 statistic values of, respectively, 2.885, 0.744 and 4.080 which lie well below the respective 5% critical values of the χ^2 distribution with two, one and three degrees of freedom (5.99, 3.84 and 7.81). This implies the non-rejection of the hypotheses under consideration.

5.4.1. ANALYSIS AND INTERPRETATION OF THE ESTIMATION RESULTS

Table 5.2 shows the coefficients' estimates and asymptotic t-values in brackets of the share equations for Portugal (PT), Spain (SP) and France (FR) using three versions of the model: unrestricted, under homogeneity and symmetry (H+S) and under homogeneity and symmetry plus the additional restriction of null cross-price effects between the share equations of Portugal and France $(H+S)^0$. The symbols $^\bullet$ and $^\circ$ indicates the 1% and 5% significance level, respectively. Table 5.3 shows the Lagrange Multiplier (LM) version and the F version of diagnostic tests for serial correlation, functional form, error normality and heteroscedasticity for the models estimated. The LM version statistic follows a Qui-square distribution and the F version follows a standard F distribution. The p values of these statistics are given in brackets.

[6] In chapter 7, we show the non-stationarity tests for the time series involved and provide statistical evidence of cointegration among the variables of the AIDS system.

Table 5.2 UK Tourism Demand (Restricted Model under Homogeneity and Symmetry)

	AIDS Model	α_i	γ_{iP}	γ_{iS}	γ_{iF}	β_i	β'_i	δ_I	θ_i
PT	Unrestricted	0.092 (5.17)•	-0.065 (-2.23)°	0.073 (2.39)°	-0.018 (-0.51)	-0.010 (-0.81)	0.005 (1.10)	-0.017 (-3.60)•	0.001 (0.88)
	Restricted (H+S)	0.091 (6.01)•	-0.072 (-2.99)•	0.072 (2.95)•	-0.0004 (-0.02)	-0.010 (-0.99)	0.006 (1.73)	-0.018 (-4.28)•	0.001 (1.17)
	Restricted $(H+S)^0$	0.090 (6.62)•	-0.072 (-3.51)•	0.072 (3.51)•	0 (none)	-0.010 (-0.99)	0.006 (1.74)	-0.018 (-4.28)•	0.001 (1.25)
SP	Unrestricted	0.412 (7.44)•	-0.068 (-0.07)	-0.375 (-3.92)•	0.526 (4.75)•	0.111 (2.90)•	-0.035 (-2.57)•	-0.053 (3.52)•	-0.011 (-2.28)°
	Restricted (H+S)	0.387 (7.86)•	0.072 (2.95)•	-0.453 (-5.62)•	0.381 (5.19)•	0.127 (4.26)•	-0.048 (-4.45)•	-0.049 (-3.53)•	-0.011 (-3.85)•
	Restricted $(H+S)^0$	0.387 (8.24)•	0.072 (3.51)•	-0.453 (-5.98)•	0.381 (5.29)•	0.127 (4.37)•	-0.048 (-4.46)•	-0.049 (-3.53)•	-0.011 (-3.92)•
FR	Unrestricted	0.496 (9.60)•	0.072 (0.85)	0.301 (3.39)•	-0.508 (-4.92)•	-0.101 (-2.83)•	0.031 (2.38)°	0.071 (5.02)•	0.009 (2.14)•
	Restricted (H+S)	0.522 (11.6)•	-0.0004 (-0.02)	0.381 (5.19)•	-0.380 (-5.22)•	-0.117 (-4.27)•	0.043 (4.24)•	0.067 (5.17)•	0.009 (3.59)•
	Restricted $(H+S)^0$	0.522 (11.8)•	0 (none)	0.381 (5.29)•	-0.380 (-5.29)•	-0.117 (-4.41)•	0.043 (4.24)•	0.067 (5.17)•	0.009 (3.84)•

Table 5.3 Diagnostic tests for the equations of the unrestricted AIDS models

Equation	DIAGNOSTIC TESTS							
	Serial Correlation		Functional Form		Normality		Heteroscedasticity	
	LM version	F version	LM version	F version	LM version	F version	LM version	F version
PT	0.32 (0.57)	0.23 (0.64)	2.40 (0.12)	1.80 (0.20)	1.333 (0.51)	na	0.45 (0.50)	0.42 (0.52)
SP	0.06 (0.81)	0.04 (0.85)	4.95 (0.03)	4.12 (0.06)	0.63 (0.73)	na	0.38 (0.54)	0.36 (0.55)
FR	0.00 (0.98)	0.00 (0.99)	3.47 (0.06)	2.72 (0.12)	0.78 (0.68)	na	0.47 (0.49)	0.44 (0.51)

The additional restriction of null-cross price effects between the equations for France and Portugal is suggested by the estimation results of model (H+S) and, accordingly, tested. The correspondent χ^2 statistic value is 4.08 indicating that the hypothesis cannot be reject even at the 1% significance level. When this restriction is included in the model, even though the individual significance of some of the coefficients increases, the changes in the magnitudes of the estimates are minimal. This indicates that the estimates of model (H+S) and those of this same model under the additional restriction of cross-price effects null, denote by $(H+S)^0$ in Table 5.2, do not differ significantly. Therefore, the interpretation of the results focuses on the estimates obtained from the model under homogeneity and symmetry (H+S) since this specification complies with the pre-requirements of consumer demand theory and its estimates do not differ from those obtained with the model $(H+S)^0$.

5.4.1.1. Interpretation of the coefficients' estimates of system

The explanatory variables are all significant at the 5% significance level or less, except for the price of Portugal in the equation for France and the price of France, trend and real expenditure variables in the equation for Portugal.[7] In general, the diagnostic tests and the goodness of fit statistics obtained on an equation-by-equation basis with the unrestricted model, show the AIDS system as a statistically robust specification. All the regressions are significant overall, indicating a relatively high explanatory power for the independent variables, and there is no evidence of serial correlation, heteroscedasticity or non-normal distributed errors. For the whole system, the null hypothesis of all coefficients being zero (equivalent to the overall significance F test on an equation-by-equation basis) was tested. The Wald statistic value for this hypothesis is 245.6, indicating a strong rejection of the null.

The regressions' coefficients can be interpreted as follows: *ceteris paribus*, γ_{ij} measures the absolute change in the i^{th} expenditure share following a 1% change in P_j and the coefficients β_i and $(\beta_i + \beta'_i)$ measure the absolute change in country i's budget share per 1% change in UK real per capita expenditure in the period 1969-1979 and in the period 1980-1997, respectively. For Portugal, both these values are insignificant, so that changes in the UK real expenditure per capita do not significantly affect the Portuguese share. However, these coefficients' estimates are significant in the equations for France and Spain.

[7] The equation for Portugal shows evidence of correlation between the trend variable and the real expenditure variable (correlation coefficient is 0.97). When this equation is estimated without the trend, the coefficient of the expenditure variable is positive and significant, but when the trend variable is included the estimated coefficient of the expenditure variable changes sign and becomes insignificant.

If $\beta_i > 0$, the share Wi increases with E and if $\beta_i < 0$, the share Wi decreases when E increases. In demand theory terms, this would indicate tourism in Spain to be a 'luxury' and tourism in France to be a 'necessity'. An alternative terminology might be to categorise Spain as a 'primary' (first choice or preference) destination and France as a 'secondary' (second choice or preference), meaning that consumers would prefer to direct additional expenditure towards Spain rather than France. Generally, a coefficient $\beta_i > 0$ gives rise to expenditure elasticities above unity, and a coefficient $\beta_i < 0$ gives rise to expenditure elasticities in the interval (0, 1). This means that for each 1% increase in the UK expenditure allocated to the region, the share of Spain (France) would respond with a more (less) than 1% increase, which confirms UK tourists' preference for Spain when their budget increases.

The values of the coefficient for the dummy variable D demonstrate that the Spanish and Portuguese political changes and the oil crises that took place in the period 1974-1981 had a negative effect on Spain and Portugal and a net positive effect on France. Hence, in this period, UK tourists' preferences moved in favour of France, relative to Portugal and Spain. The coefficients of the trend variable can be interpreted as the annual average change in the expenditure shares which would take place in the absence of any change in the other explanatory variables. France's share increases while Spain's share decreases by approximately the same amount. The increase in the Portuguese share is insignificant.

A more detailed analysis of the results requires the relevant elasticities values. These were calculated using the following formulae:

Expenditure Elasticities:

$$\varepsilon_i = \frac{d \ln q_i}{d \ln x} = \frac{1}{\overline{w}_i} \frac{d w_i}{d \ln x} + 1 = \frac{\beta_i}{\overline{w}_i} + 1$$

Uncompensated Own-Price Elasticities:

$$\varepsilon_{ii} = \frac{d \ln q_i}{d \ln p_i} = \frac{1}{w_i} \frac{d w_i}{d \ln p_i} - 1 = \frac{\gamma_{ii}}{\overline{w}_i} - \beta_i \frac{w_i^{90}}{\overline{w}_i} - 1$$

Uncompensated Cross-Price Elasticities:

$$\varepsilon_{ij} = \frac{d \ln q_i}{d \ln p_j} = \frac{1}{w_i} \frac{d w_i}{d \ln p_j} = \frac{\gamma_{ij}}{\overline{w}_i} - \beta_i \frac{w_j^{90}}{\overline{w}_i}$$

Compensated Own-Price Elasticities:

$$\varepsilon^{*}_{ii} = \varepsilon_{ii} + w_i^{90}\varepsilon_i = \frac{\gamma_{ii}}{\overline{w}_i} + w_i^{90} - 1$$

Compensated Cross-Price Elasticities:

$$\varepsilon^{*}_{ij} = \varepsilon_{ij} + w_i^{90}\varepsilon_i = \frac{\gamma_{ii}}{\overline{w}_i} + w_j^{90}$$

where $\overline{w}_i$ represents the sample's average share of destination i (i=1,...,n) and w_j^{90} represents the share of destination j (j=1,...,n) in the year base. For France, Spain and Portugal the values of these average and year base shares are the following:

	Year-base share (1990) w^{90}	Overall period average share (1969-1997) $\overline{w}$	First period average share (1969-1979) $\overline{w}_{69-79}$	Second period average share (1980-1997) $\overline{w}_{80-97}$
Portugal	0.0923	0.0775	0.0574	0.0898
Spain	0.4607	0.5635	0.6267	0.5249
France	0.4470	0.3590	0.3159	0.3853

The ensuing discussion focuses on the uncompensated elasticities as they tend to be more important for policy purposes.[8] The estimates of the expenditure and uncompensated own- and cross-price elasticities within the restricted model (H+S) and respective t-values are presented in table 5.4. [9]

[8] The values of the elasticities quantify the sensitivity of tourism demand to changes in the expenditure budget (expenditure elasticities), prices in the destination under consideration (own-price elasticities) and prices in alternative destinations (cross-price elasticities). Compensated price elasticities allow for the effects of changes in real income which accompany price changes while uncompensated price elasticities only consider price changes.

[9] Given the elasticities' definition , their variances (var), based on which the t-values are calculated, are:
$var(\varepsilon_i)=(1/w_i)^2 var(\beta_i)$;
$var(\varepsilon_{ii})=(1/w_i)^2 var(\gamma_{ii})+(w_i^B/w_i)^2 var(\beta_i) - 2[w_i^B/(w_i)^2]covar(\gamma_{ii},\beta_i)$;
$var((\varepsilon_{ij})=(1/w_i)^2 var(\gamma_{ij})+(w_j^B/w_i)^2 var(\beta_i) - 2[w_j^B/(w_i)^2]covar(\gamma_{ij},\beta_i)$.

Table 5.4: Expenditure and uncompensated price elasticities (restricted model)

	Expenditure elasticities		Own-price elasticities		Cross-price elasticities					
					PP		PS		PF	
	First Period	Second Period	First Period	Second Period	First Period	Second Period	First Period	Second Period	First Period	Second Period
WP	0.820 (4.51)	0.947 (10.10)	-2.237 (-5.46)	-1.797 (-6.87)	X	X	1.344 (2.81)	0.830 (2.77)	0.073 (0.16)	0.019 (0.06)
WS	1.203 (25.23)	1.150 (26.97)	-1.817 (-12.94)	-1.933 (-11.74)	0.097 (2.56)	0.124 (2.74)	X	X	0.517 (4.69)	0.658 (4.96)
WF	0.630 (7.25)	0.808 (24.52)	-2.039 (-9.25)	-1.901 (-10.44)	0.033 (0.41)	0.017 (0.25)	1.376 (5.46)	1.077 (5.30)	X	X

5.4.1.2. Expenditure elasticities

The expenditure elasticities of the UK demand for tourism differ considerably between the neighbouring countries; those for Portugal and France are below unity while that for Spain is above unity. The expenditure share for France is less responsive to variations in UK tourism expenditure than those of Spain or Portugal. The expenditure elasticity of the demand for France is of relatively small magnitude, so that France benefits (loses) less from increases (decreases) in UK total tourism expenditure than Spain or Portugal. The expenditure elasticity for Portugal, although inelastic, is close to unity. Spain is, however, an expenditure elastic destination for UK tourists. The values of the expenditure elasticities indicate France as a 'secondary' destination ($0<\varepsilon_F<1$) while Spain can be viewed as a 'primary' destination ($\varepsilon_S>1$).

An interesting aspect is the difference in the magnitudes of the expenditure elasticities for Spain and France between the first and the second periods. The responsiveness of the UK demand for Spain to changes in real expenditure seems to decrease while that of France seems to increase from the first to the second period. This implies that, for the most recent two decades (the second period), France and Spain have been moving in opposite directions in relation to their roles as 'primary' and 'secondary' destinations. This view can be further supported by the differences in the uncompensated own- and cross-price elasticities observed in the first and second periods for these two countries.

5.4.1.3. Own-price elasticities

The uncompensated own-price elasticities are all negative, as expected from normal commodities for which demand responds negatively to increases in prices. In the cases of Portugal and France, the value of the own-price elasticity decreases from the first period to the second period while in the case of Spain an increase is observed. The absolute magnitude of the decrease for Portugal is greater (-0.44) than that for France (-0.14) or the increase for Spain (0.12). Hence, the impact of the changes in prices on the UK tourism demand for these three destinations varies not only across countries but also between the two periods considered. In the period 1969-1979, the UK demand for tourism is more responsive to changes in Portuguese and French prices (a 1% change in prices of France or Portugal induces a demand decrease of more than 2%, *ceteris paribus*) than to price changes in Spain. However, in the period 1980-1997, the UK demand for tourism is more responsive to price variations in Spain than in France or Portugal. The second period values of the own-price elasticities also indicate that changes in effective prices in France and Spain have a greater impact on UK demand for tourism than would result from equivalent changes in Portugal. For all three countries in both periods, the positive returns that could be gained from increases in the UK tourism budget would not compensate the adverse effects of increased prices.

Comparison of the magnitudes of the expenditure elasticities in the first and second periods shows increasing values for France and Portugal and a decreasing value for Spain. On the other hand, the own-price elasticities indicate a decreasing responsiveness of UK demand towards changes in prices of France and Portugal and an increasing sensitivity of UK demand to price variations in Spain. Hence, the estimates of expenditure and own-price elasticities suggest a tendency for Spain to lose ground to France and Portugal.

5. 4.1.4. Cross-price elasticities

Substitutability and complementary among destinations are indicated by positive and negative cross-price elasticities, respectively. Clear conclusions about the complementarily or substitutability among destinations are not usually obtained in studies using the AIDS model which have produced few well-defined cross-price effects. However, the results in this study seem consistent and also coincide with *a priori* expectations. Hence, they are taken as an indication of the relative magnitude and directions of the changes in demand.

In both the first and second periods, all cross-price elasticities are positive, indicating substitutability among destinations. However, since the price of France (Portugal) in the equation for Portugal (France) and the expenditure coefficients in the equation for Portugal are not significantly different from zero, the cross-price elasticities between Portugal and France are expected to be null. Indeed, the insignificance of these cross-price elasticities is statistically confirmed by the respective t-values in Table 5.4. Hence, Portugal and Spain and France and Spain are substitute destinations, while the UK demand for Portugal (France) does not react to changes in the effective price of France (Portugal).

The cross- price elasticities for the equations for France and Spain show that the share of France is more sensitive to price changes in Spain than that of Spain is to price changes in France. However, this sensitivity alters from the first to the second period, showing a decrease in the responsiveness of UK demand for France to price variations in Spain and an increase in the responsiveness of the UK demand for Spain to price variations in France. If Spanish prices increase by 1%, the share of France increases by 1.38% in the first period and by only 1.08% in the second, while if prices increase by 1% in France, the demand for tourism in Spain will increase by 0.52% in the first period and by 0.66% in the second. This is consistent with Spain's loss of ground relative to France, mentioned above.

The results also indicate Spain and Portugal as substitutes, although price changes in Portugal have a minor effect in the UK demand for Spain compared with the effect on the UK demand for Portugal caused by price changes in Spain. This is not surprising given the difference in the sizes of the tourism markets of both countries. Yet, the situation again alters from the first to the second period. The results posit that if prices in Portugal increase by 1%, the demand for tourism in Spain will increase by 0.10% in the first period and by 0.12% in the second. These very small effects show the low sensitivity of UK demand for Spain to price changes in Portugal. However, if prices in Spain increase by 1%, the UK demand for Portugal increases by 1.35% in the first period and 0.83% in the second. Hence, although the sensitivity of the UK demand for Portugal to price changes in Spain remains greater than that for Spain to price changes in Portugal, the former diminishes in the two last decades. In contrast, the sensitivity of the UK demand for Spain to price changes in Portugal increases. This provides further evidence of the increased sensitivity of the UK demand for Spain to price variations in its neighbours.

5.4.2. COMPARABILITY OF THE RESULTS WITH THOSE FROM OTHER STUDIES

Studies that have applied the AIDS approach to the analysis of tourism demand are few in number. Among them, White (1982) and O'Hagan and Harrison (1984) modelled the US demand for tourism in Europe between 1964 and 1981, Syriopoulos and Sinclair (1993) investigated US and Western European tourism expenditure allocation to several Mediterranean countries in the period 1960-1987, and Papatheodorou (1999) studied the demand for international tourism in the Mediterranean region during the period 1957-1990.

The elasticities estimates provided in this study are only comparable with those obtained by Syriopoulos and Sinclair (S&S) and Papatheodorou (PTH) in the cases of Spain and Portugal. However, this comparison must allow for the fact that these studies examine different sample periods, different data sources and different sets of independent variables. table 5.4 shows the expenditure elasticities and the uncompensated own- and cross-price elasticities obtained in the studies.

Table 5.4. Comparison of elasticities estimates of UK tourism demand in Portugal and Spain

	Expenditure Elasticities			Uncompensated Own-price Elasticities			Uncompensated Cross-price Elasticities					
	S&S	PTH	Model (5.13)	S&S	PTH	Model (5.13)	S&S		PTH		Model (5.13)	
							PP	PS	PP	PS	PP	PS
WP	1.58	0.04	0.82 ; 0.95	-2.81	-2.85	-2.24 ; -1.80		3.75		0.88		1.34 ; 0.83
WS	0.90	1.15	1.20 ; 1.15	-1.11	-1.30	-1.82 ; -1.93	0.65		0.88		0.10 ; 0.12	

A noticeable difference between the results concerns the expenditure elasticity for Portugal. While our results posit the expenditure elasticity of the UK demand for Portugal to be inelastic but close to unity, Syriopoulos and Sinclair's (1993) results indicate Portugal as an expenditure elastic destination and Papatheodorou (1999) shows it as not being statistically different from zero. The expenditure elasticity for Spain is found to be inelastic in Syriopoulos and Sinclair's study, while the results in Papatheodorou and this study qualify it as elastic with similar values. While the own-price elasticities estimates for Portugal by S&S and PTH are similar, our results show a lower estimate in both the first and second periods. The same comment can be made for the own-price elasticity for Spain. The uncompensated cross-price elasticities' estimates in the three studies are rather different. The response of the UK demand for Portugal to price changes in Spain is highly elastic (3.75) by S&S, while PTH finds that the response is inelastic (0.88). Our results indicate the response of the UK demand for Portugal to price changes in Spain to be elastic in the first period, and inelastic in the second. The response of the UK demand for Spain to price changes in Portugal is inelastic according to S&S (0.65) and PTH (0.88) studies. The results obtained in this study also show an inelastic but much smaller response in both periods (0.10 and 0.12).

The differences between the elasticities estimates obtained by these three studies can be further explained by the use of the unrestricted model for the elasticities derivation in Syriopoulos and Sinclair's study, while Papatheodorou's study and this one use the model under homogeneity and symmetry for the same purpose. The inclusion of a structural break in the real expenditure variable in model (5.13) may also explain some of the differences.

5.5. CONCLUSION

This chapter applies the static AIDS model in two contexts not addressed in other studies that have estimated system of equations models of tourism demand: first, the model was applied to neighbouring destinations, two of which experienced a transition from features characteristic of developing countries at the beginning of the sample period, towards higher levels of income and welfare by the end of the 1990's. Second, the model specification allows for comparison of changes in tourism demand behaviour in each destination over time, in terms of the values of expenditure, own- and cross-price elasticities. The estimation results show the model to be data-consistent and statistically robust, as indicated by both the diagnostic tests and the goodness of fit statistics. Moreover, in contrast to many studies which have estimated system of equation models, the results are consistent with the properties of homogeneity and symmetry. This accords with the microfoundations of the AIDS approach and increases the credibility of the elasticity values.

The elasticities estimates obtained from the model provide interesting insights concerning the behaviour of tourism demand in the two lower income countries, both relative to each other and to their richer neighbour. Although the expenditure elasticity for Spain is marginally lower in the second period than in the first, while that for France is higher, the demand for Spain is expenditure elastic and that for France is expenditure inelastic in both periods. Therefore, a "*highly developed tourism infrastructure*" and the "*offer of a diversified product to satisfy different groups of tourists*" (Papatheodorou, 1999) do not seem to be the main justification for an elastic or inelastic demand response to changes in real expenditure, since France can hardly be seen as lagging behind Spain in these respects. Rather the proximity of France to the UK, inducing tourism products to be preferably consumed in a different way, seems to be a more acceptable reason for the differences in expenditure elasticities observed in these two countries. The proximity of France to the UK (which has been 'increased' with the Channel Tunnel opening in 1994) seems to induce a shorter average length of stay of UK tourists than that assigned to Spain. The tourists who visit Spain appear to be willing to spend more time in this country. A glance in the average length of stay of UK tourists in Spain and France (see chapter 2) confirm France as a shorter-term destination than Spain. These differences, rather than a highly developed tourism infrastructure, may better explain the elastic (more volatile) tourism demand for Spain compared with the inelastic (more stable) tourism demand for France.

In both periods, the expenditure response of the UK demand for Spain is elastic, while that for France is inelastic and that for Portugal is close to unity. These results indicate that Spain has benefited from increases in the expenditure budget relative to France. Hence, increases in the tourism expenditure budget can assist poor countries to 'catch-up' with their richer counterparts. However, Portugal did not benefit as significantly as Spain from increases in the UK tourism budget, so that the two Iberian countries have experienced different patterns of behaviour over time. These findings are clearly relevant to policy-makers who are concerned with the responsiveness of tourism demand to changes in expenditure. For example, Portugal's relatively stable gains from increases in the expenditure budget, along with Spain's apparently declining share, corroborated by the negative sign of the trend coefficient, merit further attention by the relevant tourism authorities.

The estimates of the uncompensated own-price elasticities for the first period show Spain to be a less price elastic destination than France but in the second period they are more similar. This indicates that the price sensitivity of tourism demand can vary over time. Although UK tourists remain more sensitive to price changes in France than in Spain, they are becoming increasingly aware of price changes in Spain, implying that Spain may experience more instability of demand relative to France if this trend continues. The sensitivity of the UK demand for Spain is increasing in response not only to changes in its own-price but also to changes in the prices of its competitors, France and Portugal, as shown by the values of the cross-price elasticities in the first and second periods. In contrast, the sensitivity of the UK demand for France and Portugal to price changes in Spain demonstrates a tendency to diminish between the first and second periods.

Portugal is a small country, in terms of tourism, compared with its neighbours. This fact, combined with the political and historical events that contributed to its lagging behind other European countries in development terms, may explain Portugal's late awareness of its tourism potential. However, calculations based on World Tourism Organisation data (1999) show that between 1985 and 1997, tourism in Portugal has increased faster (except for France in terms of arrivals and for the Netherlands in terms of receipts) than any other western or southern continental European country. In fact, tourists arrivals in Portugal increased by 60% between 1987 and 1996, only surpassed by France with an increase of 69%, and Portugal's tourism receipts increased 275% between 1985 and 1997, only surpassed by the Netherlands with an increase of 288% in the same period. Furthermore, Portugal's share of international tourist arrivals has exceeded 1.5% since 1990, classifying the country as one of the twenty most visited countries in the world (WTO, 1996). Nevertheless, in terms of receipts, the picture is less favourable as Portugal drops to a position below 24 among the most important tourism destinations. In terms of tourism arrivals in Portugal, Spain is the most important origin, followed by the UK, while in terms of receipts, the UK is the most important origin (Instituto Nacional de Estatistica, 1975-1997). Spanish tourists in Portugal are mainly short-term visitors who enter the country in huge numbers but spend very little. UK tourists display the opposite behaviour, entering the country in much smaller numbers but spending more money during their longer-term visits.

Although Portugal undertook a major effort of modernisation in an attempt to adjust to its EC partners' development levels, this effort, in tourism terms, has only been noticeable during the 1990s. For most of the sample period the data reflect an underdeveloped but picturesque country, squeezed between the sea and enormous Spain, with wonderful sites but no proper accommodations, lovely clean beaches but no proper roads, tasty, varied and cheap food but no proper service, with museums that few visit, art that few see, history that few know. There are, of course, places in Portugal with the cosmopolitan environment seen in similar destination countries, but they tend to be too crowded, noisy and chaotic for tourists looking for a minimum of quality to accompany rising service prices. These features are not easy to capture in an econometric model and may be one of the reasons for the statistical performance of the equation for Portugal, rendering the elasticities estimates more difficult to interpret.

However, there is sufficient evidence to indicate the changes that Portugal has been experiencing in tourism terms. For instance, although the expenditure elasticity of the UK demand for Portugal is close to unity in both periods, its estimate increases from 0.82 in the first period to 0.95 in the second. If additional observations could provide significant confirmation of this growing tendency, this would mean that an increase in the UK tourism budget allocated to the region would be accompanied by a larger share for Portugal in the second period *ceteris paribus,* indicating a growing UK preference for Portugal and assisting the country to 'catch-up' with its richer neighbours. The own-price elasticity in both periods clearly portrays a diminishing sensitivity of the UK demand for Portugal with respect to own-price changes. In fact, from the first to the second period, this elasticity estimate drops to a value even below that of France, while Spain shows an increase in its second period own-price elasticity estimate.

The cross-price elasticities provide useful information about the interrelationships between destinations. The cross-price elasticities between Portugal and France indicate the lack of sensitivity of the UK demand for Portugal (France) to price changes in France (Portugal) since for both periods, these elasticities are not significantly different from zero. The case is different for the cross-price elasticities estimates between Spain and Portugal. For both periods and in both countries, the UK demand for tourism in one country responds significantly to price changes in the other. However, as expected, the UK demand for Spain is only marginally affected by price changes in its smaller neighbour, while the UK demand for Portugal is more responsive to price changes in Spain, particularly in the first period. Since the UK demand for Portugal responds less intensively to price changes in Spain in the second period relative to the first, we can infer that the UK demand for Portugal may have achieved more stability along with greater independence from price changes in Spain in the last two decades.

Political changes are likely to have an adverse effect on tourism demand as indicated by the negative coefficient of the dummy variable D for Portugal and Spain. For these two destinations, the dummy coefficients indicate a decrease in their average shares in the period 1974-1981, while the positive coefficient of this variable in the equation for France shows an increase in its average share in the same period. The estimated coefficient for the trend variable also shows an increasing average share of UK tourism expenditure for France, while Spain's share tends to decrease. This further confirms the already stated loss of ground of Spain relative to France in UK tourists' preferences, and indicates that countries cannot rely on continuing gains from tourism based on stability of 'traditional' tastes.

This application of the AIDS model has provided new information about the long-run behaviour of the UK tourism demand for France Spain and Portugal. For example, it has shown that although lower income countries can benefit from increases in the expenditure budget relative to their neighbours, they will not do so automatically. Portugal and Spain, for example, failed to achieve similar gains from increases in the UK tourism budget. Exploration of the reasons why some lower income countries experience disproportionate gains from rises in tourism expenditure is a topic for investigation, as is the reason why countries can fail to maintain their advantageous position. The study has also shown that competition for international tourism demand, rather than complementarity, characterised the neighbouring countries under consideration. Investigation of the degree to which competitiveness rises or falls as other countries attain more similar levels of development is a further topic for research.

Nevertheless, the AIDS modelling approach adopted in this chapter does not take into consideration the dynamic features of tourism demand behaviour and, consequently, cannot account for the short-run correction mechanism which underlies the adjustment process of the variables' levels to their equilibrium path. Although, in many instances, the focus of interest of economic research is to uncover the long-run structural relationships between dependent and independent variables, the short-run dynamics underlying the steady-state equilibrium model are also of importance, particularly in cases where the short-run effects may have relevant magnitudes. Therefore, the incorporation of the short-run dynamics within the AIDS system approach is the next logical step in this comprehensive investigation of tourism demand behaviour. The next chapter concerns the specification and estimation of a dynamic AIDS model for the UK tourism demand in France, Spain and Portugal.

CHAPTER 6

DYNAMIC SYSTEM OF EQUATIONS IN TOURISM DEMAND MODELLING

6.1. INTRODUCTION

One of the interesting features of economic research on consumer demand behaviour is the close relationship between theoretical specifications and appropriate modelling and estimation techniques. Given that some notion of the consumer's optimising behaviour is assumed to underlie the data generating process, effective evaluation of the model's specification requires judgement from both a theoretical and an empirical standpoint. Many of the now standard econometric procedures available were developed in response to practical problems in interpreting demand observations and utility theory has provided a structure and a terminology for model formulation and data analysis. Although the construction of models which are both theoretically and empirically satisfactory is never straightforward, the role played by economic theory as a tool of applied economics is essential. Consumer demand analysis based on the articulation between theory and empirical evidence has been the core of important theoretical advances, introducing new modelling perspectives for the behavioural features of economic agents. This is specially the case of the integration of dynamics in theoretically consistent demand models.

Economic theory of consumer behaviour provides sophisticated mathematical descriptions of equilibrium relationships between demand and its determinants and, typically, is concerned with changes in the steady-state equilibrium caused by changes in the demand determinants. However, most of the theory underlying demand systems does not clearly specify the dynamic process by which long-run equilibrium values are attained. Yet, actual demand behaviour over time must reflect the imperfections of the adjustments that consumers entertain when confronted with changes in their demand determinants.

In chapter 5, a tourism demand steady state structure was estimated using the AIDS approach and the results were satisfactory both from an econometric and an economic point of view. However, it is possible that current changes in the budget shares of the UK demand for tourism depend not only on current changes in prices and expenditure, but also on the extent of the variables' disequilibrium values in previous periods. Consumer preferences may have been unstable over the sample period or the parameters of the utility function may have shifted over time. Evidence that this might be the case was provided by the estimation results of the AIDS model in chapter 5. Therefore, an explicitly dynamic specification within a system of equations should be considered in accounting for the adjustment process towards the long-run equilibrium.

The temporal aspect of consumer behaviour has been largely ignored in studies concerning system of equations in a tourism demand context. Early research efforts concentrate on static systems in order to test the hypothesis of utility theory. However, these specifications of consumer optimisation may be legitimate approaches under special restrictive circumstances only. Indeed, aggregate models explaining demand solely in terms of current prices and expenditure are likely to omit relevant time-changing factors which may be correlated with the variables included, making it difficult to identify the separate effects[1]. Yet, the inclusion of a trend variable may help to correct for some of the omitted factors and thereby make results more theoretically consistent. In these special circumstances, the specification of dynamics in demand models can be simple to formulate. However, this simplicity is often the result of particularly strong assumptions, for example, intertemporal separability, unchanged preferences and invariant budget constraints over time. As stressed by Blundell (1988 p. 39), "*the intertemporal separability assumption on which these models rest is precisely that which rules out explicit dynamics*".

The AIDS model discussed in chapter 5, gives a consistent temporal interpretation of the long-run relationships present, despite the underlying assumption of intertemporal separability. Within this static specification, we identified a trend-like behaviour in the UK demand for tourism, not associated with changes in prices or expenditure, which suggested that changes in consumer preferences have occurred over the sample period. To allow for this possibility we included a trend variable which supposedly 'absorbs' the effects of time-dependent shifting factors and may be interpreted as a trend change in preferences. We also

[1] In a tourism demand context these factors may be identified with, for example, reductions in the working week, increased entitlement to paid leave and increased numbers of retired or early-retired people

allowed, by means of a structural break imposed on the expenditure variable, for some variability in the budget constraint. Hence, these models can be qualified as 'quasi-' or 'seemingly-dynamic' models. Still, under more general dynamic behavioural features, the AIDS model needs to include (explicit) dynamic elements in its demand equations.

If the intertemporal separability assumption is put aside, consumers adjust their demand in response to intertemporal changes in its determinants, and an explicitly dynamic adjustment process is allowed for. In fact, it is realistic to consider that past behaviour changes preferences and, consequently, affect current behaviour. Habit implies that consumer utility functions are influenced by previous purchases which, in turn, influence present purchases. Since habits are usually unobservable, the associated changes in the demand functions are usually represented by lagged demand, prices and expenditure variables. However, as Blundell (op. cit.) points out, there may be little to be learnt from simply adding lagged variables to an otherwise static model. The resulting specification in this case, may only be intelligible under implausible behavioural hypotheses. Unless empirical models are appropriately specified and the implications of general theoretical principles are fully integrated, invalid statistical inference may rule out plausible types of behaviour, inducing research to proceed in less useful directions.

Many researchers in demand behaviour have recognised the importance of including explicit dynamic adjustments in demand systems and a number of approaches have been adopted. For example, Anderson and Blundell (1983, 1984) implement a flexible general dynamic approach which incorporates a long-run solution within a system of error-correction equations, to analyse UK consumer demand for several groups of commodities, and Burton and Young (1996) use a dynamic structure to model a system of equations which describes the adjustment process of meat consumers faced with the BSE problem. Unfortunately, specific research on tourism demand dynamics using systems of equations, is not abundant and, to the best of our knowledge, only Lyssiotou (1999) addresses these issues in an empirical study concerning the dynamics of tourism demand behaviour, using a system of equations similar to the AIDS model.

The objective in this chapter is to contribute an empirical study of tourism demand dynamics and point out areas where the scrutiny of relationships between theoretical and empirical considerations are likely to produce new insights in this area of research. The structure of the chapter is as follows. Section 6.2 provides several alternative empirical specifications derived from a general dynamic AIDS model. Section 6.3 implements statistical tests of theoretical hypotheses seeking for consistency between the models and the principles of consumer demand theory. Section 6.4, presents the estimation results of theoretically consistent models, which are analysed and compared with the results of the AIDS model estimated in chapter 5. Section 6.5 concludes.

6.2. DYNAMIC AIDS MODELLING OF THE UK DEMAND FOR TOURISM

In a purely static demand system, consumers are assumed to adjust perfectly and instantaneously to prices and income changes. However, habit persistence, unstable preferences, adjustment costs, or imperfect information leading to incorrect expectations, may prevent consumers from adjusting fully to equilibrium every period. If this is the case, a more general specification of a dynamic structure of consumer behaviour is required. This general specification can be modelled in a way that allows for testing more restrictive models such as the static hypothesis itself, and alternative dynamic forms nested within the basic general model. Once a data-coherent structure is defined, the restrictions implied by utility maximisation hypotheses can be tested.

In what follows, a flexible dynamic structure for the AIDS model is derived, based on the work of Anderson and Blundell (1983, 1984). As pointed out by these authors, nested within the flexible general structure are the static model itself and dynamic specifications, such as the partial adjustment and the auto-regressive distributed lag (ARDL) models, which can be tested against the more general alternative.

Consider the AIDS model described in equation (5.12). For simplicity and notation compatibility, we rename the variables as w_i = Wi, ln p_j = Pj where i, j = 1 ,..., n and ln(x/P*) = E. Hence, equation (5.12) is written:

$$Wi_t = \alpha_i^* + \sum_j \gamma_{ij} Pj_t + \beta_i E_t, \tag{6.1}$$

Consider equation (6.1) as the appropriate choice for the steady state structure of the following general dynamic stochastic specification:

$$\Delta Wi_t = \sum_{j=1}^{n} \gamma_{ij}^S \Delta Pj_t + \beta_i^S \Delta E_t + \lambda_i \left(\alpha_i + \sum_{j=1}^{n} \gamma_{ij}^L Pj_{t-1} + \beta_i^L E_{t-1} - Wi_{t-1} \right) + ui_t \tag{6.2}$$

where λ_i is the adjustment coefficient of the i^{th} equation, Δ is the first difference operator, subscript t-1 indicates the lagged values of the variables, and ui is the i^{th} disturbance term assumed to be characterised by a singular, independent and identical distribution over time.

The parameters with superscript S and L can be interpreted, respectively, as the short-run and long-run responses of the dependent variable to changes in its determinants.

The dynamic specification in (6.2) assumes that, in order to maintain the steady state relationship given in (6.1), consumers adjust the current values of their expenditure shares partly in response to current changes in the explanatory variables and partly in response to the desequilibrium observed in the previous period. As in Burton and Young (1996), we constrain specification (6.2) to the use of an identical adjustment coefficient across equations such that, $\lambda_i=\lambda$ for all i. Given the system's singularity, estimation may be carried out by the arbitrary delectation of the n^{th} equation.

Although the dynamic specification (6.2) is a simple first order lagged structure, the model may still be too general for any particular data generating process, resulting in a loss of estimation precision or in what Gujarati (1997) classifies as a 'shaky' statistical inference procedure. Hence, a sequence of tests may be performed to find the most restrictive dynamic specification consistent with a particular set of T observations on budget shares, prices and per capita real expenditure. Examples of such specifications are provided below.

Consider the following equivalent form of the general equation (6.2):

$$\begin{aligned}\Delta Wi_t &= \gamma_{i1}^S P1_t - \gamma_{i1}^S P1_{t-1} + \ldots + \gamma_{in}^S Pn_t - \gamma_{in}^S Pn_{t-1} + \beta_i^S E_t - \beta_i^S E_{t-1} + \\ &+ \lambda\alpha_i + \lambda\gamma_{i1}^L P1_{t-1} + \ldots + \lambda\gamma_{in}^L Pn_{t-1} + \lambda\beta_i^L E_{t-1} - \lambda Wi_{t-1} + ui_t\end{aligned} \quad (6.3)$$

Auto-regressive distributed lag (ARDL) model

In the spirit of a "*general to specific*" approach we postulate the long-run equilibrium relationship between two economic variables, say Y and X, such that:

$$Y_t = \beta_0 X_t + \beta_1 X_{t-1} + \ldots + \beta_m X_{t-m} + \delta_1 Y_{t-1} + \delta_2 Y_{t-2} + \ldots + \delta_m Y_{t-m} + u_t,$$

which is an ARDL model of order m. These models explicitly consider the behaviour of a variable over time and so are dynamic in nature. However, such models are too general since the value of m remains unspecified. This general form may be reduced to a parsimonious one by applying several criteria (see, for example, Hendry and Richard, 1983) which include the definition of m as a small number. The model's general form (6.3) can be reduced to an ARDL form as described, under the null hypothesis:

$H_0: \lambda\alpha_i = 0 \cap \gamma_{ij}^S = \gamma_{ij}^L \cap \beta_i^S = \beta_i^L$ for all i, j

If H_0 is not rejected the model reduces to:[2]

$$Wi_t = \gamma_{i1}P1_t + (\lambda-1)\gamma_{i1}P1_t + ... + \gamma_{in}Pn_t + (\lambda-1)\gamma_{in}Pn_t + \beta_i E_t + (\lambda-1)\beta_i E_{t-1} +$$
$$+ (1-\lambda)Wi_t + ui_t$$

which is a first order ARDL model.

Partial adjustment model

Consider the flexible accelerator model of economic theory which assumes that the equilibrium, or long run level of a dependent variable, say Y^*_t, is a linear function of an explanatory variable, say X_t, such that

$$Y^*_t = \beta_0 + \beta_1 X_t + u_t \qquad \text{(i)}$$

The partial adjustment hypothesis postulates

$$Y_t = \delta Y^*_t + (1-\delta)Y_{t-1} \qquad \text{(ii)}$$

Substituting (i) in (ii) we have

$$Y_t = \delta\beta_0 + \delta\beta_1 X_t + (1-\delta)Y_{t-1} + v_t \qquad \text{(iii)}$$

Considering model (6.3), the null hypothesis to be tested is

H_0: $\gamma^S_{ij} = \lambda\gamma^L_{ij} \cap \beta^S_i = \lambda\beta^L_i$ for all i, j.

If H_0 is not rejected the model reduces to:

$$\Delta Wi_t = \lambda\gamma^L_{i1}P1_t - \lambda\gamma^L_{i1}P1_{t-1} + ... + \lambda\gamma^L_{in}Pn_t - \lambda\gamma^L_{in}Pn_{t-1} + \lambda\beta^L_i E_t - \lambda\beta^L_i E_{t-1} +$$
$$+ \lambda\alpha_i + \lambda\gamma^L_{i1}P1_{t-1} + ... + \lambda\gamma^L_{in}Pn_{t-1} + \lambda\beta^L_i E_{t-1} - \lambda Wi_{t-1} + ui_t$$

or $\quad Wi_t = \lambda\alpha_i + \lambda\gamma^L_{i1}P1_t + ... + \lambda\gamma^L_{in}Pn_t + \lambda\beta^L_i E_t + (1-\lambda)Wi_{t-1} + ui_t$

which is a partial adjustment model similar to the one described in equation (iii).

Static AIDS model

To test for the static model nested within (6.3) the null hypothesis is

H_0: $\gamma^S_{ij} = \lambda\gamma^L_{ij} \cap \lambda = 1$

If H_0 is not rejected, model (6.3) reduces to

[2] The hypothesis tested in this case, includes the restriction of null intercepts ($\lambda\alpha_i = 0$). However, in a tourism demand context, the intercept has an important economic meaning. Therefore, another hypothesis to be tested for the ARDL model, would be not to include the restriction of null intercepts in H_0. In this case, the ARDL equations resulting from the non-rejection of the null would be similar to those analysed in chapter 4.

$$Wi_t = \lambda\alpha_i + \gamma_{i1}P1_t + ... + \gamma_{in}Pn_t + \beta_i E_t + u_t$$

which is the steady state orthodox AIDS model.

6.3. TESTING THE THEORETICAL CONSISTENCY OF ALTERNATIVE DYNAMIC MODELS

Dynamic generalisations of traditional static systems are an important feature of recent empirical work in several research areas. The motivation for such generalisations largely derives from the continuing lack of accord between the postulates of demand theory and empirical static demand functions using aggregate time series. Generally, consumers do not adjust perfectly and immediately to changes in their demand determinants. Hence, appropriate modelling of the short-run dynamics of consumers' expenditure shares is essential before testing utility maximisation hypotheses. This is so because the results of such tests may depend on whether a dynamic specification is considered and, if so, whether it is the appropriate one. In fact, many studies using the orthodox AIDS approach, for example, Deaton and Muellbauer (1980a), Anderson and Blundell (1983) and Syriopoulos and Sinclair (1993), suggested that a dynamic misspecification is the probable cause of rejection of the utility theory postulates. With most economic models, and particularly with demand systems, the analysis is centred on the set of hypotheses relating to the long-run structural coefficients. This analysis is independent from any short-run dynamics fitted to the data. However, as pointed out by Chambers (1993, p.728), "*the dynamic structure may itself be affected by the restrictions imposed to the long-run solution, and the incorporation of inappropriate dynamics may, in turn, affect the outcome of hypothesis tests conducted on the long-run parameters*". Therefore, adequate dynamic specifications are essential in aggregate demand systems, before plausible behavioural hypothesis can be appropriately tested against observed expenditure patterns.

Following this line of reasoning the methodological strategy in this section consists of: first, to define, in the most generic terms possible, a dynamic flexible structure for the expenditure adjustment process of UK tourism consumers; then, to test for more restrictive specifications believed to be consistent with the sample observations: finally, to test the utility maximisation restrictions on specifications not rejected by the data.

Equation (6.3) in section 6.2. represents a basic general dynamic structure which is used to derive the more restrictive specifications of the ARDL, partial adjustment and static AIDS models. In this section, the compatibility of these models with the UK tourism demand data is tested. If the models are compatible with the data, they are further subjected to utility theory constraints and tested under these hypotheses. The non-rejection of these hypotheses indicates the models as theoretically consistent. The subsequent empirical analysis focuses on dynamic specifications which are found to be both data-compatible and theoretically-consistent.

Hypothesis testing performed on alternative dynamic models provided the following results. The Wald test for the partial adjustment hypothesis against the more general dynamic specification presents a statistic value of $\chi^2(9) = 14.51$, which lies below the corresponding critical value of 16.92, at the 5% significance level. Therefore, the partial adjustment model was not rejected by the data.. However, when further constrained with the restrictions of homogeneity and symmetry, the statistical performance of this model was not satisfactory. The orthodox static AIDS hypothesis was rejected against the general dynamic model, with the Wald statistic value of 19.03. The ARDL hypothesis with null intercepts restriction, was rejected against the general dynamic model with the Wald statistic value of 87.71.[3] Only the general dynamic structure reveals itself compatible with the data and with the assumptions of consumer demand theory. Therefore, the remainder of this section focuses on this model.

The dynamic AIDS model presented in equation (6.3) can be rewritten as follows:

$$\begin{aligned}\Delta Wi_t = \lambda\alpha_i + \lambda\gamma_{iP}^{L}PP_{t-1} + \lambda\gamma_{iS}^{L}PS_{t-1} + \lambda\gamma_{iF}^{L}PF_{t-1} + \lambda\beta_i^{L}E_{t-1} - \lambda Wi_{t-1} + \\ + \gamma_{iP}^{S}\Delta PP_t + \gamma_{iS}^{S}\Delta PS_t + \gamma_{iF}^{S}\Delta PF_t + \beta_i^{S}\Delta E_t + u_{i_t}\end{aligned} \qquad (6.4)$$

Empirical evidence obtained in chapter 5 indicates the presence of a structural break in the coefficient of the real expenditure variable and the significance of a dummy variable (D) in the share equations for France, Spain and Portugal, over the period 1974-1981. this information was integrated in the general dynamic specification (6.4) by considering the presence of a structural break in the expenditure variable and the potential relevance of the dummy variable D. However, for the dynamic model, we assume that the structural break is only relevant in the long-run, whereas the dummy variable D may have a significant effect

[3] When the ARDL specification is tested without the restriction of null intercepts, the model is (marginally) not rejected with a $\chi^2(9)$ statistic value of 15.60. However, this specification is similar to the one used in chapter 4, which removes most of the interest in the empirical analysis of this model.

both in the long- and in the short-run. Investigation of whether this is the case is undertaken by defining an unrestricted general dynamic model including the variable under test in equation (6.4). Then, we assess the significance of this variable by testing the unrestricted (U) model against the restricted (R) model, which excludes the variable.

The statistical significance of the variables SE_{t-1}, D_{t-1}, and ΔD_t is tested for the system of equations, using the Wald statistic. The tests' results are presented in table 6.1 (p values in brackets). In all tests, the explanatory variables in the unrestricted (U) and restricted (R) models are displayed.

Table 6.1: Tests for the individual and joint significance of variables SE_{t-1}, D_{t-1} and ΔD_t

HYPOTHESIS UNDER TEST	WALD STATISTIC
H_0: Non-significance of SE_{t-1} (U): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} SE_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t (R): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t	$\chi^2(2)=4.24$ (0.12) Not rejected
H_0: Non-significance of ΔD_t (U): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t ΔD_t D_{t-1} (R): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t	$\chi^2(2)=2.84$ (0.24) Not rejected
H_0: Non-significance of D_{t-1} (U): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} SE_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t ΔD_t D_{t-1} (R): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t	$\chi^2(2)=6.94$ (0.03) Rejected
H_0: Joint Non-significance of D_{t-1} and SE_{t-1} (U): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} SE_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t D_{t-1} (R): PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} Wi_{t-1} ΔPP_t ΔPS_t ΔPF_t ΔE_t	$\chi^2(4)=10.90$ (0.03) Rejected

The tests indicate that the dummy variable D is only relevant in the long-run, since the non-significance of ΔD_t cannot be rejected and the non-significance of D_{t-1} is rejected. The non-significance of variable SE_{t-1}, accounting for the structural break in the expenditure variable, is not rejected. However, the rejection of this variable's individual significance is not very strong. Moreover, the joint significance of D_{t-1} and SE_{t-1} is convincingly not rejected by the data.[4] For these reasons and for comparison purposes with the static AIDS model of chapter 5, we believe this variable should be included.

The assumption of equal velocities of adjustment for all share equations in the dynamic system was also tested. The null hypothesis for this test restricts the adjustment coefficients λ_i to be equal across equations, so that $\lambda_i = \lambda$ for all i. The Wald statistic value $\chi^2(1) = 0.267$ (p value=0.605), does not rejected this hypothesis even at the 1% significance level.[5]

Given the considerations above, the i[th] equation of the general dynamic AIDS model for the UK tourism demand in France Spain and Portugal is:

$$\Delta Wi_t = a_{i1} + a_{i2}PP_{t-1} + a_{i3}PS_{t-1} + a_{i4}PF_{t-1} + a_{i5}FE_{t-1} + a_{i6}SE_{t-1} + a_{i7}D_{t-1} - \lambda Wi_{t-1} + a_{i8}\Delta PP_t + a_{i9}\Delta PS_t + a_{i10}\Delta PF_t + a_{i11}\Delta E_t + ui_t \quad (6.5)$$

where a_{ik} (i=F, S, P and k=1,...11) are parameters, FE_{t-1} represents the UK real expenditure variable in the first period (1969-1979) and SE_{t-1} represents this variable in the second period (1980-1997).[6]

The general dynamic system (6.5) is now tested against further constrained models which include the restrictions of homogeneity, symmetry and null cross-price effects between the equations for France and Portugal, both on the long- and short-run coefficients of the appropriate variables. Table 6.2. presents these tests results showing in its first column, the null hypothesis under testing; in its second column, the corresponding models' notations which are used hereafter, and in its third column, the Wald test statistic values for the hypotheses under consideration.

[4] It is possible that the demand instability, captured by the structural break in the static model, reveals itself in different ways now, given the dynamic structure adopted. For instance, it is possible that, within the dynamic specification, the instability observed is less relevant when expressed by a structural break in the expenditure variable than it would be if expressed in a different way. Indeed, it is possible that the observed demand instability could be better described by the inclusion of additional intercept dummies. This possibility is considered in chapter 7.

[5] This is not surprising since there is no evident reason to believe that different coefficients of adjustment should exist for the share equations of the destinations considered. Indeed, UK tourists have fairly similar information about these destinations, implying that they adjust to changes in their demand determinants with similar speed. The constraint of equal adjustment coefficients across equations is integrated in all subsequent restricted models derived from the general dynamic structure.

[6] The structural break in the model is included by dividing the sample period into the same two sub-periods defined in chapter 5 (1969-1979; 1980-1997) and using two dummy variables, F and S, which assume a value of unit for observations in the first and second periods respectively, and zero otherwise. These two dummies are then multiplied by E_{t-1} giving rise to the new variables FE_{t-1} and SE_{t-1} integrated in (6.5). Model (6.5) including FE_{t-1} and SE_{t-1} is an equivalent form of a model including E_{t-1} and SE_{t-1} variables. The former has the advantage of giving straightforward information on the coefficients of variable E in the first and second periods (respectively, a_5 and a_6), whereas in the latter, the information for the second period has to be obtained by summing the coefficients of E_{t-1} and SE_{t-1}.

Table 6.2: Tests for utility theory restrictions on the long- and short-run coefficients

HYPOTHESIS UNDER TEST[7]	MODEL NOTATION	WALD STATISTIC
H_0: Long-run homogeneity	(H)L	$\chi^2(3)=0.541$ (0.91) Not rejected
H_0: Long-run homogeneity and symmetry	(H+S)L	$\chi^2(4)=2.16$ (0.71) Not rejected
H_0: Long-run homogeneity and symmetry and long-run null cross-price effects between the share equations of France and Portugal	$(H+S)^0L$	$\chi^2(5)=2.56$ (0.77) Not Rejected
H_0: Long-run homogeneity, symmetry and null cross-price effects and short-run homogeneity	$(H+S)^0L$ & (H)S	$\chi^2(7)=8.72$ (0.27) Not Rejected
H_0: Long-run homogeneity, symmetry and null cross-price effects and short-run homogeneity and symmetry	$(H+S)^0L$ & (H+S)S	$\chi^2(8)=10.35$ (0.24) Not Rejected
H_0: Long-run homogeneity, symmetry and null cross-price effects and short-run homogeneity, symmetry and null cross-price effects	$(H+S)^0LS$	$\chi^2(9)=10.59$ (0.31) Not Rejected

The results in table 6.2 show that none of the hypotheses is rejected by the data. Therefore, once the dynamics of adjustment to equilibrium are fully acknowledged by an econometric structure, long- and short-run homogeneity and symmetry cannot be rejected at the 5% significance level. This is also true for the hypothesis of long- and short-run null cross-price effects between the equations of France and Portugal.

These results have important implications for the modelling and prediction of consumer behaviour. They suggest that knowledge of the way in which consumers adapt their demand behaviour to changes in its determinants requires more than a static system of long-run structural relationships. They also indicate that, for obtaining comprehensive information on the error-correction mechanism triggering the process of adjustment, it may

[7] The constraint of equal adjustment coefficients across equations is integrated in all subsequent restricted models derived from the general dynamic structure. Therefore, in all the hypotheses tested, this constraint holds previously. Consequently, the number of degrees of freedom for the χ^2 statistic includes this first restriction.

not be sufficient simply to introduce trend factors in the usual static AIDS formulations. Neither does it appear sufficient to choose dynamic models representing specific theories of short-run correction such as, for example, the partial adjustment model. A more general dynamic structure seems to be required to match data and theory in a consistent way.

6.4. EMPIRICAL RESULTS AND THEIR INTERPRETATION

The general dynamic model in (6.5) is estimated with SUR method. Table 6.3 presents the estimation results. This table shows the estimates of the coefficients (asymptotic t-values in brackets) for the share equations of Portugal (PT), Spain (SP) and France (FR) obtained with the unrestricted (6.5), $(H+S)^0L$ and $(H+S)^0LS$ models. The model labelled "unrestricted" considers the sole constraint of equal adjustment coefficients across equations. The symbols *, °and ' represent, respectively, the 1%, 5% and 10% significance levels. Goodness of fit indicators such as the residual sum squares (RSS), equation log-likelihood (ELL) and system log-likelihood (SLL) values, are also presented.

The statistical robustness of the share equations specifications is shown in table 6.4 which gives the Lagrange Multiplier (LM) version and the F version of diagnostic tests for serial correlation, functional form, error normality and heteroscedasticity, performed on an equation-by-equation basis for the unrestricted model. The p values for these statistics are given in brackets. The diagnostic test statistics in Table 6.4 show that the equations of the dynamic AIDS system are well-defined, statistically robust specifications. However, taking the unrestricted system as an whole, the explanatory power of the variables included (measured by the F test of overall significance) and the goodness of fit (measured by the adjusted R^2 coefficient) do not present brilliant results, particularly for the share equation of Portugal. This should not be surprising, given the individual non-significance of many of the coefficients (six out of twelve) in this equation. A solution to this problem, as suggested by the 'general-to-specific' practice, would be to exclude the insignificant regressors. Indeed, the imposition of theoretical constraints and null cross-price effects between the equations of France and Portugal generally improves the individual significance of the estimates. Therefore, the imposition of additional null restrictions on insignificant coefficients would be expected to strengthen further the statistical meaning of the dynamic system.[8] However, this would contradict the spirit of *general* dynamics which underlies the implementation of this model and, consequently, would omit information which we think is important to incorporate. Therefore, we do not impose further zero restrictions.

Table 6.3: Estimation results for the unrestricted dynamic model and models I and II

	Unrestricted			$(H+S)^0L$			$(H+S)^0LS$		
	PT	SP	FR	PT	SP	FR	PT	SP	FR
INT	0.0995 (4.50)•	0.2330 (2.66)•	0.4599 (7.41)•	0.0695 (5.41)•	0.2618 (3.25)•	0.4265 (7.94)•	0.0671 (5.93)•	0.2821 (4.07)•	0.4299 (7.49)•
PP_{t-1}	-0.0238 (-0.82)	-0.0479 (-0.63)	0.0717 (0.97)	-0.0494 (2.43)°	0.0494 (2.43)°	0 (none)	-0.0472 (-2.52)°	0.0472 (2.52)°	0 (none)
PS_{t-1}	0.0899 (2.69)°	-0.5036 (-5.73)•	0.4137 (5.16)•	0.0494 (2.43)°	-0.4486 (-5.72)•	0.3992 (5.38)•	0.0472 (2.52)°	-0.4545 (-5.06)•	0.4074 (4.79)•
PF_{t-1}	-0.0568 (-1.22)	0.5450 (4.08)•	-0.4882 (-3.79)•	0 (none)	0.3992 (5.38)•	-0.3992 (-5.39)•	0 (none)	0.4074 (4.79)•	-0.4074 (-4.79)•
FE_{t-1}	-0.0163 (-1.47)	0.0644 (2.23)°	-0.0481 (-1.81)'	-0.0038 (-0.50)	0.0359 (1.63)	-0.0322 (-1.60)	-0.0034 (-0.46)	0.0476 (2.60)°	-0.044 (-2.71)°
SE_{t-1}	-0.0039 (-0.82)	0.0295 (2.41)°	-0.0256 (-2.30)°	0.0018 (0.67)	0.0188 (1.81)'	-0.0207 (-2.13)°	0.0026 (1.04)	0.0177 (1.96)'	-0.0204 (-2.38)°
D_{t-1}	-0.0152 (-2.34)°	-0.0172 (-1.15)	0.0324 (2.24)°	-0.0145 (-2.37)°	-0.0118 (-0.87)	0.0263 (2.17)°	-0.0136 (-2.28)°	-0.0271 (-1.77)'	0.0407 (2.92)•
Wi_{t-1}		0.792 (7.45)•			0.758 (7.94)•			0.779 (7.78)•	
ΔPP_t	-0.1013 (-2.10)°	0.1010 (0.77)	0.0003 (0.00)	-0.0700 (-1.58)	0.0512 (0.39)	0.0188 (0.16)	-0.0569 (-1.55)	0.0569 (1.55)	0 (none)
ΔPS_t	0.1152 (1.82)'	-0.1506 (-0.92)	0.0354 (0.23)	0.0416 (0.96)	0.0286 (0.23)	-0.0702 (-0.62)	0.0569 (1.55)	-0.1761 (-1.80)'	0.1192 (1.33)
ΔPF_t	-0.0261 (-0.61)	0.2557 (2.25)°	-0.2296 (-2.13)°	0.0076 (0.22)	0.1500 (1.76)'	-0.1576 (-2.04)'	0 (none)	0.1192 (1.33)	-0.1192 (-1.33)
ΔE_t	0.004 (0.19)	0.1057 (1.89)'	-0.1017 (-2.04)'	-0.0069 (-0.33)	0.1201 (2.18)°	-0.1132 (-2.31)°	0.0020 (-0.12)	0.0636 (1.34)	-0.0656 (-1.52)
RSS	0.001	0.009	0.008	0.001	0.010	0.008	0.002	0.014	0.012
ELL	99.42	72.65	74.82	98.13	71.82	74.40	97.95	66.62	69.00
SLL		174.24			172.54			167.04	

[8] We did estimate the model under the full set of long- and short-run restrictions and additional zero constraints on the coefficients of the variables FE_{t-1} and ΔE_t in the equation for Portugal. The overall statistical quality of the model increases, and all the long-run coefficients become more significant and have the correct signs. However, the short-run coefficients do not alter much, either in their magnitudes or individual significance.

Table 6.4: Diagnostic tests for the equations of the dynamic AIDS model

Equation	DIAGNOSTIC TESTS							
	Serial Correlation		Functional Form		Normality		Heteroscedasticity	
	LM version	F version	LM version	F version	LM version	F version	LM version	F version
PT	0.05 (0.83)	0.02 (0.88)	0.00 (0.97)	0.00 (0.98)	0.92 (0.63)	na	1.77 (0.18)	1.76 (0.20)
SP	0.40 (0.53)	0.04 (0.85)	4.63 (0.03)	2.98 (0.11)	0.58 (0.75)	na	0.07 (0.79)	0.06 (0.80)
FR	0.11 (0.75)	0.06 (0.81)	2.50 (0.11)	1.47 (0.24)	1.97 (0.37)	na	0.02 (0.88)	0.02 (0.88)

For comparison purposes with the AIDS model of chapter 5, the interpretation of the estimation results focuses on the model denoted $(H+S)^0L$, for which constraints are imposed only on the long-run coefficients. Although the model under the full set of long- and short-run restrictions seems statistically more robust, model $(H+S)^0L$ presents sufficient statistical quality to be considered a reliable means to interpret the long- and short-run behaviour of the UK tourism demand in France, Spain and Portugal.

6.4.1. Interpretation of the elasticities' estimates

The interpretation of the long-run coefficients of the dynamic AIDS model is similar to that given in chapter 5, for the coefficients of the static AIDS model. However, the short-run information provided by the former has no correspondence in the estimates obtained from the latter. Hence, a thorough analysis of the dynamic model coefficients, accompanied by a comparison of long- and short-run estimates is worthwhile. This analysis is carried out on an equation-by-equation basis, leaving comparison across equations to be dealt with latter, when interpreting the elasticities' estimates.

The adjustment velocity estimate is 0.76 for all share equations, as imposed by the corresponding restriction. This estimate suggests a rapid adjustment of UK tourism demand to the long-run equilibrium, when changes in the demand determinants occur. Indeed, 76% of that adjustment is attained in the current period, and only 24% is postponed to the next period. This corroborates the idea of almost perfect information, quickly circulating among UK tourists, concerning aspects which may influence their decision to visit France, Spain or Portugal.

Since by construction of the model, the velocity of adjustment parameter (λ) is multiplied by the intercept and long-run parameters of all share equations, to obtain the actual estimates of the long-run coefficients we have to divide the coefficients of the lagged variables by the estimate of λ (0.76). As a result, the actual long-run estimates of, say, the dummy variable (D_{t-1}) coefficients, in the equations for Portugal, Spain and France are, respectively, -0.019, -0.016 and 0.034.

In general, all coefficients' signs are consistent with theoretical expectations. For instance, all coefficients of the expenditure variable, when statistically significant, have the expected signs both in the long- and short-run and in the first and second periods. In the share equation for Spain, these coefficients are all positive, indicating an elastic response of the UK demand for Spain to changes in tourism budget. Conversely, in the share equation for France, these coefficients are all negative, indicating an inelastic response of the UK demand for France to changes in the UK tourism budget. In the case of Portugal, none of these coefficients is statistically significant.[9] Moreover, when significant, both the long- and short-run own-price coefficients are negative, as expected with normal commodities, and the cross-price coefficients are positive, as expected from destinations which are competitors rather than complements. The results also indicate that the political and economic events of 1974-1981 had a negative effect of the UK tourism demand for Spain and Portugal and a positive effect for France. Overall, the magnitudes and signs of the long-run estimates obtained from the dynamic model accord with the corresponding estimates obtained from the static model in chapter 5. Therefore similar values for the long-run elasticities estimates should be expected.

For all equations, the short-run coefficients are, generally, statistically insignificant. This may indicate that the effects on the UK tourism demand, induced by short-run changes in its determinants, are not of relevant magnitude. Supporting this hypothesis are the statistical robustness of the static model and that of the dynamic model despite its short-run insignificance, the consistency of the long-run estimates provided by both models, and the high adjustment velocity of UK demand to changes in tourism budget and prices.

A more detailed analysis of the results requires the relevant elasticities values which may confirm the similarities of the long-run information obtained from the AIDS model of

[9] The non-significance of a given coefficient does not necessarily imply the non-significance of the corresponding elasticity as the formulae for its calculation may include other coefficients as well as the average and/or the base year shares.

chapter 5 and the dynamic model $(H+S)^0L$. As before, we compute the expenditure and uncompensated price elasticities using the coefficients' estimates of the relevant dynamic specification, and the formulae and shares' values given in chapter 5. Table 6.5 shows these elasticities estimates and respective t values in brackets, for the AIDS model denoted $(H+S)^0$ in Chapter 5, and for the dynamic AIDS model denoted $(H+S)^0L$ in this chapter.

Table 6.4: Expenditure and uncompensated own- and cross-price elasticities

		Expenditure elasticities		Own-price elasticities		Cross-price elasticities					
						PP		PS		PF	
		First Period	Second Period	First Period	Second Period	First Period	Second Period	First Period	Second Period	First Period	Second Period
PORTUGAL	Static $(H+S)^0$ LONG-RUN	0.820 (4.51)	0.947 (10.16)	-2.237 (-5.46)	-1.797 (-6.87)	X	X	1.344 (2.81)	0.830 (2.77)	0.073 (0.16)	0.019 (0.06)
PORTUGAL	Dynamic $(H+S)^0L$ LONG-RUN	0.913 (5.22)	1.027 (26.22)	-2.128 (-4.61)	-1.729 (-5.93)	X	X	1.176 (2.62)	0.714 (2.39)	0.039 (0.49)	-0.011 (-0.69)
PORTUGAL	Dynamic $(H+S)^0L$ SHORT-RUN	0.911 (3.35)		-1.895 (-3.40)		X		0.573 (1.01)		0.138 (0.28)	
SPAIN	Static $(H+S)^0$ LONG-RUN	1.203 (25.23)	1.150 (26.97)	-1.817 (-12.94)	-1.933 (-11.74)	0.097 (2.56)	0.124 (2.74)	X	X	0.517 (4.69)	0.658 (4.96)
SPAIN	Dynamic $(H+S)^0L$ LONG-RUN	1.076 (21.39)	1.047 (36.27)	-1.980 (-10.82)	-2.150 (-10.12)	0.097 (2.30)	0.120 (2.41)	X	X	0.807 (5.40)	0.983 (5.38)
SPAIN	Dynamic $(H+S)^0L$ SHORT-RUN	1.213 (12.38)		-1.048 (-4.56)		0.071 (0.32)		X		0.171 (1.02)	
FRANCE	Static $(H+S)^0$ LONG-RUN	0.630 (7.25)	0.808 (24.52)	-2.039 (-9.25)	-1.901 (-10.44)	0.033 (0.41)	0.017 (0.25)	1.376 (5.46)	1.077 (5.30)	X	X
FRANCE	Dynamic $(H+S)^0L$ LONG-RUN	0.865 (9.65)	0.929 (25.34)	-2.608 (-8.76)	-2.336 (-9.36)	0.012 (1.50)	0.007 (1.93)	1.730 (4.98)	1.400 (5.05)	X	X
FRANCE	Dynamic $(H+S)^0L$ SHORT-RUN	0.685 (5.02)		-1.298 (-5.45)		0.082 (0.25)		-0.050		X	

The long-run elasticities obtained from the dynamic model are similar in magnitude and signs to those obtained from the AIDS model of chapter 5. Therefore, the discussion and comments about these elasticities estimates provided then apply, in general terms, to the long-run elasticities estimates obtained from the dynamic model. Indeed, the latter estimates not only present values similar to those obtained from the former but also behave in similar ways (increasing or decreasing), in the first and second periods. This should be expected as the AIDS model of chapter 5 is a "trended", "apparently-dynamic" specification which seems to allow for the correction of omitted temporal factors through the addition of a trend variable and the consideration of a non-constant expenditure coefficient. However, if comparison were made using an orthodox static AIDS model for the same data sample, the results would show this specification not to be compatible with the data or consistent with demand theory restrictions. This result was suggested by the rejection of the orthodox AIDS model when tested against the "unorthodox" form of chapter 5, and further supported by the rejection of the same model against the general dynamic model adopted in this chapter.

Nevertheless, the examination of the short-run elasticities estimates and a comparative analysis of short- and long-run demand behaviour are pertinent within the general dynamic framework. The analysis will focus on the second period of the sample (the last two decades), given that it relates to more recent behaviour of the UK tourism demand. While the estimates of the expenditure elasticities in the equation for Portugal are close to unity in both the short- and in the long-run, the corresponding estimates for the shares of Spain and France present significant differences in their short- and long-run magnitudes. The long-run expenditure response for France is close to unity but in the short-run, the UK demand for France is clearly inelastic. In the equation for Spain, the long-run response is also close to unity, but the short-run response is clearly elastic.

The estimate for Spain has the lowest value of all short-run own-price elasticities. This indicates that UK tourists seem to be less sensitive to short-term price changes in Spain than in France or Portugal. This information, supplemented with the fact that Spain presents the highest estimate for the short-run expenditure elasticity, suggests Spain as a primary destination for UK tourists in the short-run. Hence, Spain may have a comparative advantage in relation to its competitors and a wider scope for manoeuvre concerning policies involving short-term price changes. However, decisions in this area should not overlook the increasing sensitivity of UK demand to long-term changes in Spanish prices as compared with its decreasing sensitivity towards identical changes in France and Portugal.

The UK tourism demand for Portugal seems to be more sensitive to short-run own-price changes in this destination than in France or Spain. Indeed, the own-price elasticity estimates present their largest short-run value in the equation for Portugal. However, the long-run elasticity estimates for the equations of all destinations are close to -2 which indicates that, in the long-run, UK tourists are highly sensitive to price changes and will penalise, in a similar way, any of the destinations for an increasing-price policy.

The inferences concerning the cross-price effects drawn from the AIDS model of chapter 5 apply, in general terms, to the long-run estimates of the dynamic model. Indeed, the lack of sensitivity of the UK demand for tourism in France (Portugal) to price changes in Portugal (France), the decreasing response of the UK demand for tourism in France and Portugal to price changes in Spain and the increasing response of the demand for Spain to price changes in France or Portugal are common features of the estimation results provided by both models. However, the dynamic model permits the analysis of short-run cross-price elasticities not possible with the AIDS approach of chapter 5.

In particular, it is worth noting an interesting feature of the short-run cross-price elasticities when compared with their corresponding values in the long-run. None of the short-run elasticities is statistically significant at the 5% level. In contrast, their long-run counterparts are all significantly different from zero except, of course, in the case of France versus Portugal. These results indicate that UK tourism demand for one destination, does not respond significantly to short-run price changes in another, while in the long-run UK tourists seem to be able to compare prices across destinations and adapt their preferences accordingly. Put another way, in the short-run, the ability of UK demand to respond significantly to cross-price changes in competing destinations is immaterial, whereas in the long-run UK tourists seem fully aware of price changes across destinations and adapt their demand with significant effects for the destinations considered. Hence, destinations are more likely to retain their tourism receipts if they are able to avoid long-run increases in their own-prices and to maintain any adverse price changes from competitors within short-run periods.

6.5. CONCLUSION

The goal of this chapter is to point out areas where the analysis of tourism demand within an AIDS system approach is likely to contribute new aspects of theoretical and empirical relevance. The development of dynamic generalisations of traditional systems is an important element of recent research work in demand behaviour analysis. The motivation for such generalisations has largely derived from the lack of accordance between utility theory constraints and static demand models. Static specifications typically rule out theoretically plausible features involving short-run dynamics of demand adjustment. Hence, the conflict observed between theory and empirical evidence has generally been attributed to dynamic misspecifications within static approaches. Mis-specified econometric models give rise to unreliable estimation results, invalid statistical inference and inaccurate forecasting procedures. Flawed estimates produce misleading economic analysis which may induce inappropriate policy measures.

These problems, however, have been largely ignored in tourism demand research. Dynamic generalisations of demand modelling are a rare feature in empirical studies of tourism, and demand analysis in this context has generally been based on quantitative models which do not comply with the necessary pre-requisites of theoretical and statistical quality. Nevertheless, the possibility of knowing how the allocation of tourism expenditure evolves over time, and how tourists adjust their demand behaviour in order to achieve a steady state equilibrium is of considerable interest for tourism analysis and policy making.

The unorthodox AIDS system estimated in chapter 5, has specification features which allow for dynamic-like elements to be incorporated in its share equations. As a consequence, and in contrast with the orthodox static system, this 'seemingly-dynamic' model reveals itself consistent with the data and with the utility maximisation assumptions of consumer theory. However, information about the mechanism underlying the short-run adjustment process cannot be accessed within this approach and a clear separation between short- and long-run effects cannot be made. Hence, an explicit dynamic specification was found to be the appropriate means to obtain reliable estimates of both the long- and short-run responses of UK tourism demand to changes in its determinants.

In this chapter we estimated a flexible general dynamic form of the AIDS system. The estimation results show this model to be data coherent and theoretically consistent providing empirical evidence of the robustness of this methodology for conducting tourism demand analysis in a temporal context. Moreover, the dynamic model offers dependable evidence on both the capacity of the 'seemingly-dynamic' AIDS to provide reliable long-run information, and on the inadequacy of the orthodox AIDS and other restricted specific dynamic models, to reconcile consistently data and theory within their formulations.

The results of this extensive modelling exercise for the UK tourism demand are encouraging and indicate directions for future research. For instance, the results show that an appropriate dynamic specification does matter when modelling systems of equations, and can have considerable impact on the results of tests concerning the validity of theoretical hypotheses. Moreover, the general flexible form of the dynamic AIDS model allows for its testing against more specific demand formulations. Indeed, if models prove to be data-consistent, they are subject to further testing for evaluation of their consistency with the postulates of utility theory. The general dynamic AIDS model passes all the tests providing estimation results which are statistically robust, empirically plausible and theoretically consistent.

In dynamic specifications, the utility theory constraints are generally tested for the long-run coefficients. The motivation for testing theory restrictions, preferably in the steady-state, rests on the idea that, in the short-run, consumers may not have fully adjusted to changing circumstances and, hence, homogeneity and symmetry may not be observed in short-run behaviour. Given that the general dynamic structure of the AIDS model is not rejected by the data, the inherent implication is that tourists adjust their behaviour to changes in their demand determinants with a lag. That is, the adjustment process takes more than the current period to be accomplished fully. Hence, if tourists do not adjust instantaneously, then homogeneity and symmetry are not expected to hold in the short-run. When tested, however, this hypothesis could not be rejected, suggesting that the rationality of utility maximisation postulates is observed in the behaviour of UK tourists' demand for France, Spain and Portugal, both in the long- and in the short-run.

At this point, we think it would be interesting to call upon the findings Anderson and Blundell (1984) who, when confronted with a similar situation (although still rejecting the hypothesis in the short-run) make the following comment: "*with homogeneity and symmetry imposed on the long-run coefficients, ... short-run homogeneity produced a surprising result since the test statistic of 15.42 implies only a marginal rejection. The consideration of this and further restrictions on short-run behaviour would seem a fruitful area for future research*".

In our case, the 'surprising' result of homogeneity and symmetry holding in the short-run is not unexpected. The general dynamic model seems to be a sufficiently robust specification to track accurately the UK demand behaviour over the sample period. Thus, we can be fairly sure of the statistical reliability of the information provided by this model.

Its estimates suggest that UK tourists adjust very fast to changes in their demand determinants. As a consequence, we should expect non-significance and/or irrelevant magnitudes for the short-run coefficients. Indeed, that is the general indication of the estimates provided by the dynamic AIDS model. With small or insignificant short-run coefficients, the statistical process by which constraints are imposed on the model leads to an 'easy' non-rejection (a possible under-rejection) of the hypothesis tested. Therefore, the faster consumers adjust their demand behaviour, the less significant short-run effects are, and the likelier is the non-rejection of utility theory postulates imposed on the short-run. This hypothesis requires, of course, further empirical support which can only be delivered in the context of future research.

The empirical results provided in this chapter show how the estimation of a dynamic AIDS system can provide new information about the behaviour of UK tourism demand. This modelling approach allows for intertemporal rationality of consumer behaviour by explicitly considering the mechanism underlying the short-run adjustment process. Estimates for tourism price and budget elasticities were obtained, permitting a comparative analysis of, on the one hand, the relative ability of the AIDS system to provide accurate information on the long-run behaviour of UK tourists and, on the other hand, the relative magnitudes and statistical relevance of long- and short-run sensitivity of the UK tourism demand to changes in its determinants.

Nevertheless, there are theoretical and empirical issues which still have to be addressed to endorse the AIDS models as quality specifications. One of the most important regards the spurious regression problem; since the AIDS systems include nonstationary time series, their estimation results can be spurious unless the variables are cointegrated. Others, are linked with the *a priori* division of endogenous/exogenous variables assumed by the AIDS approach. Finally, an important matter in quality evaluation of econometric models is their forecasting ability. Statistical models can be good means to describe long-run economic relationships but if they are not equally good forecasters, they lose much of their relevance for policy analysis purposes. We address these issues in the next chapters: chapter 7 involves cointegration analysis within a vector auto-regressive framework of the long-run relationships defined within an AIDS system; chapter 8 investigates the ability of the different econometric models previously estimated to predict the UK tourism budget shares of France, Spain and Portugal.

CHAPTER 7

VECTOR AUTOREGRESSIVE MODELLING OF THE UK DEMAND FOR TOURISM

7.1. INTRODUCTION

Static, steady-state equilibrium models are often termed long-run specifications. In contrast, dynamic short-run models are often linked with the concept of disequilibrium as a process of adjustment. Thus, inherent to the distinction between short- and long-run models is the notion of equilibrium. The long-run is a state of equilibrium where change is not likely to occur since economic forces are in balance. However, as pointed out by Harris (1995, p.5), "*there is no necessity actually to achieve equilibrium at any point in time, even as $t \rightarrow \infty$. All that is required is that economic forces move the system towards the equilibrium defined by the long-run relationship posited. ... Thus, what matters is the idea of a steady-state relationship between variables which are evolving over time.*" This idea is a central aspect of theoretical and empirical economic analysis and the estimation of long-run relationships has been a main concern in applied econometric research.

Frequently, the estimation of long-run relationships involving time series brings about problems that arise from the presence of non-stationary variables in the model. Ignoring them and proceeding with the estimation of a regression containing unit-roots in the data can lead to spurious results. The spurious regression problem is reflected by estimation results suggesting the existence of statistically significant long-run relationships between variables when, in fact, all that is shown is evidence of contemporaneous correlations due to common, but unrelated, trends in the nonstationary series. Simple differentiation to remove unit roots from the data is not the answer since, while avoiding the spurious regression problem, it also removes long-run information which is crucial to establish the existence of meaningful causal relations among the variables' levels.

The spurious regression problem leads to the following question: "is it possible to infer a genuine long-run relationship between non-stationary variables"? The affirmative answer is given in Engle and Granger (1987) through the concept of cointegration. The economic interpretation of cointegration states that if two or more time series variables are linked to form a long-run stationary equilibrium relationship then, even if the variables themselves are non-stationary, they will nevertheless present a meaningful co-movement over time, owing to underlying equilibrating forces. These forces are represented by an error-correction mechanism, which compels the economic system to converge towards its long-run equilibrium. Thus, the concept of cointegration portrays the existence of a short-run dynamic equilibrating mechanism which pushes the variables' levels towards their long-run equilibrium path, even if these variables are nonstationary. In other words, if two or more variables are cointegrated, then there must be an error correction mechanism underlying their co-movements and, conversely, the existence of an error-correction mechanism generates cointegrated time series.

In chapter 4, we addressed the importance of distinguishing between stationary and non-stationary variables, since failure to do so could lead to the problem of spurious regression. We tested for the presence of unit roots in the time series data entering the model and obtained indications of their presence. When dealing with nonstationary variables, the concept of cointegration is synonymous with the concept of long-run equilibrium. Failure to establish cointegration often means the non-existence of a steady state relationship among the variables. Therefore, the next step was to find out if a long-run relationship existed, linking the variables of the dynamic single equation ARDL model specified. From this model, information on both short- and long-run relationships could be retrieved. We tested for cointegration using the residuals of the long-run specification and obtained indications that the variables were cointegrated. The linkage between the error-correction mechanism and the long-run equilibrium relationship could also be established.

However, there are several disadvantages in the use of a single equation approach. In the single equation framework used in chapter 4, we started by modelling three separate equations which specified the UK tourism demand for France, Spain and Portugal as a function of tourism prices and the UK real per capita income. However, in this specific tourism demand context, it is possible that a shock affecting the UK demand for tourism in France also affects tourism in Spain, and a revolution in Portugal influence tourists' decision to visit this or neighbouring countries. Yet, the single equation approach does not explicitly account for the possible existence of interrelationships between destinations and cross-equation restrictions cannot be tested within this approach.

Furthermore, despite the statistical evidence of cointegration in the single equations specified in chapter 4, several problems still remain within this econometric approach. When there are n > 2 variables in an equation, there can be up to n-1 linearly independent cointegrated relationship each corresponding to a distinct long-run equilibrium. Only when n=2 is it possible to infer that the cointegrating vector found is unique. If this is not the case, adopting a single equation approach is inefficient since we can only obtain one linear combination of the existing cointegrating vectors and not valid information about *all* the possible long-run relationships existing in the model.[1] Moreover, even if there is only one cointegrating relationship, the single equation approach is still inefficient, unless all the right-hand side variables are weakly exogenous. If there are endogenous variables as regressors, information will be lost causing inefficiency. However, this information is not wasted if, instead, a system of equations is estimated allowing each endogenous variable to appear as a dependent variable. Therefore, the results estimated from a single equation approach involving nonstationary time series are reliable only if there is a unique cointegrating vector among the variables and all regressors are (weakly) exogenous.

In chapter 5 we estimated the UK demand for tourism in France, Spain and Portugal using a static AIDS system of equations. A dynamic version of this model was estimated in chapter 6. However, the methodological framework of the AIDS approach does not contemplate cointegration analysis and the assumptions underlying its specification include an assumed endogenous-exogenous division of variables that may be questionable. The AIDS approach of chapters 5 and 6 models the UK demand for tourism as a system of equations, expressing the destinations' expenditure shares as functions of exogenous variables which include the UK per capita real tourism expenditure and own and competing tourism prices. Although there is some theoretical basis justifying the assumption of exogeneity for the real per capita expenditure, there is the possibility of feedback effects between prices and the demand for tourism. Yet, the AIDS approach does not consider this possibility. In addition, the tourism price variables included in the AIDS models were tested and found nonstationary in chapter 4. Hence, the estimation results obtained from

[1] As shown in chapter 4, we can obtain valid information about these issues with the Pesaran and Shin approach. However, if the tests indicate the presence of more than one cointegrating vector, the ARDL single equation is inefficient.

these models can be deemed spurious and the statistical inference invalid, if no cointegrated relationship(s) are found linking the variables of these specifications.

Consequently, if data series are nonstationary, there seems to be a risk involved in the estimation of econometric models which regress endogenous variables on several assumed exogenous variables, without sanctioning their statistical validity with cointegration analysis. Given that the number of potential cointegrating vectors is unknown, and given the possibility of simultaneous determination of the variables, empirical analysis must go one step further and establish a methodology to specify econometric models which can be efficiently estimated and validly tested within a system of equations approach. Summarising, in the presence of non-stationary time series and feedback effects, an efficient econometric approach for estimating long-run relationship(s), must be a system of equations which allows all potentially endogenous variables to appear as dependent variables and for appropriate cointegration analysis.

An econometric methodology with these features was proposed by Sims (1980) and developed by Johansen (1988). Sims' vector autoregressive (VAR) approach establishes innovative specification methods that are considered to be valid alternatives to both the single equation and the traditional structural multi-equation approaches. The Johansen (1988) procedure adds cointegration analysis and an efficient estimation method of the structural parameters needed to establish the number of long-run relationships and to supply information on both the short- and long-run responses.

In this chapter we use the VAR methodology and the full information likelihood system approach, developed by Johansen (1988), to analyse the UK tourism demand for France, Spain and Portugal. The chapter is structured as follows. Section 7.2, addresses the main features of Sims' methodology. Section 7.3, establishes the order of integration of the time series and the appropriate lag-length for the variables included, specifies the (unrestricted) VAR model for the UK demand for tourism and presents the estimation and forecasting results obtained with this specification. Section 7.4, applies the Johansen procedure for determining the number of cointegrating vectors existing in the specification and presents the estimation results of a cointegrated structural VAR under exactly- and over-identifying restrictions. Section 7.5, presents the forecast results obtained with the cointegrated VAR and compares them with the forecasts of the unrestricted specifications. Section 7.6 concludes.

7.2. MAIN FEATURES OF THE VECTOR AUTOREGRESSIVE METHODOLOGY

In a single equation model, the emphasis is on estimating the average value of the dependent variable (say, Y) conditional upon the fixed values of the explanatory variables (say, X's). The cause-effect relationship in such models is therefore assumed to run from the explanatory variables to the dependent variable. Hence, in these models, the X's are assumed to be exogenous variables and Y is assumed to be endogenous.

Although in some economic contexts this assumption can be justified, it is often controversial. In many situations such unidirectional cause-effect relationship is meaningless, particularly in cases where feedback effects are present and the "independent" variables are correlated with the error terms. This occurs when Y is determined by the X's and some of the X's are determined by Y. In this case, a feedback mechanism is operating, making these variables jointly or simultaneously determined. In such cases, there is a bi-directional cause-effect relationship between the Y and some of the X's which renders the distinction between "dependent" and "independent" variables dubious. Moreover, the observed simultaneity of Y and some of the X's has important consequences for the entire process of empirical analysis. It can be shown that the endogenous variables on the right-hand side of a single equations are correlated with the error terms violating one of the crucial assumptions of the least squares estimation method, namely, that the X variables are either non-stochastic or if stochastic, are distributed independently of the disturbance term. If neither of these conditions is met, the least-squares estimators are not only biased but also inconsistent and the statistical inference based on such estimators is invalid.

This flaw of the single-equation approach, resulting from the correlation of the error terms with one or more of the explanatory variables, is known as simultaneous equation bias. This problem was tackled by theoretical advances in econometrics during the 1950s and 1960s which were mainly concerned with the development of appropriate structural specifications and estimation methods that could take into consideration the simultaneous determination of a set of variables within the same equation. These theoretical efforts gave rise to what is known in the literature as the traditional structural multi-equation system approach. This approach is described in many textbooks, for example, Johnston (1984), Griffiths *et al.* (1993) and Gujarati (1995) and only the main features will be outlined here.

In contrast with the single equation approach, the structural multi-equation methodology separates the set of variables that are viewed as being simultaneously determined, from the remaining set of regressors. Such models specify one equation for each of the jointly dependent or endogenous variable, forming a multi-equation system of structural relationships. Hence, these models can accommodate the endogenous-exogenous dichotomy of the variables and further distinguish the set of endogenous from the set of predetermined variables, which includes exogenous and lagged-endogenous variables.

Once a system of equations is formulated, the traditional structural approach suggests several procedures, which can remove the problem of simultaneous bias. This problem is present in a structural system of equations because this specification expresses each endogenous variable as a separate function of other endogenous variables, along with predetermined (exogenous and lagged endogenous) variables and a stochastic error term. Since the endogenous explanatory variables are correlated with the error term, the coefficients associated with the structural system (structural coefficients), cannot be consistently estimated by least squares methods. However, this problem can be removed if the structural equations are solved for the existing endogenous variables, making them dependent solely on the predetermined variables and the stochastic disturbances.

A structural system solved for the endogenous variables is known as the reduced-form of the model, and the coefficients associated with this form are called reduced-form coefficients. Since only predetermined variables and stochastic disturbances appear in the right-hand side of the reduced-form equations, and since the predetermined variables are assumed not to be correlated with the error terms, the OLS estimation method applied to the reduced-form equations generates consistent and asymptotically efficient estimates. Hence, the simultaneous bias problem is removed from the reduced-form system of equations. However, the estimates obtained with this procedure are those of the reduced-form coefficients and not of the structural coefficients which are ultimately of interest. Since the latter are combinations of the former, the possibility exists that the structural coefficients can be retrieved from the reduced-form coefficients. Whether this is the case, brings about one of the major problems of the structural system approach: the problem of identification.

The identification problem arises because different sets of structural coefficients may be compatible with the same set of data; that is, a given reduced-form equation may be compatible with different structural equations, making it difficult to tell which particular model is being investigated. As an illustration, consider the following example of a simple demand-supply model:

Demand function: $Q_t^d = \alpha_0 + \alpha_1 P_t + u1_t$ (7.1)

Supply function: $Q_t^s = \beta_0 + \beta_1 P_t + u2_t$ (7.2)

Equilibrium condition: $Q_t^d = Q_t^s = Q$

where Q^d is quantity demanded, Q^s is quantity supplied, P is price of the commodity and α and β are parameters. It is easy to see that P and Q are jointly determined as a change in u1 will cause the demand curve to shift which, in turn, causes both P and Q to change. Similarly a change in u2 will shift the supply curve which, in turn, will affect P and Q. Hence, there is simultaneous equation bias in the system, as P and Q are correlated with both u1 and u2.

However, model (7.1-7.2) might be perfectly sensible from a theoretical point of view. Nevertheless, intuition would suggest that with only two time series data on Q and P and no additional information, it might be difficult to estimate four unknown parameters. Even if we ignore the intercepts by, for example, working with cantered variables, how can we estimate two structural parameters (α_1 and β_1) from a regression with only one explanatory variable? If we regress Q on P how can we know whether we are estimating the demand equation or the supply equation? The identification problem consists in seeking answers to these questions and it is apparent that these answers must precede the problem of how to estimate the structural parameters.

Another way of seeing the identification problem is to realize that it is impossible to retrieve the structural parameters from the reduced-form coefficients of a system. In this case the system is said to be under-identified. Using the demand-supply example above, we can illustrate this point.

By the equilibrium condition we obtain:

$$\alpha_0 + \alpha_1 P_t + u1_t = \beta_0 + \beta_1 P_t + u2_t \tag{7.3}$$

Solving for P we get the equilibrium price

$$P_t = \Pi_0 + v_t \tag{7.4}$$

Substituting P_t given by (7.4) in (7.1) we obtain the equilibrium quantity

$$Q_t = \Pi_1 + w_t \tag{7.5}$$

where $$\Pi_0 = \frac{\beta_0 - \alpha_0}{\alpha_1 - \beta_1} \quad \text{and} \quad \Pi_1 = \frac{\alpha_1\beta_0 - \alpha_0\beta_1}{\alpha_1 - \beta_1} \tag{7.6}$$

Equations (7.4) and (7.5) are reduced-form equations and Π_0 and Π_1 are reduced-form coefficients. It is clear from (7.6) that there is no way of recovering the four unknown structural coefficients from the two reduced-form coefficients.

However, equations (7.1) and (7.2) become identified if additional predetermined variables are added to the system. For example, if we add an exogenous variable to the demand equation (income or interest rate, for example) and the lagged endogenous variable P_{t-1} to the supply equation, we can retrieve the structural coefficients from the reduced-form coefficients. In this case, there are six unknown structural parameters and six reduced-form coefficients, enabling us to determine unique estimates for the structural parameters. In this case, the structural form is said to be exactly-identified.

Adding more variables to one and/or the other equation would cause the system to be over-identified and the structural parameters estimates not to be unique. In this case we would have too much supplementary information, not needed to exactly-identify the structural parameters. The "secret" for achieving identification, then, seems to be related to the adding of variables to an equation which do not appear in others. For these other equations, the coefficients attached to the "missing" variables are zero. Intuitively, it appears that identification is more likely if the matrix of the structural parameters contains several zero elements in specific rows and columns. In structural econometrics terminology, the presence of a zero element in this matrix is defined as a zero restriction. Without these restrictions, most of the structural systems would not be identified.

It seems to be the case, then, that much is left for the structural econometrician to decide prior to undertaking the actual estimation procedure of a structural system. First, in order to specify the system, it is necessary to establish which variables are endogenous and which are exogenous. Second, in order to achieve identification, a number of zero restrictions have to be imposed. The main criticism of multi-equation structural modelling has been centred on these *a priori* decisions: the role of the zero restriction assumption and the categorisation of the variables into exogenous and endogenous. In fact, some models seem to be formulated with variables added to some equations and deleted from others merely to achieve identification, and without much economic justification. On the other hand, plausible economic models lack empirical support because they are not identified.

These aspects of structural modelling led Sims (1980) to consider the zero restrictions as "incredible" and "haphazard" and to propose a radically different strategy for the specification and estimation of multi-equation systems. Sims (1980, p. 14-5) states:

"Because existing large models contain too many incredible restrictions, empirical research aimed at testing competing macroeconomic theories too often proceeds in a single- or few equations framework. For this reason alone, it appears worthwhile to investigate the possibility of building large models in a style which does not tend to accumulate restrictions so haphazardly. ...It should be feasible to estimate large-scale macro-models as unrestricted reduced-forms, treating all variables as endogenous".

Sims' (1980) new methodology of vector autoregressive (VAR) modelling differs from the traditional structural approach in three main aspects: all variables are treated as endogenous; no zero restrictions are imposed; the VAR model is not based on any specific economic theory. As pointed out by Charemza and Deadman (1997), the third aspect derives from the other two since, "*if there is no variable excluded from any equation of the model and nothing is exogenous, it means that everything causes everything else and there is no room for assuming much more than very general economic principles as a starting point* (p.157)." Indeed, this is one of the major criticisms attached to Sims' approach since, as Pesaran (1997) underlines, much of the long-run analysis within this 'purely-statistical' approach "*is conducted without providing an explicit account of the type of equilibrium theory that may underlie it*" and that "*empirical applications of these methodology have focused on the statistical properties of the underlying economic time series, often at the expense of theoretical insights and economic reasoning*" (p. 178). Hence, Sims' methodology is often labelled by its critics, as an a-theoretical approach to long-run equilibrium analysis.

The features which make the VAR approach so flexible, easy to specify and simple to estimate also mark the area within which weaknesses remain. Unlike the traditional structural systems, an unrestricted VAR model is a-theoretical because it does not use any *a priori* information supplied by economic theory. The formulation of a general (unrestricted) VAR is based on a system of equations, regressing each variable of the model on its own and all the other variables' lagged values. A reduced-form VAR[2] requires little more than the choice of appropriate variables to include in the model and the determination of a suitable lag-length. With so little theoretical input, one should not expect more than little economic content in the output. In fact, unless the underlying structural model can be

[2] The right-hand side of all equations in a general VAR contains only predetermined (endogenous lagged) variables. This form of the model is considered to be similar to the reduced-form of a structural multi-equation system with no exogenous variables. Therefore, a general (unrestricted) VAR specification is also known as reduced-form VAR.

identified from the reduced-form VAR, the economic interpretation of the estimates is difficult, if not impossible. On the other hand, the VAR is a truly simultaneous system in that all variables are regarded as endogenous and no *a priori* endogenous-exogenous division is assumed.

However, the biggest practical challenge in a VAR specification is to choose the appropriate lag-length for its variables. The longer the lag-length, the faster degrees of freedom are eroded. If the lag-length is p and there are n variables, each of the n equations contains n.p unknown parameters plus the intercept. In view of the limited number of observations generally available in most empirical analysis, the introduction of several lags for each variable can be a problem. Nevertheless, appropriate lag-length selection is crucial. Since a VAR contains lagged dependent variables as regressors, if autocorrelation is present in the error terms, the predetermined variables on the right-hand side of the equations can be correlated with the error terms, leading to inconsistent estimators. So, purging the error terms from autocorrelation is essential in VAR modelling and lagging all variables a certain number of times can do this. Yet, if p is too small, the model can be mis-specified; if p is too large, the model can be over-parameterised and degrees of freedom wasted. In addition, the Johansen procedure, which establishes the number of cointegrating vectors existing in a VAR, can be quite sensitive to the lag-length chosen for the VAR structure. Therefore, a great deal of attention should be paid in defining p.

When the appropriate lag-length is imposed, the error terms of each equation can be assumed to be serially uncorrelated. Since a VAR expresses the current values of each endogenous variable as a function solely of predetermined (lagged endogenous) variables,[3] these are not correlated with the serially uncorrelated error terms, and each equation in the system can be estimated by OLS. In this case, the OLS method provides consistent and asymptotically efficient estimates.[4]

There is an important difference between using an unrestricted VAR model for forecasting and using it for economic analysis. If economic analysis is the objective, knowledge of the structural parameters is important. However, an unrestricted VAR model will be under-identified and there is no way (without imposing restrictions), in which the structural parameters can be recovered from the reduced-form VAR. On the other hand, a

[3] An additional set of deterministic components, such as an intercept, deterministic trend and seasonal dummy variables is often added to the equations of a VAR model.
[4] If the appropriate lag-length is not the same for all variables, or if the equations include different regressors, SUR is more efficient than OLS for estimating a VAR.

VAR model can be viewed as a reduced-form system with no exogenous variables specially adapted for forecasting purposes. In fact, if the objective is forecasting, the underlying structural form of the model does not need to be estimated, as its parameter estimates are irrelevant in the process of obtaining forecasts. In this case, even being under-identified, an appropriately specified (unrestricted) VAR supplies forecasts which are unbiased and have minimum variance. In addition, unlike the standard econometric forecasting procedures where forecasts must be conditioned upon knowledge of the exogenous variables values, a VAR user not only can overlook any economic theory underlying the model, but also does not need to know the values of the exogenous variables. There are no exogenous variables in a VAR model.

Sims' methodology consists of regressing each current (non-lagged) variable on all variables in the model lagged p times. Consequently, an unrestricted VAR model is over-parameterised in the sense that some of the lagged variables entering the model could be properly excluded on the basis of statistical insignificance. However, the advocates of this methodology, including Sims himself, advise against this procedure, arguing that the imposition of zero restrictions may suppress important information and that the regressors in a VAR specification are likely to be highly collinear, so that the t-tests on individual coefficients are not reliable guides for down-sizing the model.

In practice, it is impossible to avoid the consideration of prior restrictions on a VAR system. Sample size constraints mean that there will be a limit to the number of variables included and the number of lags imposed. Even if the sample size is not a problem, the possibility of giving some structure to a VAR model and using it for economic analysis alongside forecasting purposes, requires the imposition of restrictions. Well-founded theoretical constraints may help to transform an unrestricted VAR specification into a restricted VAR model "*consistent with even highly detailed economic theories*" (Charemza and Deadman, 1997, p.157). The consideration of such restrictions allows for identification of the model and economic interpretation of the structural parameters in a way not possible with the reduced-form. Furthermore, *a priori* information concerning the parameters allows for testing and including restrictions that can improve the precision of estimates and reduce the forecast error variance. Hence, even if the main interest is forecasting, the down-sizing of an over-parameterised VAR can help to improve the results.

The specification of a reduced-form VAR can be illustrated using a simple example with three time series variables. Consider the variables $\{x_t\}$, $\{y_t\}$ and $\{w_t\}$ and assume that the time path of x_t (y_t , w_t) is affected by current and past realizations of the other two variables. Establishing a maximum lag-length of one, we can specify the following system:

$$\begin{aligned} x_t &= b_{10} - b_{12}y_t - b_{13}w_t + \gamma_{11}x_{t-1} + \gamma_{12}y_{t-1} + \gamma_{13}w_{t-1} + \varepsilon x_t \\ y_t &= b_{20} - b_{22}x_t - b_{23}w_t + \gamma_{21}x_{t-1} + \gamma_{22}y_{t-1} + \gamma_{23}w_{t-1} + \varepsilon y_t \\ w_t &= b_{30} - b_{32}x_t - b_{33}y_t + \gamma_{31}x_{t-1} + \gamma_{32}y_{t-1} + \gamma_{33}w_{t-1} + \varepsilon w_t \end{aligned} \tag{7.7}$$

where εx_t , εy_t and εw_t are assumed to be white-noise disturbances

System (7.7) represent a structural first-order vector autoregressive model VAR(1). The equations in (7.7) are not reduced-form equations since each endogenous variable has a contemporaneous effect on the other two. However, it is possible to transform the system into a more tractable form. Using matrix algebra we can write (7.7) as:

$$\begin{bmatrix} 1 & b_{12} & b_{13} \\ b_{21} & 1 & b_{23} \\ b_{31} & b_{33} & 1 \end{bmatrix} \begin{bmatrix} x_t \\ y_t \\ w_t \end{bmatrix} = \begin{bmatrix} b_{10} \\ b_{20} \\ b_{30} \end{bmatrix} + \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} \\ \gamma_{31} & \gamma_{32} & \gamma_{33} \end{bmatrix} \begin{bmatrix} x_{t-1} \\ y_{t-1} \\ w_{t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon x_t \\ \varepsilon y_t \\ \varepsilon w_t \end{bmatrix}$$

or $\quad Bz_t = \Gamma_0 + \Gamma_1 z_{t-1} + e_t \qquad (7.8)$

where $B = \begin{bmatrix} 1 & b_{12} & b_{13} \\ b_{21} & 1 & b_{23} \\ b_{31} & b_{33} & 1 \end{bmatrix}$; $z_t = \begin{bmatrix} x_t \\ y_t \\ w_t \end{bmatrix}$; $\Gamma_0 = \begin{bmatrix} b_{10} \\ b_{20} \\ b_{30} \end{bmatrix}$; $\Gamma_1 = \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} \\ \gamma_{31} & \gamma_{32} & \gamma_{33} \end{bmatrix}$; $e_t = \begin{bmatrix} \varepsilon x_t \\ \varepsilon y_t \\ \varepsilon w_t \end{bmatrix}$

Pre-multiplying (7.8) by B^{-1} allows us to obtain the reduced-form VAR model:

$$z_t = A_0 + A_1 z_{t-1} + \epsilon_t \tag{7.9}$$

where $A_0 = B^{-1}\Gamma_0$; $A_1 = B^{-1}\Gamma_1$; $\epsilon_t = B^{-1}e_t$

If we define a_{i0} as the element i of vector A_0, a_{ij} as the element of row i and column j in matrix A_1 and ei_t as the element i of vector ϵ_t we can write (7.9) in the equivalent form:

$$\begin{aligned} x_t &= a_{10} + a_{11}x_{t-1} + a_{12}y_{t-1} + a_{13}w_{t-1} + e1_t \\ y_t &= a_{20} + a_{21}x_{t-1} + a_{22}y_{t-1} + a_{23}w_{t-1} + e2_t \\ w_t &= a_{30} + a_{31}x_{t-1} + a_{32}y_{t-1} + a_{33}w_{t-1} + e3_t \end{aligned} \tag{7.10}$$

System (7.7) is called the structural VAR model and system (7.10) is called the reduced-form VAR model. Note that the right-hand side of all equations in (7.10) includes only predetermined variables. As the error terms are assumed not to be serially correlated, each equation in the system can be efficiently estimated using OLS.

7.3. SPECIFICATION OF A VAR MODEL FOR THE UK TOURISM DEMAND IN FRANCE, SPAIN AND PORTUGAL

The arguments presented in the previous section lead to the conclusion that when empirical analysis about the existence of long-run relationships among more than two non-stationary time series must be conducted, and there are doubts about the exogeneity assumption with respect to some of the variables and/or whether zero restrictions should be imposed, the appropriate econometric approach consists of treating all variables as endogenous within a reduced-form VAR framework. Then, using one or more of the methods available, tests for the exogeneity of the set of variables in doubt should be undertaken. Once the endogenous-exogenous division is established, the Johansen (1988) procedure can be used to test for the existence of cointegrating relationship(s). The number of cointegrated vectors found establishes the number of meaningful long-run relationships existing among the variables. Imposing restrictions to exactly-identify the underlying structural VAR can assess the estimates of the long-run coefficients. Once the structural VAR is identified, over-identifying restrictions making the VAR compatible with specific economic theories, can also be tested.

In this section, we specify a basic general VAR model for the UK tourism demand in France, Spain and Portugal, which is used for both forecasting and economic analysis purposes. The analysis within the basic model includes the variables in vector $z_t' = \{WF_t, WS_t, WP_t, PP_t, PS_t, PF_t, E_t\}$. All variables in vector z_t, are the same as those included in the AIDS systems, already characterised in chapter 5.

The specification of a VAR model starts with establishing the order of integration and the appropriate lag-length of its variables. Hence, in this section, we use unit root tests to determine whether the time series in vector z_t are stationary, and diagnostic tests to establish the lag-length of the VAR which eliminates serial correlation from the error terms.

7.3.1. ORDER OF INTEGRATION OF THE VARIABLES INCLUDED IN THE VAR

The issue of whether the variables in a VAR need to be stationary has been a matter for debate among researchers. As nonstationary variables can lead to spurious regressions and invalid statistical inference, the usual assumption (see, for example, Griffiths *et al.*, 1993, p.693 and Judge *et al.*,1988, p.754) is that all variables in a VAR should be stationary. Thus, if time series entering a VAR contain unit roots, the data must be appropriately transformed (by means of differentiation, for example).

However, differentiating nonstationary variables in order to achieve stationarity excludes important long-run information concerning the co-movements of the variables. In fact, researchers dealing with the VAR approach, for example Sims (1980) and Johansen (1988), recommend against differentiation even if, as stressed by Sims, the variables contain unit roots. In this author's view, the main objective of a VAR is to determine the existence of long-run relationships among the variables. Hence, differentiation does not add much to this main purpose of the model and, instead, "throws away" important information about the co-movements of the data which can help to identify possible cointegrating vectors. If cointegrating vector(s) are present in a VAR, meaningful long-run relationship(s) exist among the variables and an error correction mechanism is active within the short-run dynamics. This information is crucial for economic analysis and policy purposes and can be retrieved from the reduced-form VAR. Furthermore, the Johansen (1988) procedure for determining the number of cointegrating vectors in a VAR is considered to be more reliable when all the level variables involved are I(1).

Although the usual approach adopted by VAR "aficionados" is to work with variables in levels, even if these are non-stationary *"it is important to recognise the effect of unit roots on the distribution of estimators"* (Harvey, 1990, p.83). Thus, one should exercise special caution regarding the specification of unrestricted VAR models which include non-stationary time series because, unless the variables are cointegrated, the estimation of a VAR model with non-stationary variables may lead to spurious results and invalid inference.

Table 7.1 shows the statistic values and respective critical values at the 5% significance level, of the DF and ADF unit root tests for the levels and first differences of the variables WF, WS, WP, PP, PS, PF and E. It also shows the AIC and SBC criteria for the lag-length selection of the Dickey-Fuller test equations. The values in bold indicate the valid test statistic for comparison with the critical value, according to both selection criteria.

The tests clearly indicate that all variables in levels are non-stationary and, except for ΔPP, all variables in first differences are stationary. This means that according to the DF and ADF tests, all variables, except PP, can be considered to be I(1) variables.

Table 7.1: Unit root DF and ADF tests for the variables WF, WS, WP, PF, PS, PP and E

Variable	Test	Statistic	AIC criterion	SBC criterion	Critical value
WF	ADF(0)	**-2.100**	54.21*	52.87*	-2.971
	ADF(1)	-2.008	50.81	48.87	-2.975
ΔWF	ADF(0)	**-5.163**	49.71*	48.42*	-2.975
	ADF(1)	-3.284	46.36	44.47	-2.980
WS	ADF(0)	**-1.965**	52.09*	50.76*	-2.971
	ADF(1)	-1.756	48.71	46.77	-2.975
ΔWS	ADF(0)	**-5.187**	48.08*	46.72*	-2.975
	ADF(1)	-3.510	44.78	42.89	-2.980
WP	ADF(0)	**-1.738**	86.03*	84.70*	-2.971
	ADF(1)	-1.458	81.64	79.69	-2.975
ΔWP	ADF(0)	**-5.429**	81.49*	80.20*	-2.975
	ADF(1)	-4.827	78.25	76.36	-2.980
PF	ADF(0)	**-1.869**	32.98*	31.64*	-2.971
	ADF(1)	-2.311	32.24	30.30	-2.975
ΔPF	ADF(0)	**-3.435**	30.59*	29.23*	-2.975
	ADF(1)	-3.417	28.69	26.80	-2.980
PS	ADF(0)	**-2.083**	33.50*	32.17*	-2.971
	ADF(1)	-2.428	31.74	29.80	-2.975
ΔPS	ADF(0)	**-3.788**	29.78*	28.48*	-2.975
	ADF(1)	-2.968	27.24	25.35	-2.980
PP	ADF(0)	-1.418	31.74	30.41	-2.971
	ADF(1)	**-2.605**	34.17*	32.23*	-2.975
ΔPP	ADF(0)	**-2.480**	31.81*	30.51*	-2.975
	ADF(1)	-2.245	29.16	27.27	-2.980
E	ADF(0)	**-2.362**	20.51*	19.18*	-2.971
	ADF(1)	-1.624	18.48	16.53	-2.975
ΔE	ADF(0)	**-3.696**	18.07*	16.77*	-2.975
	ADF(1)	-2.675	16.94	15.06	-2.980

The case of the variable for the price of Portugal (PP) is a special one. As previously mentioned in chapter 4, the DF and ADF tests may not be adequate for testing stationarity in a time series that may present changing means over the sample period.[5] Owing to the political upheaval of the 1970s in Portugal, the possibility of a change in the mean of the variable PP is very likely. Hence, an adequate unit root test allowing for a shift in the mean of this time series requires a different approach. The Phillips-Peron (1988) test seems to be an appropriate way for testing stationarity in the case of variable PP. The Phillips-Peron test is based on a simple DF regression, such that

$$\Delta PP_t = \beta_0 + \beta_1 PP_{t-1} \tag{7.11}$$

The OLS estimation results for equation (7.11) are

$$\Delta \hat{PP}_t = \underset{(0.536)}{0.0079} - \underset{(-1.4177)}{0.14139} PP_{t-1}$$

The t ratio of β_1 coefficient is simply the ADF(0) statistic for variable PP in Table 7.1. The non-parametric correction to this statistic proposed by Phillips and Peron is carried out by deriving the White and Newey-West adjusted covariance matrix and computing the adjusted variances. The estimation results for regression (7.11) using the adjusted covariance matrix are as follow:

$$\Delta \hat{PP}_t = \underset{(0.5112)}{0.0079} - \underset{(-2.9039)}{0.14139} PP_{t-1}$$

The t ratio for the β_1 coefficient is now the adjusted t ratio and also the valid statistic to compare with the critical value. As $|-2.9039| < |-2.971|$ we cannot reject the hypothesis of PP being non-stationary. This confirms the result obtained with the ADF test displayed in Table 7.1 for the same variable. However, the problem resides in the fact that the ADF test indicates the first difference of PP (ΔPP) to be non-stationary as well. If this is the case, then PP cannot be considered an I(1) variable. Therefore, we perform the Phillips-Peron test for ΔPP, running the following regression:

$$\Delta\Delta PP_t = \beta'_0 + \beta'_1 \Delta PP_{t-1} \tag{7.12}$$

The OLS estimation results for regression (7.12) are

$$\Delta\Delta \hat{PP}_t = \underset{(-0.1211)}{-0.001692} - \underset{(-2.4803)}{0.51046} \Delta PP_{t-1}$$

The estimation results for (7.12) using the adjusted covariance matrix are

[5] In chapter 4 we showed that the PP variable is accepted as an I(1) variable with MacKinnon critical values at the 10% significance level and with Charemza and Deadman critical values at the 5% level.

$$\Delta\Delta\hat{P}P_t = -\underset{(-0.1576)}{0.001692} - \underset{(-3.9587)}{0.51046}\Delta PP_{t-1}$$

The t ratio for β_1 is the adjusted t ratio which is the valid statistic for comparison with the critical value. As $|-3.9587| > |-2.975|$ we cannot reject the hypothesis of ΔPP being stationary. Hence, according to the Phillips-Peron test, the variable in levels PP is I(1). Therefore, we conclude that there is sufficient statistical evidence to consider all variables in levels included in the VAR as integrated of order one or I(1) variables.

7.3.2 Determination of the order of the VAR

The maximum lag-length (p) imposed on the structure of a VAR model is generally termed the order of the VAR and denoted by VAR(p). The order of a VAR plays a crucial role in empirical analysis since its dimension can determine whether the estimation method applied is consistent and, thus, whether the estimation result's and the statistical inference are valid. Therefore, in the process of choosing p, special care should be taken to ensure that it is high enough for the disturbances not to be serially correlated, and small enough for the VAR not to be over-parameterised.

The observations available for this VAR modelling exercise do not allow for an extensive lag-structure. In fact, given the number of variables involved and the existing sample size, the lag-length cannot exceed two. Taking these limitations into consideration, we used the AIC and SBC selection criteria and the adjusted (for small samples) Likelihood Ratio (LR) test for selecting the order of the VAR with the maximum lag-length permitted. Table 7.2. presents the LR test statistic and the two selection criteria for choosing p.

Table 7.2: AIC and SBC criteria and adjusted LR test for selecting the order of the VAR

Order (p)	AIC	SBC	Adjusted LR test
2	296.91	246.37	-------
1	296.59	269.37	$\chi^2(36) = 37.67(0.393)$
0	185.90	182.01	$\chi^2(72) = 366.02(0.000)$

The LR test rejects order zero but cannot reject a first order VAR. The SBC criterion clearly indicates the order of the VAR to be one. The AIC criterion indicates p to be two, but by a very small margin. It is common for the SBC criterion to select a lower order VAR as compared with the AIC criterion so, in this case, we can be quite confident about the order to be selected. Yet, it is always prudent to examine the residuals of the individual equations in order to check for statistical evidence of no serial autocorrelation.

7.3.3. THE UNRESTRICTED VAR SPECIFICATION OF THE UK DEMAND FOR TOURISM

The reduced form of the first order unrestricted VAR for the UK tourism demand in France, Spain and Portugal (denoted by VAR I) can be written as:[6]

$$\begin{aligned}
WF_t &= a_{10} + a_{11}WF_{t-1} + a_{12}WS_{t-1} + a_{13}PP_{t-1} + a_{14}PS_{t-1} + a_{15}PF_{t-1} + a_{16}E_{t-1} + e1_t \\
WS_t &= a_{20} + a_{21}WF_{t-1} + a_{22}WS_{t-1} + a_{23}PP_{t-1} + a_{24}PS_{t-1} + a_{25}PF_{t-1} + a_{26}E_{t-1} + e2_t \\
PP_t &= a_{30} + a_{31}WF_{t-1} + a_{32}WS_{t-1} + a_{33}PP_{t-1} + a_{34}PS_{t-1} + a_{35}PF_{t-1} + a_{36}E_{t-1} + e3_t \qquad (7.12) \\
PS_t &= a_{40} + a_{41}WF_{t-1} + a_{42}WS_{t-1} + a_{43}PP_{t-1} + a_{44}PS_{t-1} + a_{45}PF_{t-1} + a_{46}E_{t-1} + e4_t \\
PF_t &= a_{50} + a_{51}WF_{t-1} + a_{52}WS_{t-1} + a_{53}PP_{t-1} + a_{54}PS_{t-1} + a_{55}PF_{t-1} + a_{56}E_{t-1} + e5_t \\
E_t &= a_{60} + a_{61}WF_{t-1} + a_{62}WS_{t-1} + a_{63}PP_{t-1} + a_{64}PS_{t-1} + a_{65}PF_{t-1} + a_{66}E_{t-1} + e6_t
\end{aligned}$$

or in the equivalent form:

$$z^*_t = A_0 + A_1 z^*_{t-1} + \epsilon_t \qquad (7.13)$$

where $z^*_t = [WF_t, WS_t, PP_t, PS_t, PF_t, E_t]$

Since all observations of the share variables (WF, WS and WP) sum to unity, exact multicollinearity exists between these variables' equations if all three are included in the VAR. Hence, one of the share equations must be omitted. We omit the equation for the expenditure share of Portugal (WP_t). However, as noted in chapter 5, the estimation results are invariant whichever share equation is omitted and, by the adding-up property, all coefficient estimates for the omitted equation can be retrieved from the estimates of the coefficient of the first two equations.

The statistical quality of VAR I model can be accessed by estimating (7.13) and computing the relevant diagnostic statistics. Table 7.3 shows the estimation results (t ratios in brackets), the AIC and SBC criteria and a set of test statistics - adjusted R^2, F statistic and χ^2 statistic for diagnostic tests of serial correlation, functional form, error normality and heteroscedasticity (p values in brackets) - for all equations included in the basic unrestricted VAR I structure.

[6] This basic specification of the unrestricted VAR (denoted by VAR I) is later subjected to modifications due to the introduction of dummy variables and the exogeneity assumption concerning some of the regressors.

Table 7.3. Estimation results and statistical performance of the basic VAR I model

REGRESSORS	EQUATIONS					
	WF_t	WS_t	PP_t	PS_t	PF_t	E_t
WF_{t-1}	0.9518 (2.10)	-0.5617 (-1.14)	-2.0208 (-1.57)	-1.0503 (-0.75)	-1.3501 (-1.32)	-0.3106 (-0.15)
WS_{t-1}	0.6820 (1.44)	-0.2932 (-0.57)	-1.9954 (-1.48)	-0.7034 (-0.48)	-0.9772 (-0.78)	-1.3497 (-0.61)
PP_{t-1}	-0.0712 (-1.14)	0.1062 (1.56)	1.2525 (7.05)	0.2108 (1.10)	0.5033 (3.06)	-0.4728 (-1.62)
PS_{t-1}	0.4620 (4.85)	-0.4622 (-4.45)	-0.1341 (-0.49)	0.7453 (2.54)	-0.0014 (-0.01)	0.0602 (0.14)
PF_{t-1}	-0.3195 (-2.38)	0.3157 (2.16)	-0.4922 (-1.29)	-0.2276 (-0.55)	0.0598 (0.17)	0.6837 (1.09)
E_{t-1}	-0.0071 (-0.79)	0.0000 (0.00)	-0.0312 (-1.23)	-0.0027 (-0.10)	-0.0318 (-1.35)	0.9315 (22.40)
Intercept	-0.2630 (-0.60)	0.8449 (1.77)	1.9533 (1.57)	0.7627 (0.57)	1.1671 (1.01)	1.1990 (0.59)
SLECTION CRITERIA AND DIAGNOSTIC STATISTICS						
AIC	61.24	58.82	31.99	29.77	34.10	18.12
SBC	56.58	54.15	27.33	25.11	29.44	13.45
Adjusted R^2	0.789	0.824	0.771	0.562	0.612	0.991
F statistic	17.88	22.12	16.11	6.79	8.09	511.34
Serial Correlation	0.78(0.38)	2.48(0.12)	0.73(0.39)	0.19(0.67)	0.37(0.54)	0.58(0.45)
Functional Form	3.81(0.05)	8.66(0.00)	1.85(0.17)	7.39(0.01)	0.00(0.98)	0.18(0.67)
Normality	1.02(0.60)	1.56(0.46)	0.56(0.75)	4.82(0.09)	11.65(0.00)	1.66(0.44)
Heteroscedasticity	0.39(0.53)	1.19(0.28)	1.33(0.25)	0.28(0.60)	0.03(0.87)	0.21(0.65)

There seems to be no statistical evidence of serial correlation in the error terms of VAR I. Hence, the lag-length selected seems to be adequate. However, the tests indicate problems in the functional form and in the normality of the error terms of some equations. We are particularly interested in the expenditure share equations and for these equations the problems of functional form are severe. These problems may be related to the assumption of endogeneity for all variables, and to the omission of relevant dummy variables.

In the AIDS system of equations specified in chapter 5, the UK tourism demand for France, Spain and Portugal, represented by the tourism shares of France (WF), Spain (WS) and Portugal (WP), are assumed to be the only endogenous dependent variables. In the AIDS model, changes in these variables are explained by a set of exogenous variables which include tourism prices (PF, PS and PP) and the UK real per capita tourism expenditure (E). In chapter 3, we explained why, in a tourism demand context, prices may be considered exogenous and, in chapter 5, we explained how the multi-stage budgeting process underlying the AIDS approach sets the variable 'real per capita expenditure' (E) as a determinant of the demand shares.

Within a VAR model framework, we are willing to question the pre-assumed exogeneity of the price variables. However, there seems to be no obvious theoretical basis for considering the UK real per capita expenditure on tourism, E, an endogenous variable within the system of equations defined in (7.13). On the contrary, there appears to be a number of theoretical and empirical reasons to considering it as (weakly) exogenous. If all variables in a VAR model are endogenous, there is a bi-directional cause-effect relationship (feedback) between them. However, even if it is reasonable to consider that changes in the UK real per capita expenditure affect the tourism shares of important UK holiday destinations such as France, Spain or Portugal, it does not seem to make much sense to consider that variations in these shares affect the way in which UK consumers allocate their budget. The multi-stage budget allocation underlying the rationality of an expenditure share system does not comprise such bi-directional cause-effect features. In fact, UK consumers (tourists) may allocate their income to expenditure on tourism and other commodities prior to (or in spite of) any knowledge of each item's current or past share values in their list of acquisitions.

Empirical evidence also seems to support this line of reasoning. It can be inferred from the estimation results of the equation for E_t depicted in table 7.3, that the 99% of this variable's variations explained by the model lie exclusively on its own lagged value. No other variable in that equation is individually or jointly statistically significant. In fact, the F test for the joint significance of all explanatory variables, excluding E_{t-1}, presents a value of 0.86 implying that the hypothesis of these variables' coefficients being zero cannot be rejected. Furthermore, the estimation results of the VAR equations for WF_t and WS_t indicate that the lagged value of E_t does not affect significantly the current tourism shares of France or Spain.

In a model, the existence of a causal link between the destinations expenditure shares (Wi_t) and the UK real per capita expenditure (E_t) can be investigated by analysing the relationships between the error term of the model and the stochastic disturbance in the assumed data generating process (d.g.p.) of the E_t variable. Following Harris (1995, p. 4) we consider the i^{th} share variable equation to be

$$Wi_t = \alpha_0 + \alpha_1 E_t + \alpha_2 Wi_{t-1} + ui_t \qquad \text{where } i = F, S, P \qquad (7.14)$$

Assuming that E_t is a stochastic variable, and letting the underlying d.g.p. be given by

$$E_t = \beta_1 E_{t-1} + \varepsilon_t \;\; ; \;\; \beta_1 < 1 \text{ and } \varepsilon_t \rightarrow N(0, \sigma^2) \qquad (7.15)$$

If ui_t and ε_t are not correlated we can state that $EV(ui_t, \varepsilon_s) = 0$ for all t, s (where EV stands for expected value, not to be confused with the time series variable E_t). Then, it is possible to treat E_t as if it were fixed, that is, E_t is independent of ui_t such that $EV(E_t, ui_t) = 0$. Hence, we can treat E_t as exogenous in terms of (17.14), with the *current* value of the UK real expenditure per capita (E_t) being said to Granger-cause Wi_t. Equation (7.14) is a conditional model since Wi_t is conditional on E_t, with E_t being determined by the marginal model (7.15). Note that if (7.15) is reformulated as $E_t = \beta_1 E_{t-1} + \beta_2 Wi_{t-1} + \varepsilon_t$, $EV(E_t, ui_t)=0$ is still valid. However, since past values of Wi_t now determine E_t, this variable can only be considered weakly exogenous in the conditional model (7.14). The *current* value E_t still causes Wi_t but not in the Granger sense, since past values of Wi_t now determine E_t.

As Hendry (1995, p.164) states, "*a variable cannot be exogenous per se*". A variable can only be exogenous with respect to a set of parameters of interest. Therefore, if E_t is deemed to be exogenous with respect to the parameters α_j (j = 0, 1, 2) in (7.14), the marginal model can be neglected and the conditional model (7.14) is complete and sufficient to sustain valid inference. Consequently, knowledge of the marginal model will not significantly improve the statistical and forecasting performance of the conditional model. Only in this sense can a set of variables be considered exogenous.

As an indication of the possibility that E_t can be treated as an exogenous variable with respect to the parameters of the expenditure share equations in VAR I model, we run regression (7.14) for the expenditure shares of France, Spain and Portugal and regression (7.15) as a representation of the d.g.p. of E_t. We retrieve the residual series of these four regressions, namely u_{Ft}, u_{St}, u_{Pt} standing for the residual series of (7.14) for, respectively, France, Spain and Portugal, and ε_t standing for the residuals of (7.15). We then run individual regressions of the current and lagged values (up to the fifth lag) of the residuals

u_{it} (i = F, S, P) on the current and lagged (up to the fifth lag) values of ε_t. The estimation results of all 90 regressions indicate no significant linear relationship linking the current or lagged residuals of the conditional models to the current and lagged residuals of the marginal model. We accept this statistical evidence as an indication that knowledge of the marginal model does not improve the statistical or forecasting performance of the conditional (on E_{t-1}) equations for WF_t, WS_t, PP_t, PS_t and PF_t in the VAR and so treat E_t as a (weakly) exogenous variable in the VAR structure given by (7.13).

The claim that in the relationships between the variable 'UK real per capita expenditure' (E_t) and the destination expenditure shares (Wi_t, i=F, S) feedback effects might be absent in this particular VAR specification, can be further investigated using the causality test proposed by Granger (1969). The Granger-causality test is a well known and widely used procedure to establish the existence of significant feedback effects between variables in VAR specifications. A test for causality is related to whether the lags of a variable are statistically significant in the equation of another variable. In other words, and using the definition given in Griffiths *et al.* (1992, p. 695), "*a variable $y1_t$ is said to be Granger-caused by $y2_t$ if current and past information on $y2_t$ help improve forecasts in $y1_t$.*" Therefore, if the lagged values of $y2_t$ do not help to improve the forecasts of $y1_t$, that is, if the lagged values of $y2_t$ are statistically insignificant in the reduced form equation for $y1_t$, then $y2_t$ does not Granger-cause $y1_t$.[7]

The multivariate generalisation of the Granger-causality concept is known as "block Granger-causality" and it can be used for establishing if one or more variables in a VAR should or should not integrate the group of endogenous variables. The testing of the null hypothesis that the coefficients of a subset of variables in a VAR are zero using the log-likelihood ratio (LR) statistic, is known as block Granger *non*-causality test. This test provides a statistical measure of the extent to which lagged values of a set of variables (say E_t), are important in predicting another set of variables (say Wi_t), once lagged values of the latter (say Wi_{t-1}) are included in the model. The LR test for block Granger non-causality of the variable E_t which tests the null hypothesis that the coefficients of E_{t-1} are null in the block equations for WF_t, WS_t, PP_t, PS_t and PF_t presents the statistic value of $\chi^2(5) = 10.578$. As the critical values at the 5% level is 11.071, the null hypothesis cannot be rejected.

[7] Note that Granger's concept of causality does not imply a cause-effect relationship but rather is based on the concept of "predictability". As Charemza and Deadman (1997, p. 165) states, "*in econometrics, causality has a meaning more on the lines of 'to predict' rather than 'to produce'.*"

Accordingly, we now consider E_t not to integrate the set of endogenous variables in the VAR model. Rather, it integrates the set of pre-determined and/or exogenous variables. Hence, we reformulate the basic model to include this assumption and estimate a new version of the VAR (denoted VAR II). The AIC and SBC selection criteria and the same set of test statistics for accessing the statistical quality of the previous model are now used to access the quality of this second version of the basic VAR and presented in Table 7.4.

Table 7.4. AIC and SBC selection criteria and diagnostic tests for the VAR II model

Selection criteria and diagnostic testes	EQUATIONS				
	WF_t	WS_t	PP_t	PS_t	PF_t
AIC	61.39	58.83	33.58	30.02	34.62
SBC	56.73	54.17	28.92	25.36	29.96
Adjusted R^2	0.792	0.825	0.795	0.570	0.656
F statistic	18.11	22.15	18.47	6.97	8.53
Serial Correlation	0.83(0.38)	2.16(0.14)	0.27(0.50)	0.46(0.67)	0.47(0.49)
Functional Form	3.28(0.07)	8.49(0.00)	0.70(0.40)	5.56(0.02)	0.02(0.90)
Normality	0.86(0.65)	1.55(0.46)	0.05(0.97)	3.04(0.22)	6.49(0.04)
Heteroscedasticity	0.42(0.52)	1.23(0.27)	0.84(0.36)	0.21(0.65)	0.00(0.95)

The selection criteria and the diagnostic tests indicate this new specification of the VAR to be, overall, statistically more adequate than the previous one. However, problems with the functional form of the equations for the share variables still remain.

As explained in previous chapters, there is reason to believe that political and economic events in the 1970s (change of political regimes in Portugal and Spain and the oil crises) may have affected the time path of the variables included in the VAR. In addition, we detected, in chapter 5, a structural break in the variable 'real per capita expenditure' (E_t) (linked to Spain and Portugal's integration process in the EU) and modelled it with a slope dummy separating this variable's coefficient into two. The 1970s events are taken into account by adding a dummy variable D1, which takes the value of unity in the period 1974-1981 and zero otherwise. Since the modelling of slope coefficients has not been fully addressed either theoretically or empirically within the VAR approach, to mimic the structural break detected previously, we use dummy variables D2 (taking the value of unity in the period 1982-1997 and zero otherwise) and D3 (taking the value of unity in the period 1989-1997 and zero otherwise) to separate the integration process of Spain and Portugal in the EU into two periods: the integration period (1982-1988) and the post-integration period (1989-1997). These dummies are assumed to be exogenous, leading to the modelling of a partial VAR system with exogenous variables. This third version of the basic VAR is denoted VAR III. Table 7.5 shows the AIC and SBC criteria and a set of diagnostic statistics to access the statistical quality of the VAR III model.

Table 7.5. AIC and SBC selection criteria and diagnostic tests for the VAR III model

Selection criteria and diagnostic testes	EQUATIONS				
	WF_t	WS_t	PP_t	PS_t	PF_t
AIC	69.79	65.69	32.27	27.28	33.69
SBC	63.13	59.03	25.61	20.62	27.02
Adjusted R^2	0.892	0.899	0.788	0.508	0.623
F statistic	25.87	27.64	12.16	4.10	5.96
Serial Correlation	0.35(0.55)	0.02(0.90)	0.04(0.85)	0.21(0.64)	0.10(0.66)
Functional Form	0.04(0.84)	0.21(0.64)	0.57(0.45)	7.24(0.01)	0.14(0.71)
Normality	1.28(0.53)	0.63(0.73)	0.08(0.96)	3.50(0.17)	11.64(0.03)
Heteroscedasticity	0.10(0.75)	0.01(0.92)	0.68(0.41)	0.26(0.61)	0.03(0.86)

The results of Table 7.5 indicate that the statistical quality of VAR III over-performs those of VAR I and VAR II, particularly, with respect to the expenditure share equations which are the relationships we are interested in. Moreover, the LR test for block Granger non-causality of variable E_t performed on the VAR III (which includes the three dummy variables), confirms the results of the similar test performed on VAR I (not including the dummy variables). In fact, the LR test presents now the statistic value of $\chi^2(5) = 8.376$, which is well below the 5% critical value (11.071), further supporting the null hypothesis that the coefficients of E_{t-1} are statistically null in the block equations of the VAR.

For a clear view of this and other statistical tests performed, we present Table 7.6 which shows the results obtained from testing hypothesis involving the variables included in the VAR models. The tests concern the significance of the intercept, the joint significance of dummy variables D1, D2 and D3 and the Granger block non-causality of variable E_{t-1}. The first column of Table 7.6 presents the various null hypothesis for each test and shows the variables entering the VAR model under the "unrestricted" hypothesis (U), and under the "restricted" null hypothesis (R) (for simplicity, the time subscripts are omitted). In each case, the set of endogenous variables entering the model is separated from the set of deterministic components and exogenous variables by the symbol '&'. The second column presents the maximum value of the likelihood function (ML) for the unrestricted and restricted alternatives. In the third column the log-likelihood ratio (LR) statistic is computed. Under the null, the LR statistic is asymptotically distributed as χ^2 with degrees of freedom (i) equal to the number of restrictions. The null hypothesis is rejected if the LR statistic is larger than the relevant critical value of the χ^2 distribution.

Table 7.6. LR tests for restrictive hypothesis on the intercept and variables of the VAR

Model	ML	LR → χ^2(i)	Critical value (5%)	Result
H_0: Non-significance of intercept (INT)				
U: WF WS PP PS PF E & INT	351.68	$\chi^2(6)$=18.05	12.59	Rejected
R: WF WS PP PS PF E &	342.05			
H_0: Non-significance of dummy variables D1 D2 D3				
U: WF WS PP PS PF E & D1 D2 D3 INT	385.40	$\chi^2(18)$=67.43	28.87	Rejected
R: WF WS PP PS PF E & INT	351.68			
H_0: Block non-causality of E with no dummy variables				
U: WF WS PP PS PF E & INT	351.68	$\chi^2(5)$=10.58	11.07	Not rejected
R: WF WS PP PS PF & E INT	346.39			
H_0: Block non-causality of E with dummy variables				
U: WF WS PP PS PF E & D1 D2 D3 INT	385.40	$\chi^2(5)$=8.37	11.07	Not rejected
R: WF WS PP PS PF & E D1 D2 D3 INT	381.21			

Although VAR III is statistically more robust than VAR I, the latter is a more general model, not being restricted in any way, while the former is a specific model, being a partial system conditioned on exogenous variables. Therefore, we believe it is interesting to use both models for comparison purposes, in accessing the main features of the UK tourism demand for France, Spain and Portugal, within the reduced-form VAR specification.

A reduced-form VAR specification is not the ideal means of conducting a reliable economic analysis of the structural long-run relationships existing among the variables. The reasons are first, an *a-theoretical* model is unlike to produce estimation results interpretable within the limits of economic theory; second, the economic interpretation of the structural parameters is only possible if the underlying structural model can be identified from the reduced-form VAR; finally, the lag-structure included in the reduced-form VAR is likely to lead to over-parameterisation causing imprecision in the coefficients' estimates and obscuring the economic meaning of the long-run parameters. Yet, if the main purpose of an econometric modelling exercise is forecasting, the reduced-form VAR specification has advantages. For instance, the estimation of the structural model is irrelevant in the forecasting process since, although under-identified, a VAR model can provide unbiased and efficient forecasts. These reasons justify the use of the reduced-form VAR I and VAR III specifications for forecasting rather than economic analysis purposes.

7.3.4. The forecasting ability of the reduced-form VAR I and VAR III models

Except for the choice of the variables, VAR I model does not bear any other *a priori* assumption or restriction. Thus, we can qualify this model as a "pure" VAR. However, VAR III model incorporates theoretical and empirical assumptions such as conditioning the model on E_t and three dummy variables, which can be considered constraints to the more general VAR. It would be interesting to obtain forecasts of the tourism shares for France, Spain and Portugal from these two models, and compare their forecasting ability. For this purpose, the models are now estimated for the in-sample period 1969-1993, including the last four observations in the forecasting period. Tables 7.7, 7.8. and 7.9 report the actual and forecast values of, respectively, the expenditure shares of France, Spain and Portugal obtained with VAR I and VAR III models. It also reports the forecast errors, and a number of summary statistics (mean of absolute errors - MAE; mean of squared errors – MSE; root mean of squared errors - RMSE) for evaluating the models' forecasting accuracy. These forecast evaluation criteria are defined in chapter 8, section 8.3.2.

Table 7.7: Forecasting results for the UK expenditure share of France

		1994	1995	1996	1997
Actual values		**0.39700**	**0.38540**	**0.38967**	**0.40481**
VAR I	Forecast	0.34422	0.36840	0.39676	0.42782
	Forecast error	0.052780	0.017006	-0.007086	-0.023007
VAR III	Forecast	0.34274	0.44818	0.41524	0.40121
	Forecast error	0.054261	-0.062778	-0.025567	0.003595

SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS			
		Estimation period: (1970 -1993)	Forecast period: (1994 –1997)
VAR I	MAE	0.015321	0.024970
	MSE	0.000315	0.000914
	RMSE	0.017734	0.030226
VAR III	MAE	0.008510	0.036550
	MSE	0.000119	0.001888
	RMSE	0.010917	0.043451

Table 7.8: Forecasting results for the UK expenditure share of Spain

		1994	1995	1996	1997
Actual values		**0.51857**	**0.52625**	**0.52292**	**0.50691**
VAR I	Forecast	0.57000	0.54972	0.52281	0.49058
	Forecast error	-0.051431	-0.023469	0.000108	0.016328
VAR III	Forecast	0.56593	0.45798	0.48770	0.50172
	Forecast error	-0.047361	0.068264	0.035218	0.005191

SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS			
		Estimation period: (1970 -1993)	Forecast period: (1994 –1997)
VAR I	MAE	0.016475	0.022834
	MSE	0.000301	0.000866
	RMSE	0.019780	0.029422
VAR III	MAE	0.010340	0.039009
	MSE	0.000192	0.002043
	RMSE	0.013847	0.045195

Table 7.9: Forecasting results for the UK expenditure share of Portugal

		1994	1995	1996	1997
Actual values		**0.08443**	**0.08835**	**0.08741**	**0.08828**
VAR I	Forecast	0.08578	0.08189	0.08043	0.08160
	Forecast error	-0.001349	0.006462	0.006977	0.006679
VAR III	Forecast	0.09133	0.09384	0.09706	0.09797
	Forecast error	-0.006900	-0.005488	-0.009651	-0.008786

SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS

		Estimation period: (1970 -1993)	Forecast period: (1994 –1997)
VAR I	MAE	0.008054	0.005367
	MSE	0.000088	0.000034
	RMSE	0.009365	0.005850
VAR III	MAE	0.006276	0.007706
	MSE	0.000060	0.000062
	RMSE	0.007770	0.007875

For the estimation period (1969-1993) and for all three expenditure share equations, VAR III model is shown to be statistically more accurate than VAR I model. Nevertheless, the forecast errors and all the quality criteria indicate VAR I model to be a better forecasting device than VAR III. Yet, as the differences between the two sets of summary statistics are small (never exceeding 0.02), we can consider both models to have fairly similar accuracy in forecasting the UK demand for tourism.

The fact that the variables included in VAR I and VAR III models are I(1) implies that estimation, statistical tests and forecasting procedures based on them are strictly valid if, and only if, cointegrating relationship(s) link the variables. This means that we must conduct statistical tests which can establish whether long-run steady state relationship(s) exist within the VARs. Indeed, as Harris (1995, p. 76) points out, "*the implication that nonstationary variables can lead to spurious regressions unless at least one cointegrating vector is present means that some form of testing for cointegration is almost mandatory*".

However, the 'mandatory' testing for cointegration must be performed within a system of equations framework and not on an equation-by-equation basis. Certainly, given the consequences for the single equation approach if more than one cointegrating vector exists, earlier use of cointegration analysis within the single equation framework (performed on the ARDL models of chapter 4) must give way to the determination of cointegration rank within a VAR system of equations approach. Therefore, the cointegration rank test of Johansen procedure becomes the inevitable next step in proceeding the empirical analysis of the UK tourism demand for France, Spain and Portugal, within a vector autoregressive framework.

Besides establishing the procedure for cointegration rank testing, the Johansen approach provides a general framework for identification, estimation and hypothesis testing in cointegrated systems (see, for example, Johansen, 1988 and Johansen and Juselius, 1990). Hence, given a cointegrated VAR system, this approach provides a method for identifying the long-run structural relationships underlying the unrestricted reduced-form of the VAR, and permits all kinds of hypothesis testing involving its structural parameters. This is of paramount importance for validly using a VAR specification for both forecasting and economic analysis purposes.

However, the "empirical identification" process suggested by Johansen to exactly-identify the structural long-run coefficients can prove inadequate, particularly in economic contexts where theory provides strong, sensible and testable restrictions. In these cases, the cointegrating vectors must be subject to exact- and over-identifying restrictions obtained from economic theory and other relevant *a priori* information, rather than being subject to some normalisation process which does not consider the theoretical and empirical framework within which the phenomenon under analysis evolves.[8] Using the UK labour market model of Clements and Mizon (1991) as an example, Harris (1995) adopts the same line of reasoning stating that "*what is becoming increasingly obvious is the need to ensure that prior information motivated by economic arguments forms the basis for imposing restrictions, and not the other way around*" (p.117). This is particularly important in a demand context involving n tourism shares, n tourism prices and the origin's per capita real expenditure. In this case, the number of steady state equilibrium relationships predicted by theory is, precisely, the (n−1) share equations incorporated in an orthodox AIDS model.

Therefore, if theory is right, the cointegration tests involving a VAR of the UK tourism demand with variables WF_t, WS_t, PP_t, PS_t, PF_t and E_t (with or without dummy variables) must indicate that the only relevant long-run relationships existing in the model are those established by the two share equations. Furthermore, the exact-identification

[8] For a comparison of Johansen's 'empirical identification' process with the AIDS theoretical exact-identifying restrictions in a VAR model of budget shares, see Pesaran and Shin (1997).

process of the structural coefficients must confirm that their steady state form is that of the equations in a static AIDS model. Consequently, in either the VAR I or VAR III models, we expect to find exactly two cointegrating vectors and to identify the structural parameters with restrictions that match those of the normalisation process used for identifying the share equations of a static AIDS model. If this is the case, then we can subject the long-run equilibrium relationships to further hypothesis testing, such as homogeneity and symmetry, and contribute an empirical basis for the validation of the principals of consumer behaviour theory within an AIDS system framework, using a VAR specification. This would complete our goal in this chapter and provide theoretical, empirical and statistical evidence of the robustness of the AIDS system in offering a sound and reliable basis for the analysis of long-run equilibrium relationships between destinations' tourism shares, their tourism prices and a specific origin's tourism budget.

In the next section, we use the Johansen procedure to establish the number of cointegrating vectors in VAR I and VAR III models and specify restricted models for the estimation of the long-run coefficients of interest. We then use the restricted models for economic analysis and forecasting purposes and compare the results with those obtained in this section.

7.4. THE JOHANSEN PROCEDURE

Conditioning a model on the (weak) exogeneity of a sub-set of variables implies a reduction of the system from an n-variables dimension to a n_1-variables dimension ($n_1 < n$). If (weak) exogeneity holds, this ensures that such reduction is without loss of information on the parameters of interest. Therefore, in the terminology of the 'general-to-specific' approach, VAR I is the general system and VAR III is a specific case of the more general model. In addition, including dummy variables and/or imposing a 'structure' to the general unrestricted VAR, via exactly and over-identifying restrictions, implies other specifications which also constitute special cases of the 'more general' VAR I model. Whichever the specifications are, if nonstationary variables are present, the statistical inference based on them needs to be sanctioned by evidence of at least one cointegrating vector. This evidence can be provided by the Johansen's reduced rank regression method.

The Johansen maximum likelihood approach is applied to both the general VAR I and specific VAR III models. In this way, besides finding how many cointegrating vectors exist in each model, we can also test the relevance of several constraints imposed on their parameters, and the legitimacy of conditioning the models, on a set of exogenous variables.

The application of the Johansen procedure to the VAR models of the UK demand for tourism includes the following aspects:

1. Testing the order of integration of each variable included in the VAR.[9]
2. Testing the appropriate lag-length of the VAR in order to assure white noise errors.
3. Determining whether the equations in the system should include deterministic components such as an intercept and/or a time trend.
4. Determining whether the equations in the system should include dummy variables.
5. Testing for reduced rank to find out how many cointegrating vectors exist.
6. Testing for exactly- and over-identifying restrictions on the equations' parameters in the cointegrated VAR model.

All these steps ensure a correct specification of the VAR and the possibility of determining the structural relationships embedded in the cointegrating vectors. It also allows for testing theoretical hypotheses involving single and cross-equation parameters, and permits the economic interpretation of the long-run coefficients. Therefore, we follow these steps in order to establish the existence and estimate the structural components of cointegrating vectors within a vector autoregressive framework for the UK demand for tourism in France, Spain and Portugal.

Let z_t be a vector of n potentially endogenous variables. Then, it is possible to model z_t as an unrestricted VAR involving up to p-lags such that (see Harris, 1995, p.77):

$$z_t = A_1 z_{t-1} + \ldots + A_p z_{t-p} + u_t \qquad u_t \rightarrow IN(0, \Sigma) \tag{7.16}$$

where z_t is a (n x 1) vector and each of the A_i is a (n x n) matrix of parameters.

The VAR model in (7.16) can be reformulated as a vector error-correction (VEC) model such that:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \ldots + \Gamma_{p-1} \Delta z_{t-p+1} + \Pi z_{t-p} + u_t \tag{7.17}$$

where $\Gamma_i = -(I - A_1 - \ldots - A_i), i = 1, \ldots, p-1$ and $\Pi = -(I - A_1 - \ldots - A_p)$.

This specification provides information on both the short- and long-run coefficients, via the estimates of, respectively, Γ and Π, and it is the basis for the implementation of the

[9] Although it is possible to find cointegrated relationship(s) when there is a mix of I(0), I(1) variables, the standard Johansen approach is designed to handle I(1) variables and is more reliable under these circumstances.

Johansen procedure. Matrix Π can also be described as $\Pi = \alpha\beta'$, where α represents the speed of adjustment matrix and β is the matrix of long-run coefficients such that, $\beta' z_{t-p}$ implicit in (7.17) represents up to (n−1) cointegrating vectors in the multivariate VAR. These cointegrating vectors define the relationships which ensure that the variables in z_t converge to their long-run solutions.

If the variables in z_t are nonstationary I(1) variables then, Πz_{t-p} in (7.17), containing the long-run relationships, must be stationary for $u_t \sim I(0)$ to be a white noise. This occurs when $\Pi = \alpha\beta'$ has reduced rank, that is, when there are $r \leq (n-1)$ cointegrating vectors in β (meaning that r columns in β form r linearly independent stationary combinations of the variables in z_t,) alongside (n−r) nonstationary vectors. Hence, testing for cointegration amounts to determining the number of r linearly independent columns in Π (the rank of Π). By the same token, testing for cointegration amounts to determining if the last (n−r) columns of α are zero. In fact, once $r \leq (n-1)$ cointegrating vectors are found, we can infer that the last (n−r) columns of α are zero and, consequently, the last (n−r) columns of β are nonstationary. This means that they do not enter the specification (7.17). In this case, we can reduce α and β dimension to (n x r).

As noted by Chambers (1993), in many cases, and in particular with demand systems, "*the focus of interest is on a set of hypotheses relating to the long-run structure of the model, which is quite independent of any short-run dynamics fitted to the empirical model*" (p.727). Likewise, in the case of the UK tourism demand VAR system, our interest is focused on the structural equations upon which parameter restrictions are to be imposed, testing theoretical hypotheses about the long-run equilibrium relationships.

In sections 7.3.1. and 7.3.2. we established that all variables involved in the VAR system were I(1) and that the appropriate lag-length was one. We considered that an intercept, representing the average value of the left-hand side variables when all the right-hand side variables are set to zero, was a relevant deterministic component and should be added to integrate (exclusively) the long-run form of the model. In addition, we assumed that no time trends should be included in the specification, as no significant trend co-movements were detected in the time series. Finally, we found sufficient empirical evidence for not rejecting the possibility of treating variable E_t as an exogenous variable, and for integrating dummy variables D1, D2 and D3 as statistically relevant regressors. Hence, we end up with VAR specifications such as the VAR I and VAR III models.

Theory predicts that there should be two cointegrating relationships among the variables included in VAR I and VAR III. This hypothesis was tested by using the Johansen reduced-rank test for determining the number of cointegrating vectors. The Johansen procedure consists of obtaining estimates for matrixes α and β using the reduced rank regression approach that provides n eigenvalues $\hat{\lambda}_1 > \hat{\lambda}_2 .... > \hat{\lambda}_n$ and their corresponding eigenvectors $\hat{V} = (\hat{v}_1 ..... \hat{v}_n)$. Those r elements in V, which determine the linear combinations of stationary relationships, are the cointegrated vectors. Therefore, testing the hypothesis that there are at most r cointegrating vectors (and, hence, n−r nonstationary relationships) implies testing the null H_0: $\lambda_i = 0$; $i = r+1,..., n$.

This hypothesis can be tested using a statistic known as the eigenvalue trace statistic (λ_{trace}) which tests the null that r = q (q = 0, 1,..., n−1) against the alternative that r ≥ q + 1. Another test for the number of cointegrating vectors is known as the maximum eigenvalue statistic (λ_{max}) which tests the null that there are r = q cointegrating vectors against the alternative that r = q + 1 exist. Asymptotic critical values for these statistics are provided in most computer packages used for time series analysis.

The results obtained from applying Johansen's reduced rank test to VAR I and III models are given in Table 7.10. The first column shows the eigenvalues associated with each of the I(1) variables, ordered from highest to lowest, necessary to calculate the λ_{max} and λ_{trace} statistics. The second column shows the various hypotheses to be tested, starting with no cointegration (r = 0 or n−r = 6 in the case of VAR I, and r = 0 or n−r =5 in the case of VAR III) and followed by increasing numbers of cointegrating vectors. The following columns present the estimated λ_{max} and λ_{trace} statistics with the respective 5% and 10% critical values. The last column presents the SBC model selection criterion.

Table 7.10: Tests of the cointegration rank for VAR I and VAR III models

Eigen values	H_0		$\hat{\lambda}_{max}$ statistic	λ_{max} critical		$\hat{\lambda}_{trace}$ statistic	λ_{trace} critical		SBC
	r	n − r		5%	10%		5%	10%	
VAR I									
λ_1=0.9201	r = 0	n−r = 6	**70.76**	40.53	37.65	**153.96**	102.56	97.87	274.71
λ_2=0.7489	r = 1	n−r = 5	**38.69**	34.40	31.73	**83.20**	75.98	71.81	290.09
λ_3=0.5889	r = 2	n−r = 4	24.89	28.27	25.80	44.50	53.48	49.95	**292.78**
λ_4=0.3273	r = 3	n−r = 3	11.10	22.04	19.86	19.61	34.87	31.93	291.89
λ_5=0.1838	r = 4	n−r = 2	5.69	15.87	13.81	8.51	20.18	17.88	287.45
λ_6=0.0958	r = 5	n−r = 1	2.82	9.16	7.53	2.82	9.16	7.53	283.63
VAR III									
λ_1=0.8798	r = 0	n−r = 5	**59.31**	46.77	43.80	**142.48**	119.77	114.38	265.82
λ_2=0.7734	r = 1	n−r = 4	**41.57**	40.91	38.03	83.16	90.60	85.34	272.15
λ_3=0.5833	r = 2	n−r = 3	24.51	34.51	31.73	41.59	63.10	59.23	**272.94**
λ_4=0.2767	r = 3	n−r = 2	9.07	27.82	25.27	17.08	39.94	36.84	268.54
λ_5=0.2489	r = 4	n−r = 1	8.01	20.63	18.24	8.01	20.63	18.24	259.74

For VAR I model, at the 5% significance level, both the λ_{max} and λ_{trace} statistics give evidence of two cointegrating vectors corresponding to the higher eigenvalues attached to the share equations for France and Spain. In fact, both statistics associated with the null of r = 0 and r = 1 reject these hypotheses (statistic value > critical value) but cannot reject r = 2 (statistic value < critical value). The SBC criterion further supports these findings by selecting the model which contains two cointegrated relationships.

For VAR III model, at the 5% significance level, the λ_{max} statistic suggests the existence of two cointegrating vectors while the λ_{trace} statistic does not reject the hypothesis of only one cointegrating vector in VAR III model. This apparent disagreement in the tests for cointegrating rank is not uncommon, particularly in cases of small samples and added dummy variables. So, a choice must be done at this stage. As Harris (1995, p.89) suggests "*it is important to use any additional information that can support the choice of r*" and we do have additional information that can support the choice of r. We have one test, the λ_{max} statistic, clearly rejecting the existence of only one in favour of two cointegrating vectors; we have the SBC criterion selecting the VAR III model with two cointegrating vectors. We have theory suggesting the existence of two and not one long-run relationships which is unmistakably supported by both test statistics and selection criteria in the more general specification of VAR I model. Hence, given that evidence against theory prediction seems to be rather weak, we proceed by setting $r = 2$ for both VAR I and VAR III models.

Considering that both VAR I and VAR III models comply with theory with respect to the number of cointegrating vectors predicted, we must now identify the structural long-run coefficients of the two cointegrating vectors. In order to do so, we follow Pesaran and Shin (1997) and use the theoretical exact-identifying restrictions implicit in the specification of the share equations of a static AIDS model.

Consider the following notation for the matrix β of the two cointegrating vectors of VAR I model for variables WF_t WS_t PP_t PS_t PF_t E_t and intercept,

$$\beta'_{I} = \begin{bmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} & \beta_{51} & \beta_{61} & \beta_{71} \\ \beta_{12} & \beta_{22} & \beta_{32} & \beta_{42} & \beta_{52} & \beta_{62} & \beta_{72} \end{bmatrix}$$

and for the matrix β of the two cointegrating vectors of VAR III model for variables WF_t, WS_t, PP_t, PS_t, PF_t, E_t, D1, D2, D3 and intercept

$$\beta'_{III} = \begin{bmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} & \beta_{51} & \beta_{61} & \beta_{71} & \beta_{81} & \beta_{91} & \beta_{101} \\ \beta_{12} & \beta_{22} & \beta_{32} & \beta_{42} & \beta_{52} & \beta_{62} & \beta_{72} & \beta_{82} & \beta_{92} & \beta_{102} \end{bmatrix}$$

The theoretical restrictions which exactly-identify the cointegrating vectors as the share equations of an AIDS specification in both VAR I and VAR III models are given by:

$$H_{AIDS} : \begin{Bmatrix} \beta_{11} = -1 & \beta_{12} = 0 \\ \beta_{21} = 0 & \beta_{22} = -1 \end{Bmatrix}$$

The exactly identified estimates of the two cointegrating vectors of VAR I and VAR III models are given in Table 7.11 (asymptotic t ratios in brackets). The third vector, corresponding to the share equation for Portugal, is retrieved from the estimates of the other two by applying the adding-up property of the AIDS model.

Table 7.11: Long-run coefficient estimates of the exactly-identified share equations

Variables	VAR I			VAR III		
	Vector 1 (WF)	Vector 2 (WS)	Vector 3 (WP)	Vector 1 (WF)	Vector 2 (WS)	Vector 3 (WP)
WF	-1	0		-1	0	
WS	0	-1		0	-1	
PP	-0.4965 (-0.86)	0.1781 (0.60)	0.3184 (0.95)	-0.0027 (-0.05)	0.1090 (1.69)	-0.1062 (-3.28)
PS	0.8214 (2.09)	-0.5937 (-2.97)	-0.2277 (-0.97)	0.2256 (4.68)	-0.3075 (-5.22)	0.0820 (2.90)
PF	-0.2244 (-0.48)	0.3119 (1.20)	-0.0875 (-0.33)	-0.3394 (-3.59)	0.3044 (2.57)	0.0350 (0.56)
E	-0.0684 (-1.12)	0.0159 (0.51)	0.0525 (1.45)	0.0153 (2.33)	-0.0183 (2.29)	0.0030 (0.78)
D1				0.0380 (5.27)	-0.0154 (-1.74)	-0.0226 (-5.18)
D2				-0.0565 (-3.65)	0.0528 (2.78)	0.0037 (0.41)
D3				0.0574 (5.17)	-0.0590 (-4.34)	0.0016 (0.24)
INT	0.8309 (2.20)	0.3818 (1.97)	-0.2127 (-0.95)	0.3687 (16.26)	0.5443 (19.50)	0.0870 (6.33)

There is a sharp difference between the estimates of the long-run coefficients obtained with VAR I and with VAR III models, both in magnitude and statistical significance and in the signs which should accord with theoretical expectations. For instance, in the share equation for Portugal, the estimates of VAR I model indicate all coefficients to be statistically insignificant at the 5% level, and in the share equations for France and Spain only the price of Spain and the intercept to be significant at the same level. In addition, in the share equation for Portugal, the own-price and the intercept estimates have and 'wrong' signs. Both these estimates and the intercept estimate in the equation for France present implausible magnitudes

However, in VAR I model, the price of Portugal in the equation for France and the price of France in the equation for Portugal are not significantly different from zero, as predicted in our conjecture about the competitive behaviour of these neighbouring destinations. This same feature is supported by the corresponding estimates obtained with VAR III model. Yet, this is as far as similarity goes between the estimation results of the two models. The estimation results of VAR III model are overall statistically significant, present the expected signs and magnitudes and give accurate and plausible information about the ways the events represented by the dummy variables affect the UK demand for tourism in the three destinations. Indeed, the coefficients of D1 indicate that the political changes in Portugal and Spain and the oil crises, affected significantly and negatively the UK tourism flows of these destinations, favouring France instead during the period 1974-1981. The coefficients of D2 indicate that the integration process of Spain and Portugal in the EU caused UK tourism flows to divert from France to the Iberian peninsula, although favouring Spain more than Portugal. The coefficients of D3, representing the post-integration period, indicate a recovery of the share for France at the expense of Spain's share which steadily declines.[10] In the same period, the share for Portugal shows an increase, although not statistically significant.

Theory suggests that restrictions corresponding to the features of homogeneity and symmetry should hold within a system of demand equations, reflecting the rationality of consumer behaviour. In addition, according to our own assumption of competitive behaviour between neighbouring tourism destinations, tourism price changes in France (Portugal) should not affect the UK demand for tourism in Portugal (France), while price changes in Spain should affect significantly the UK demand for tourism both in France and in Portugal. These hypotheses were tested by imposing the appropriate restrictions on both VAR I and VAR III models and using the LR statistic to evaluate the result. If the LR statistic is smaller than the appropriate critical value, the restricted model cannot be rejected. Otherwise, the restrictions do not hold. Table 7.12 shows the tests results for the hypothesis of null cross-price effects between the share equations for France and Portugal, homogeneity and symmetry and both these sets of hypotheses simultaneously.

The tests indicate that the set of hypotheses cannot be rejected in either VAR I or VAR III models. Therefore, both models comply with the theoretical restrictions of homogeneity and symmetry and with our assumption of null cross-price effects between the share equations for France and Portugal.

[10] The opening of the Channel Tunnel in 1994 may also have contributed to this result.

Table 7.12: Tests for over-identifying hypothesis in VAR I and VARIII models

Hypothesis	Model	LR→χ^2(i)	5% Critical value
Null cross-price effects	VAR I	$\chi^2(2)=1.53$	5.99
	VAR II	$\chi^2(2)=0.36$	
Homogeneity & symmetry	VAR I	$\chi^2(3)=3.31$	7.81
	VAR II	$\chi^2(3)=7.28$	
Homogeneity & symmetry & null cross-price effects	VAR I	$\chi^2(4)=4.01$	9.49
	VAR II	$\chi^2(4)=7.53$	

The estimates for the two cointegrating vectors of VAR I and VAR III models under the exact- and over identifying restrictions considered above are given in Table 7.13 (asymptotic t ratios in brackets). As before, the third vector corresponding to the share equation for Portugal, is retrieved from the estimates of the other two vectors by applying the adding-up property of the AIDS model. There is a general improvement in the estimation results of VAR I model, in the sense that they are now statistically more relevant and seem more plausible than before. However, the model still presents some incongruent results such as, for example, in the share equation for Portugal, a positive own-price coefficient and a negative intercept. Moreover, the irrelevance of the coefficients of variable E_t (UK real per capita expenditure) in the equations for France and Spain, compounded with the statistical significance of this variable's coefficient in the equation for Portugal seems questionable, given empirical evidence found previously with the AIDS models of chapters 5 and 6. Hence, we believe that VAR I model is not the ideal specification for explaining the UK tourism demand behaviour for these destinations. In contrast, the estimation results of VAR III are overall statistically robust, theoretically consistent and empirically plausible, indicating this model as an adequate vehicle for analysing the UK tourism demand in France Spain and Portugal.

Table 7.13: Long-run estimates of the share equations under exact- and over-identifying restrictions of homogeneity, symmetry and null cross-price effects for France and Portugal

Variables	VAR I			VAR III		
	Vector 1 (WF)	Vector 2 (WS)	Vector 3 (WP)	Vector 1 (WF)	Vector 2 (WS)	Vector 3 (WP)
WF	-1	0		-1	0	
WS	0	-1		0	-1	
PP	0	-0.1308 (-1.76)	0.1308 (1.76)	0	0.0895 (5.80)	-0.0895 (-5.80)
PS	0.6037 (6.51)	-0.4729 (-4.97)	-0.1308 (-1.76)	0.2891 (4.79)	-0.3785 (-5.97)	0.0895 (5.80)
PF	-0.6037 (-6.51)	0.6037 (6.51)	0	-0.2891 (-4.79)	0.2891 (4.79)	0
E	-0.0362 (-3.40)	-0.0001 (-0.01)	0.0361 (3.53)	0.0091 (1.16)	-0.0121 (-1.38)	0.0030 (0.95)
D1				0.0354 (3.83)	-0.0115 (-1.13)	-0.0239 (-7.05)
D2				-0.0373 (-2.36)	0.0360 (2.02)	0.0013 (0.20)
D3				0.0429 (4.56)	-0.4184 (-3.95)	-0.0011 (-0.27)
INT	0.6152 (10.77)	0.4916 (7.12)	-0.1068 (-1.76)	0.3849 (13.17)	0.5230 (16.96)	0.092 (11.69)

A detailed investigation of these results requires the analysis of the relevant elasticity values. We estimate the expenditure, own- and cross-price uncompensated elasticities of the tourism shares, based on the long-run coefficients estimates of VAR III model, under exact- and the over-identifying restrictions of homogeneity, symmetry and null cross-price coefficients between the share equations for France and Portugal. The elasticities are computed using the formulae given in chapter 5. Table 7.14 presents these elasticities estimates and, for comparison purposes, the corresponding long-run elasticities estimates obtained with the AIDS models of chapter 5 and 6 only for the second period (the last two decades), as these estimates represent more recent behaviour of the UK demand for tourism.

Table 7.14. Expenditure and uncompensated price elasticities derived from VAR III

	Models	Expenditure elasticities	Own-price elasticities	Cross-price elasticities		
				PP	PS	PF
WP	VAR III	1.039	-2.158	X	1.137	-0.017
	AIDS (2nd period)	0.947	-1.797	X	0.830	0.019
	Dynamic AIDS (2nd period)	1.027	-1.729	X	0.714	-0.011
WS	VAR III	0.979	-1.057	0.161	X	0.523
	AIDS (2nd period)	1.150	-1.933	0.124	X	0.658
	Dynamic AIDS (2nd period)	1.047	-2.150	0.120	X	0.983
WF	VAR III	1.026	-1.817	-0.002	0.793	X
	AIDS (2nd period)	0.808	-1.901	0.017	1.077	X
	Dynamic AIDS (2nd period)	0.929	-2.336	0.007	1.400	X

The elasticities' estimates of the VAR III and the AIDS model of chapter 5 are remarkably similar with one exception: the own-price elasticity in the equation for Spain.[11] Indeed, the estimates of the expenditure elasticities have values close to unity in every model and for all destinations. This accords with the theoretical postulates predicting unitary long-run expenditure (income) elasticities. Except in the equation for Spain of the VAR model, the estimates of the own-price elasticities have values close to -2. If the VAR estimate for this elasticity in the case for Spain is to be trusted, Spain has a comparative advantage relative to its neighbours in the long-run, as its own-price elasticity estimate is roughly half that of its neighbours. The cross-price elasticities' estimates of the VAR give the same indication as those of the AIDS specifications; for instance, insignificant cross-price effects between the equations of France and Portugal and significant cross-effects indicating competitive behaviour between the equations for France and Spain, on the one hand, and for Spain and Portugal, on the other. As before, the estimates of these elasticities

[11] The dynamic AIDS model over-estimates the price elasticities (both own- and cross-price) relative to the VAR model in the cases for France and Spain, and under-estimates these elasticities in the case of Portugal. As will be shown in chapter 8, the VAR III model is the best forecaster in the cases of France and Spain while in the case for Portugal, the dynamic AIDS model provides the best forecasts. As the forecasting performance of an econometric model is also a selection criterion, it seems reasonable to infer that the elasticities estimates provided by the VAR III model are more trustworthy than those provided by the dynamic AIDS in the cases for France and Spain, while the reverse is acceptable in the case of Portugal.

suggest that, in the long-run, the UK demand for Portugal or France is more sensitive to price-changes in Spain than that for Spain is to price changes in France or Portugal.

This similarity of results should be expected as the VAR III specification fully complies with the theoretical predictions underlying the AIDS approach. Hence, the results support the AIDS specification as an empirically and theoretically robust means for economic analysis of the UK demand for tourism. Moreover, the econometric methodology applied to obtain these estimates substantiates the importance of the VAR approach and the Johansen procedure in finding statistical support for the existence of cointegrated vectors, which sanction the share equations of the AIDS system as the only meaningful long-run relationships among the variables included. However, a cautionary note should be added for practitioners of VAR model building: the "empirical identification" process of the Johansen procedure is not the appropriate means to exactly-identify the cointegrated relationships in the case of a system of tourism share equations. Instead, the identification of the long-run structural coefficients of tourism share equations requires *a priori* information provided by economic theory and knowledge of the destinations considered. Market conditions, policy regulations, institutional characteristics, geographical attributes and location, and social and political features need to be modelled and appropriately tested within the restricted VAR framework, to obtain theoretically consistent, empirically plausible and statistically reliable estimates of the structural parameters underlying the long-run equilibrium relationships in a VAR. Only by using this information were we able to establish with reasonable accuracy, the long-run equilibrium path of the UK demand for tourism in France, Spain and Portugal. However, our analysis is not complete without information on the forecasting performance of the cointegrated VAR models. This is our next task.

7.5. FORECASTING WITH COINTEGRATED VAR SPECIFICATIONS

Although the cointegrated VAR I specification did not seem to be the best means for explaining the UK tourism demand in France, Spain and Portugal it may, nevertheless, be a very competent forecasting device. For assessing its forecasting ability and comparing it with that of the cointegrated VAR III, we estimate both models for the in-sample period 1969-1993, leaving the last four observations to integrate the forecasting period. In order to obtain a complete picture of the predictive ability of these models, we use two versions of the cointegrated VAR models for forecasting the destinations' expenditure shares: the VAR models under exactly-identifying restrictions alone (denoted by 'exact-VAR'); and the VAR models under the full set of exactly- and over-identifying restrictions (denoted by 'over-VAR'). Tables 7.15., 7.16. and 7.17 show the actual and forecast values of the expenditure shares of France, Spain and Portugal, respectively, obtained with the cointegrated exact-VAR I and over-VAR I and with the cointegrated exact-VAR III and over-VAR III models. As before, the tables report the forecast errors and the set of summary statistics (MAE, MSE and RMSE) used previously.

Table 7.15: Forecasting results for the UK expenditure share of France

Actual values		1994	1995	1996	1997
		0.39700	**0.38540**	**0.38967**	**0.40481**
Exact-VAR I	Forecast	0.34988	0.33138	0.32715	0.32610
	Forecast error	0.047117	0.054027	0.062526	0.077871
Over-VAR I	Forecast	0.36319	0.36478	0.36285	0.36129
	Forecast error	0.033805	0.020626	0.026820	0.043523
Exact-VAR III	Forecast	0.38011	0.37760	0.37451	0.37335
	Forecast error	0.016885	0.007803	0.015162	0.031463
Over-VAR III	Forecast	0.39505	0.39260	0.39485	0.39475
	Forecast error	0.001942	-0.007194	-0.005178	0.010054

SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS			
		Estimation period: (1970 -1993)	Forecast period: (1994 –1997)
Exact-VAR I	MAE	0.016278	0.060595
	MSE	0.000358	0.003811
	RMSE	0.018907	0.061732
Over-VAR I	MAE	0.017425	0.031194
	MSE	0.000456	0.001045
	RMSE	0.021364	0.032333
Exact-VAR III	MAE	0.013428	0.017828
	MSE	0.000256	0.000391
	RMSE	0.015990	0.019785
Over-VAR III	MAE	0.016342	0.006092
	MSE	0.000362	0.000046
	RMSE	0.019016	0.006772

Table 7.16: Forecasting results for the UK expenditure share of Spain

Actual values		1994	1995	1996	1997
		0.51857	**0.52625**	**0.52292**	**0.50691**
Exact-VAR I	Forecast	0.56316	0.58313	0.58840	0.58956
	Forecast error	-0.044588	-0.056878	-0.065126	-0.082649
Over-VAR I	Forecast	0.54121	0.54052	0.54303	0.54514
	Forecast error	-0.022636	-0.014275	-0.020117	-0.038235
Exact-VAR III	Forecast	0.53975	0.55446	0.55924	0.56056
	Forecast error	-0.021177	-0.028212	-0.036326	-0.053649
Over-VAR III	Forecast	0.51884	0.52718	0.52568	0.52543
	Forecast error	-0.000264	-0.000932	-0.002761	-0.018518

SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS			
		Estimation period: (1970 -1993)	Forecast period: (1994 –1997)
Exact-VAR I	MAE	0.017295	0.062310
	MSE	0.000422	0.004074
	RMSE	0.020540	0.063827
Over-VAR I	MAE	0.023393	0.023816
	MSE	0.000723	0.000646
	RMSE	0.026893	0.025410
Exact-VAR III	MAE	0.014235	0.034841
	MSE	0.000289	0.001361
	RMSE	0.017006	0.036886
Over-VAR III	MAE	0.015845	0.005619
	MSE	0.000371	0.000088
	RMSE	0.019268	0.009374

Table 7.17: Forecasting results for the UK expenditure share of Portugal

Actual values		1994	1995	1996	1997
		0.08443	**0.08835**	**0.08741**	**0.08828**
Exact-VAR I	Forecast	0.08696	0.08550	0.08481	0.08434
	Forecast error	-0.002530	0.002850	0.002600	0.003942
Over-VAR I	Forecast	0.09560	0.09470	0.09411	0.09357
	Forecast error	-0.011169	-0.006352	-0.006703	-0.005288
Exact-VAR III	Forecast	0.08014	0.06794	0.06625	0.06610
	Forecast error	0.004292	0.020408	0.021164	0.022185
Over-VAR III	Forecast	0.08611	0.08022	0.07947	0.07982
	Forecast error	-0.001678	0.008124	0.007939	0.008463

SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS			
		Estimation period: (1970 -1993)	Forecast period: (1994 –1997)
Exact-VAR I	MAE	0.008959	0.002980
	MSE	0.000136	0.000009
	RMSE	0.011647	0.003034
Over-VAR I	MAE	0.009119	0.007378
	MSE	0.000133	0.000060
	RMSE	0.011520	0.007713
Exact-VAR III	MAE	0.005628	0.017012
	MSE	0.000055	0.000344
	RMSE	0.007440	0.018540
Over-VAR III	MAE	0.006155	0.006551
	MSE	0.000063	0.000051
	RMSE	0.007914	0.007132

Comparison and the choice of the best forecasting model is facilitated by the inclusion, in Table 7.18, of the three accuracy statistics for the forecasting period of all the VAR models used in this chapter. The lower the value of the statistic, the more accurate are the forecasts. Therefore, to make the choice of the best forecasting model easier, we signal the first, second and third lowest values of the statistics with, respectively,. •, * and °.

Table 7.18: Forecast accuracy statistics for the different VAR specifications

Expenditure shares	Statistics	VAR MODELS					
		Unrestricted VAR I	Exact-VAR I	Over-VAR I	Unrestricted VAR III	Exact-VAR III	Over-VAR III
France	MAE	0.0250°	0.0606	0.0312	0.0366	0.0178*	0.0061•
	MSE	0.0009°	0.0038	0.0010	0.0019	0.0004*	0.0000•
	RMSE	0.0302°	0.0617	0.0323	0.0435	0.0198*	0.0068•
Spain	MAE	0.0228*	0.0623	0.0238°	0.0390	0.0348	0.0056•
	MSE	0.0009°	0.0041	0.0006*	0.0020	0.0014	0.0001•
	RMSE	0.0294°	0.0638	0.0254*	0.0452	0.0369	0.0094•
Portugal	MAE	0.0054*	0.0030•	0.0074	0.0077	0.0170	0.0066°
	MSE	0.0000*	0.0000•	0.0001	0.0001	0.0003	0.0001°
	RMSE	0.0059*	0.0030•	0.0077	0.0079	0.0185	0.0071°

The over-VAR III model appears to be the best forecasting device for the shares of France and Spain and the third best for the share of Portugal. For this share, however, the differences between the values of the quality statistics for the different VAR models are so small that any of them can be considered as an excellent forecaster. Hence we can consider the VAR III specification under the full set of theoretical restrictions as the best device for forecasting the UK demand for tourism in France, Spain and Portugal.

However, it should be noted that the unrestricted VAR I specification is qualified as the second best forecasting model. This brings about the matter of the forecasting ability of a "pure" VAR model alongside its remarkable qualities of simplicity of form, estimation ease and minimal assumptions. If the main purpose of a research project is to provide reliable forecasts of expenditure share variables, the estimation of a general (unrestricted) VAR specification with the appropriate variables and lag-length is the preferable approach, since the quality differences of the forecasts do not justify undergoing the complexity and difficulty of implementing the Johansen approach. Yet, the researcher must be aware that the validity of statistical inference based on reduced-form VAR systems depends on the stationarity of the time series included or, if these are nonstationary, on the existence of cointegrating relationships(s) underlying the long-run equilibrium path of the variables.

7.6. CONCLUSION

Economic models portraying steady state long-run relationships between time series variables are a central aspect of theoretical and empirical research. However, these models are often subject to debate, in the sense that they may include alternative theoretical hypotheses about the ways in which the relevant variables affect each other. Part of the role of applied work is to establish appropriate formal specifications which can be considered valid means of estimating the long-run equilibrium path of the relevant economic variables and of testing competing theoretical hypotheses within those specifications.

Over the last few decades, the validity of modelling and estimation procedures involving nonstationary time series within the single equation and the traditional structural system of equations approaches has been questioned. The estimation of long-run relationships involving nonstationary data may lead to spurious results, invalidating statistical inference and forecasting procedures. Moreover, pre-assuming exogeneity in variables that may be jointly determined and zero restrictions with no economic basis can also lead to invalid inference and forecasting procedures.

Steps towards resolving these problems were taken with the work of Sims (1980, 1982), Engle and Granger (1987), Johansen (1988, 1995), Johansen and Juselius (1990) and others, who introduced, developed and applied such concepts as cointegration analysis and vector autoregression (VAR) specifications, adding a new dimension to the methodological processes of modelling and estimating economic relationships. This new methodology, applied to a system of equations, can be used to model, estimate and test its structural long-run parameters, avoiding the problems of spurious regression, simultaneous equations bias, *a priori* division of endogenous-exogenous variables and unfounded zero restrictions.

The main line of reasoning in cointegration analysis is that statistical inference based on a system of equations which includes nonstationary time series can only be valid if one or more cointegrating relationships exist among its variables. Hence, appropriate cointegration tests must be applied in order to determine whether a system is cointegrated and, if so, proceed with suitable estimation and testing techniques which can provide valid estimates of its structural parameters and reliable information about its forecasting ability.

The main line of reasoning in the VAR approach is that "incredible" zero restriction and inappropriate endogenous-exogenous *a priori* division of variables can lead to invalid estimation procedure and inaccurate forecasting results. Hence, the estimation of reduced-form (unrestricted) systems treating all variables as endogenous, has been proposed as a feasible and reliable alternative to the traditional structural approach, particularly if forecasting is the main objective of the modelling exercise.

The Johansen methodology can be viewed as the integration of these two lines of reasoning in that it establishes the cointegration rank test procedure for determining the number of cointegrating relationships in a system, and provides a general framework for identification, estimation and hypothesis testing of cointegrated vectors within a VAR modelling approach. However, the identification process used by Johansen is not always appropriate, and in economic contexts where theory provides testable restrictions, these should be used to identify the structural parameters of the cointegrated relationships.

The AIDS models specified in chapters 5 and 6 are systems of equations which include nonstationary data series and assume exogeneity for the right-hand side variables. The implications of these features for the empirical analysis of the UK demand for tourism in France, Spain and Portugal could seriously damage the validity of the estimation methods, statistical inference and forecasting procedures used, if no cointegrating relationships exist among the variables of such models. Hence, the objective of this chapter is to find empirical evidence for validating (or otherwise) the inference procedures and the estimation results obtained with the AIDS specifications of previous chapters. For this purpose, we specified a reduced-form unrestricted VAR system for the relevant variables included in the AIDS model. We used several statistical tests to establish the lag-length, deterministic components and the endogenous/exogenous division of the variables that should include the general VAR specification. Once the appropriate form of the VAR was in place, we used the Johansen cointegrated rank test to determine the number of long-run cointegrated relationships among the variables included.

The economic theory underlying an AIDS model of n expenditure share equations predicts the existence of exactly (n-1) long-run (cointegrated) relationships among its variables. In the AIDS system of the UK tourism demand there are three expenditure share equations and, hence, two cointegrated vectors should be accounted for. The Johansen cointegrated rank test provided statistical evidence to support the theoretical predictions.

The theoretical framework of an AIDS model of the UK tourism demand also establishes the structural form of the cointegrated relationships it predicts. This form is that of the equations included in the system which describe the destinations' tourism shares as functions of own and competing tourism prices and the UK real per capita tourism budget.

Therefore, the structural parameters of the cointegrated vectors in the corresponding VAR should be exactly-identified with restrictions matching those of the normalisation process used to identify the tourism share equations of the static AIDS system.

The resulting structural form of the cointegrated VAR, identical to that of the AIDS specification, was then subjected to additional over-identifying restrictions, such as homogeneity, symmetry and null cross-price effects between the share equations of Portugal and France. These hypotheses were tested for the parameters of the cointegrated VAR and could not be rejected. Consequently, further evidence was obtained about the capability of the AIDS model to comply with theoretical predictions underlying the rationality of the UK tourism demand behaviour and the destinations' competitive conduct.

The estimates of the structural coefficients of the cointegrated VAR under the full set of theoretical restrictions were then used for computing the uncompensated expenditure, own- and cross-price elasticities of the tourism shares. The results obtained proved to be similar to the corresponding elasticity values obtained from the AIDS model estimated in chapter 5. Hence, the economic interpretation of these results is identical to that provided in chapter 5, namely, that the estimated magnitudes of the long-run expenditure elasticities approximate unity; that all own-price elasticities estimates have the expected signs and plausible magnitudes, with the UK demand for Portugal being more sensitive to own-price changes than the demand for France or Spain is to own-price changes in these destinations; the cross-price elasticities estimates confirm Portugal and Spain on the one hand, and France and Spain on the other hand, as competing destinations, as increasing own-prices in one destination lead to increasing shares of the other, while own-price changes in Portugal (France) do not affect significantly the tourism share for France (Portugal).

Given the theoretical and empirical consistency of these results, the predictive ability of the cointegrated VAR model under the full set of restrictions should not come as a surprise. Indeed, the quality criteria measuring the accuracy of this model's forecasts indicate it as the most precise forecasting device among all VAR specifications used in this chapter to predict the UK demand for tourism. However, these same quality criteria suggest the general reduced-form (unrestricted) VAR to be the next best forecasting model. This gives empirical support to the claimed competence of the VAR approach for forecasting purposes, and emphasizes the valuable qualities of modelling simplicity and estimation ease of the "pure" unrestricted VAR. Thus, it seems reasonable to conclude that, if the main goal of a multivariate modelling exercise is to obtain reliable forecasts, an appropriately specified reduced-form VAR is likely to suffice.

The main goal of this chapter was to investigate, empirically, for validating the theoretical principles underlying the AIDS approach and the estimation and inference procedures implemented with the AIDS model in previous chapters. We did so using a VAR model framework and the Johansen cointegrated vector methodology. We believe we have provided enough empirical evidence for considering the AIDS approach as a theoretically consistent and statistically robust means to produce valid and reliable estimates of the long-run equilibrium parameters underlying the relationships between destinations' tourism shares, tourism prices and an origin's tourism real per capita expenditure. In addition, we confirmed the competence of the VAR model, either in its more general unrestricted form or under the full set of theoretical restrictions, for providing accurate forecasts of the destinations' tourism shares.

A good econometric model offers a formal quantitative framework that can be useful for understanding economic activities, interpreting economic relationships, testing economic theories, evaluating economic policy and predicting economic behaviour. Hence, the predictive power of an econometric model is an integral part of its quality judgement. However, the forecasting performance of the ARDL models estimated in chapter 4 and the AIDS systems estimated in chapters 5 and 6 have not yet been assessed. This task is carried out in the next chapter.

CHAPTER 8

THE FORECASTING ABILITY OF ALTERNATIVE ECONOMETRIC MODELS OF THE UK DEMAND FOR TOURISM

8.1 INTRODUCTION

A fundamental objective of economic research is to describe, explain and predict the complexities involved in economic behaviour. Economic phenomena are complex because they are interrelated and contain dynamic components. Shocks and policy changes may take time before their influences are fully transmitted throughout the economy. Moreover, economic variables affect, and are affected by other variables. The dynamic and interrelated nature of most economic variables implies that technological, social and political changes impact on their levels not only in the current period but also in subsequent periods.

The nature of economic behaviour may be captured by econometric models, the structure of which describes the intricacies of the relationships between the current value of a dependent variable and the current and/or lagged values of several explanatory variables, as well as lagged values of the dependent variable itself. The use of appropriate methodologies in the construction, estimation and evaluation of such models can provide accurate and reliable means of explaining and predicting the economic behaviour under investigation. However, if deficient methodologies are used in any of these three basic steps of empirical modelling, the resulting estimates may be misleading, the inference procedures invalid and the forecasting ability mediocre.

In previous chapters we constructed, estimated and evaluated alternative econometric specifications for modelling the UK demand for tourism in France, Spain and Portugal. The statistical evaluation of these models involved several quality criteria which are usually applied for qualifying empirical models as accurate, reliable or useful representations of the reality they supposedly portray. According to these criteria, we found that the ARDL and ECM models estimated in chapter 4, the AIDS models specified in chapters 5 and 6 and the VAR models constructed in chapter 7 were theoretically coherent, empirically plausible and statistically robust representations of the economic phenomena under investigation. In addition, we showed that the different models could give different insights and emphasize different perspectives of the economic behaviour analysed. Therefore, they can be useful forms of explaining the UK demand for tourism in France, Spain and Portugal under the specific purposes for which they were intended. However, the forecasting ability of econometric models is a fundamental aspect of their quality evaluation and, except for the VAR models of the previous chapter, this aspect has not yet been addressed. Given its major importance, we now undertake the analysis of the forecasting ability of the models specified previously.

Forecasting is a crucial input of the decision-making process. Rather than an isolated activity, it should be considered an integral part of this process. In this sense, this activity is not only essential but also an unavoidable task of a comprehensive economic analysis. Forecasting problems are as diverse as the behavioural features of the variables they intend to address. This diversity is such that it seems unreasonable to expect a unique best method to tackle all prediction issues. Indeed, a considerable number of methodological tools is presently available for solving the variety of forecasting problems arising from the inherent complexities of modern economics. The wide range of different problems met in practice justifies a wide range of alternative forecasting approaches, implying that there is no single right way to address them all. In fact, every variable forecasting problem has its unique features, demanding careful analysis of its past and present behaviour and of the main interrelated factors that might influence it.

The problem of how to predict the future, given information on the past, takes us into the subject matter of time series analysis. Understanding the past is a sound base for looking more knowledgeably into the future. However, even after careful historical study of the problem at hand, it will often be the case that more than one appropriate solution emerges. Although there is no single right forecasting approach to time series analysis, some might be considered more appropriate than others in specific contexts, and it might even be the case that a combination of different methods serves as the best solution for a particular problem in sight. The identification of what might be a reasonable methodological approach requires consideration of several factors. These include what purposes forecasts are needed for, what quantities require prediction, what forecasting horizon is relevant, what resources are available and what amount of effort is worthwhile in view of the potential return from increased forecast accuracy. Having established what must be predicted, why and how far, the next step involves the consideration of which might be the best methodological approach for obtaining accurate forecasts, given the resources and information available.

This chapter concerns these and other aspects of the process of predicting the UK tourism shares of France Spain and Portugal. The chapter is structured as follows. Section 8.2, briefly addresses the qualities and limitations of some quantitative forecasting methods, and gives an overview of recent research work on tourism demand forecasting. Section 8.3, presents the forecast results obtained using the econometric models specified in chapters 4 to 7. Section 8.4 compares the forecasting ability of these different econometric models and section 8.5 concludes.

8.2. MERITS AND LIMITATIONS OF QUANTITATIVE FORECASTING METHODS

The forecasting methodological apparatus presently available includes approaches that, as Clements and Hendry (1998) point out, range from pure guessing based on luck, to the building of complex large-scale macroeconometric systems or, as Newbold and Bos (1994, p. 13) state "*from the very informal to the formal, from the use of unaided human judgement to the exclusive reliance on quantitative models, from the very simple to the highly complex*". In spite of this wide range of methodological tools, the forecasting methods available can be divided into two main groups: quantitative and non-quantitative methods. While agreeing with Newbold and Bos (1994) in that most successful forecasting exercises involve a combination of formal quantitative methods and less formal judgemental expertise, we believe that the former are, in general, more appropriate in forecasting problems involving economic time series, and more useful in the particular forecasting exercise we are interested in addressing in this chapter.

A number of forecasting methods based on quantitative approaches have been proposed and implemented in the literature.[1] Two of these have been given particular

[1] See, for example, Fair (1986) for a partial survey, Wallis (1989) for an appraisal of some important contributions of the 1970s and 1980s and Harvey (1993) for a study of more recent forecasting methods.

attention: the univariate time series forecasting approach and the econometric regression analysis approach. The univariate approach, in its basic variants, involves the prediction of future values of a time series process exclusively on the basis of its own past values. The econometric approach involves the specification of a statistical model designed to isolate and describe relationships among time series processes, which are interpretable by economic theory and might be used for forecasting purposes.

There is a similarity between these two forecasting procedures springing from the construction method of the models on which they are based. Assuming that the objective is to obtain reliable forecasts of the future values of an economic time series, these forecasting procedures are based on a statistical model-building approach which specifies, as accurately as possible, the assumed underlying data generating process. This statistical approach provides a framework in which the appropriate forecast function is approximated through the analysis of observed data so that the prediction method is tailored to the specific characteristics of a given time series process. Moreover, the statistical models on which univariate time series and econometric forecasting procedures are based allow for the derivation of measures of forecast uncertainty and associated tests of forecast accuracy, not available for *ad hoc* procedures. However, the univariate and econometric methodologies differ substantially in several other aspects well worth mentioning.

A crucial difference between these methodologies resides primarily in the objectives for which they are implemented. While univariate time series methods exclusively serve forecasting purposes, formal multivariate econometric models fulfil much more useful roles than being simple forecast-generating devices. As Clements and Hendry (1998, p.16) point out, econometric models "*consolidate existing empirical and theoretical knowledge of how economies function, provide a framework for progressive research strategy and help to explain their own failures. They are open to adversary scrutiny, are replicable and, in this sense, offer a scientific basis to research*". A second group of differences between univariate and econometric forecasting models can be identified with the theoretical basis on which they rest and with the degree of complexity of their statistical specifications.

There is a wide range of time series forecasting procedures exclusively based on its own past values. This type of models, often qualified as naive, is implemented through the estimation of very simple statistical models. Examples of such models are no-change, previous-change, pure autoregressive (AR), stepwise AR and autoregressive integrated moving average (ARIMA) methods, among others.[2] The ARIMA class of models, usually referred as Box-Jenkins (1970) approach, is particularly useful for short-term economic forecasting and has been extensively applied for this purpose in time series literature. These models are generally viewed as quite sophisticated when compared with other univariate models. Yet, viewed in a different way, this approach can be considered as naïve as other univariate forecasting processes. Indeed, the construction of univariate time series models completely ignores whatever relevant knowledge, expertise, information or theory may be available at the time. It is based exclusively on the evidence provided by the past values of the variable under consideration, taking no account of the broader context in which the time series is generated. In this sense, these are purely data-analysis statistical models and differ from the formal econometric specifications in that *"they do not begin with a conceptual framework provided by economic theory, that specifies a relationship between economic variables. Thus, behavioural or technical equations... are not considered. Instead, the emphasis is on making use of information from past values of a variable for forecasting its future values by using what amounts to a sophisticated extrapolation procedure."* (Griffiths *et al.* 1993, p. 639). Nevertheless, in circumstances involving a cost-benefit perspective and short-run purposes, univariate methods can be useful processes to address specific forecasting problems. Past observations yield important information about the future and, frequently, no additional relevant information is available. Furthermore, these methods are relatively easy to implement and often provide forecasts of sufficient quality for the purposes at hand.

However, ignoring relevant available information cannot always be thought of as a good strategy when engaging in forecasting. Reliable forecasting procedures have to be based on all currently available information. Forecasts of a time series can often be improved considerably through the incorporation of information on related time series. In addition, the inclusion of what is known from economic theory into the model-building process is of paramount importance. Such theory is likely to postulate behavioural relationships among related time series which cannot be ignored without seriously affecting the accuracy of the forecasting process. In fact, economic theory is essential in the model-building and model-selection processes. Theory can suggest a particular model, a group of relevant variables, a suitable functional form or lag structure. Moreover, theory can play a part in the quality evaluation of the models selected, through the plausibility of the signs

[2] For an introductory description of several of these methods see, for example, Newbold and Bos (1994).

and magnitudes of the coefficients' estimates and the outcome of hypothesis testing. These considerations indicate that, in cases for which the behaviour of a time series variable is likely to depend on the behaviour of other time series and economic theory plays a major role in explaining the nature of their interrelated behaviour, it is advisable to move away from a-theoretic naive univariate models and embrace the theory-based, multivariate econometric specifications for forecasting purposes.

At this point, however, a cautionary note should be issued. The building of behavioural, high-powered theory, large-scale, complex econometric models for forecasting purposes is not always advisable. The large quantities of information needed, the enormous amount of time and effort spent on their construction and the heavy computational requirements of their estimation and testing, can make them appear impractical and useless in many situations. Frequently, bigger is not better and successful forecasting models are likely to be parsimonious, involving just a small number of unknown parameters.

This brings us to the matter of comparative forecasting performance. How accurate are forecasts, how should this accuracy be evaluated and how to compare differently sourced forecasts are important questions that many empirical studies have been addressing over the last decades. Several of these studies assess the quality of a set of forecasts obtained from econometric models through the comparison of accuracy with a set of competing forecasts obtained from univariate models. Frequently, the univariate models chosen for comparison are ARIMA models and, more recently, their multivariate generalisations, VAR models. Although, historically, the univariate approach has been claimed to perform well relative to econometric methods and, as noted in Charemza and Deadman (1997), simple univariate time series methods seem to provide more accurate forecasts than the large and complex models constructed under the traditional methodology, the available evidence is not conclusive. While the results of several studies, for example Cooper (1972), underline the superior performance of univariate forecasting methods, others, for example Zarnowitz (1978), indicate that structural econometric models out-perform ARIMA models by quite substantial margins, particularly for extended forecast horizons. While in Granger (1981), the irrelevance of the univariate forecasting approach for policy purposes is pointed out, it is also recognized that the same remark can be made about poorly specified econometric models and that "*a badly specified econometric model is even more dangerous because it may appear to have something to say about alternative policies and yet actually may be very misleading*" (p. 127).

In studies which address the issue of comparative forecasting performance, for example, Clements and Hendry (1998), it has been acknowledged that the forecasting success of univariate models relative to econometric approaches is likely to be due to dynamic misspecifications in the econometric models and/or to the failure of adequately accounting for the data being used. Perhaps because of the disappointing results of early evaluations, econometricians have been paying extra attention to the dynamic specifications of their models, and it seems not a coincidence that recent research on the subject of comparative forecasting performance indicates a more favourable position towards the forecasting ability of well-specified econometric models when compared with that of naive univariate models.[3] However, as Newbold and Bos (1994) stress, given the difference in scale and effort involved, it is difficult to be overwhelmingly impressed with these findings and scope for further improvement in the quality of econometric forecasts still remains.

In the specific context of recent tourism demand analysis, several empirical studies address the matter of comparative forecasting accuracy based on alternative forecasting methodologies. Many of these studies only emphasise univariate methods, others focus only on econometric approaches and few include the comparative forecasting performance of econometric versus univariate procedures.[4] The results of most studies comparing the forecasting accuracy of econometric and simple univariate methods seem systematically to favour naive and *ad hoc* forecasting techniques (such as the no-change model) relative to what are claimed to be "more complex" forecasting procedures (such as the ARIMA method) and what is usually qualified as the "regression" or "econometric" approach. However, in several cases, the "regression" models used for comparison are not what could be viewed as structured multivariate econometric causal specifications, but rather some simplified version of single equation models, with no account for integrated data, dynamics and/or for the theoretically established full rank of explanatory variables.

Some studies, although stating their intention of addressing alternative tourism demand forecasting methods, do not explicitly do so. For example, Fujii and Mak (1981) analyse alternative estimation techniques for forecasting purposes, rather than alternative forecasting models. In their study, the alternative estimation techniques (ridge regression,

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[3] For an enlightening debate on the forecasting accuracy of large-scale macroeconometric models see Kmenta and Ramsey (1981). For a survey in the comparison of time series and econometric forecasts see Granger and Newbold (1986). For a review of econometric forecasting methods see Granger (1990).

[4] For surveys in tourism demand forecasting see, for example, Archer (1976), Uysal and Crompton (1985) and Witt and Witt (1995).

GLS and OLS) are applied to a single equation model which regresses the number of US visitors to Hawaii on its lagged values, real airfare transportation cost and real per capita US income. The study concludes that the ridge regression method out-performs the GLS and OLS methods for forecasting purposes. Other examples are Martin and Witt (1987) and Lee, Var and Blaine (1996). The former study, although under the title "*Tourism demand forecasting models*", focuses on the suitability of alternative measures of the explanatory variable 'cost of tourism' and the latter, although underlining the importance of forecasting in tourism demand analysis, focuses on specification aspects and estimation methods for tourism demand regressions, rather than engaging in forecasting procedures for predicting the future values of the dependent variable.

An example of a comparative analysis of alternative forecasting methods is Gonzalez and Moral's (1996) study of forecasting trends of international tourism demand for Spain. This study analyses the forecasting performance of a Basic Structural Model (BSM) which is supposed to capture the trend, seasonal, cyclical and irregular components of the dependent variable time series, a Box-Jenkins ARIMA model and a "classic linear regression model" which is derived from the BSM model by imposing some restrictions on its parameters. The explanatory variables of this 'econometric' regression are a linear deterministic trend and a seasonal component. The forecasting performance of the BSM model is found to be superior when compared with that of the ARIMA and the "deterministic regression" models.

In other studies, for example, an early work by Geurts and Ibrahim (1975) and a more recent one by Witt *et al.* (1994), econometric specifications are not included for comparison purposes of forecasting accuracy. In the former, the Box-Jenkins and the exponential smoothing approaches are applied to forecast the number of incoming tourists to Hawaii. In the latter, besides these two forecasting methods, five other univariate approaches (no-change, constant-change, trend, Gompertz and stepwise autoregressive) are used for comparative purposes in predicting the UK outbound tourism flows.

Several research studies of tourism demand forecasting do not address forecasting methods other than an econometric approach. For instance, a group of papers on tourism demand forecasting, using solely an econometric approach to predict long-term international tourism flows, includes Smeral *et al.* (1992), Smeral and Witt (1996) and Smeral and Weber's (2000) studies. In the study by Smeral *et al.* (1992), a system of tourism demand equations is estimated over the period 1975-1988 and used to generate econometric forecasts of tourism imports and exports flows for 18 industrialised countries, up to the year 2000. Smeral and Witt's (1996) study modifies the previous model to allow for destination-specific demand structures in the export functions, extends the estimation period by four years up to 1992 and the forecast horizon up to the year 2005, and considers two different income growth scenarios. The Smeral and Weber's (2000) study contemplates a similar model for 20 countries estimated over the period 1975-1996, and extends the forecasting horizon up to 2010 taking into consideration, for the alternative scenarios adopted, the countries' participation in the European Monetary Union.

Examples of studies explicitly comparing the forecasting performance of econometric models with that of univariate naive models include Martin and Witt (1989), Witt and Witt (1992), Kulendran (1996) and Kulendran and King (1997) Kim and Song (1998), Song *et al.* (1999), Song *et al.* (2000) and Song and Witt (2000). While Martin and Witt (1989), and Song *et al.* (1999) suggest that the forecasting ability of naive no-change models tends to over-perform other type of forecasting techniques including ARIMA and econometric approaches, Witt and Witt (1992), Kulendran (1996) and Kulendram and King (1997) conclude that ARIMA models can be powerful competitors with econometric specifications. Kim and Song (1998) and Song *et al.* (2000) contradict these findings showing that error-correction econometric models are able to produce better forecasts than univariate time series models. Using two quality criteria (MAE and RMSE), Song and Witt (2000) compare the performance of several models – Engle-Granger ECM, unrestricted and cointegrated VAR, ARDL, time varying parameter (TVP) and no-change (NC) – in forecasting the demand for tourism of two origin countries (UK and USA). Depending on the forecasting horizon, quality criteria and the origin considered, the ranking of the best forecasting model varies considerably. For instance, while the one-year-ahead forecasts for the UK show that the no-change model is ranked first using MAE and the TVP model is ranked first using RMSE, for the USA using both criteria, the TVP is first and the no-change model is third. In two-years-ahead forecasting and using both criteria, the Johansen cointegrated VAR is best for the UK and the no change model is best for the USA. The overall forecasting performance of the TVP is ranked first for the USA while, for the UK, it seems to depend on the quality criteria used.

Given these rather confusing results, research on comparative forecasting issues has still a long way to go. Nevertheless, early evidence produced on the poor forecasting performance of econometric models may be attributed more to misspecifications of the models used, rather than to some intrinsic defect of the econometric approach. Furthermore, in the specific research matter we are concerned with, it appears that more than simple *ad hoc* no-change models or any other univariate extrapolation procedure, however sophisticated, is needed to obtain accurate forecasts of the UK tourism demand for France, Spain and Portugal.

Indeed, based on the estimation results obtained from the econometric models specified in previous chapters, we showed that the behaviour of the dependent variable representing the UK demand for tourism is structurally linked to the behaviour of several explanatory variables, implying that changes in these regressors have a separate and additional, significant impact on the dependent variable. Moreover, for all the econometric specifications used previously, we provided empirical evidence of their theoretical consistency and statistical robustness, implying that these specifications are valid means of describing, interpreting and forecasting the UK tourism demand for France, Spain and Portugal. Hence, we believe that a univariate model, basing the explanation of the dependent variable's behaviour solely on its lagged values (with or without lagged disturbances), is a mis-specified representation of the underlying data generating process. The consequences of mis-specification in statistical models are well known: the estimation and statistical inference procedures are invalid. As a result, we consider that univariate time-series models are not appropriate means of predicting the behaviour of the UK tourism demand in France, Spain or Portugal. Hence, comparisons between the forecasting ability of these naive models with that of the theoretically consistent, data-coherent, well-specified and statistically robust econometric models of previous chapters seem to be either a trivial "confirm-the-obvious" task, or a pointless exercise of self-indulgence.

In what follows we concentrate on the derivation, quality evaluation and comparison of forecasts of UK tourism demand in France, Spain and Portugal obtained from the ARDL models of chapter 4, AIDS models of chapters 5 and 6, and VAR models of chapter 7.

8.3. THE FORECASTING PERFORMANCE OF ECONOMETRIC MODELS OF THE UK DEMAND FOR TOURISM IN FRANCE, SPAIN AND PORTUGAL

In order to explain how the forecasts of the UK tourism demand in France, Spain and Portugal are derived, evaluated and compared, it is important to identify which time series is to be forecast and to clarify the terminology used to describe the basic concepts and methods used in forecasting procedures. These include the definition of the in-sample estimation period and out-of-sample forecasting period, the forecasting horizon, the forecasting equations valid in each case, the type of forecast (interval/point, ex-post/ex-ante, conditional/unconditional, static/dynamic), and the statistical criteria used for assessing the quality of the forecasts.

8.3.1. Forecasting Procedures: Basic Concepts

Econometric forecasts are quantitative estimates of a variable's future values, obtained from past and current information included in the specification of an econometric model. By projecting this model beyond its estimation period, we derive forecasts for the future values of the variable under analysis. Two types of forecasts can be obtained by this process: point-forecasts which predict a single forecast value, and interval-forecasts which define a confidence interval used to provide a margin of error around the point-forecast.

The time series variable for which we are interested in obtaining econometric forecasts is the level of the UK demand for tourism in France, Spain and Portugal. The derivation of these forecasts is based on the ARDL models specified in chapter 4, the AIDS models specified in chapters 5 and 6 and the VAR models specified in chapter 7. For this forecasting exercise, the existing sample of 29 observations over the period 1969-1997 is divided into two sub-periods: the in-sample estimation period comprising 25 observations from 1969 to 1993, and the out-of-sample forecasting period involving the last four observations available. For both these sub-periods, the actual values of all variables involved in the estimation and forecasting equations are known. However, if the forecasting period had to be extended beyond 1997, no information about the actual values of these variables would be available.[5] This fact allows us to distinguish between ex-post and ex-ante and unconditional and conditional forecasting procedures.

Both ex-post and ex-ante procedures predict values of the dependent variable beyond the estimation period of the model. However, while in the former the actual values of both the dependent and the explanatory variables are known with certainty, in the latter the values of the variables may or may not be known with certainty. Therefore, an ex-ante forecasting procedure permits the checking of the forecasts against existing data and, hence, provides an immediate means of evaluating the accuracy of the forecasts.

[5] This is true with one exception. In the case of the AIDS models of chapters 5, the future values of the trend variable on the right-hand side of all equations are previously known.

In an unconditional forecasting procedure, the values of all the explanatory variables in the forecasting equation are known. Hence, any ex-post forecast is also an unconditional forecast. However, an ex-ante forecast may also be unconditional if, for example, the explanatory variables include trends or lagged values. In a conditional forecasting procedure, values of one or more explanatory variables are not known. So they must be previously forecast to be used in the dependent variable forecast equation.[6] This introduces further uncertainty into the forecasts of the dependent variable which, for this particular forecasting exercise, we want to avoid. Therefore, to obtain point-forecasts of the UK tourism demand levels from the econometric models previously specified, we use ex-post unconditional forecasting procedures over the out-of-sample period 1994-1997.

8.3.2. Evaluation Criteria of Forecasts

In any forecasting evaluation process a fundamental measure of accuracy on which most evaluation criteria are based, is the difference between the actual value of the time series and its corresponding forecasted value. This difference is known as forecast error and the best forecast is defined as "*the one which yields the forecast error with minimum variance*" (Pindyck and Rubinfeld, 1998, p.204). Therefore, once an econometric model have been estimated and used to predict the dependent variable, one important statistic for evaluating the forecast precision is the forecast error variance and its associated confidence interval $100(1-\alpha)\%$.

Assume that the time series variable to be predicted is Y_t, where $t=1,\ldots,T$, and that the actual values of this variable and all its determinants are known for all t. Assume further that an econometric model has been used to regress Y_t on its determinants using a sample of N out of the T available observations ($N<T$), leaving the remaining n observations for forecasting purposes ($N + n = T$). We define the h-step ahead forecast for this variable as Y^f_{N+h}, $h=1,\ldots, n$, and the actual value of this variable in the same period as Y_{N+h}.[7] Hence, the forecast error corresponding to this ahead period is.

[6] These concepts follow those of Pindyck and Rubinfeld (1998). Yet, in other textbooks, (Newbold and Bos, 1994, for example), conditional and unconditional forecasts are defined in the opposite way; that is, in conditional forecasting the values of all variables are known and in unconditional forecasting the values of the explanatory variables for the forecasting period have to be forecast.

[7] If a regression contains the lagged dependent variable on its right-hand side, the forecast of this variable, updated period by period, is used to create what is generally known as a *dynamic* forecast. If no lags of the dependent variable are present among the regressors, the forecasting process generates *static* forecasts.

$$e_{N+h} = Y_{N+h} - Y^f_{N+h}. \tag{8.1}$$

and the associated $100(1\text{-}\alpha)\%$ confidence interval is

$$[Y^f_{N+h} - t_{\alpha/2}.sd(e_{N+h})\ ;\ Y^f_{N+h} + t_{\alpha/2}.sd(e_{N+h})] \tag{8.2}$$

where $t_{\alpha/2}$ is the critical value of the t statistic for the appropriate degrees of freedom and significance level, and $sd(e_{N+h})$ is the estimated standard deviation of the forecast error. If the actual value of the forecasted variable lies within this interval, the forecast for this actual value is considered to be accurate at the $\alpha\%$ significance level.

In addition or alternatively, three other measures of average forecast quality are frequently used in the literature: the mean absolute prediction errors (MAE), the mean squared prediction errors (MSE) and the root mean squared prediction errors (RMS). These measures are defined as follows. If n forecast errors e_t are available, the mean absolute prediction errors is

$$MAE = \sum_{t=1}^{n} \frac{|e_t|}{n} \tag{8.3}$$

the mean squared prediction errors is

$$MSE = \sum_{t=1}^{n} \frac{e_t^2}{n} \tag{8.4}$$

and the root mean squared prediction errors is

$$RMSE = \sqrt{\sum_{t=1}^{n} \frac{e_t^2}{n}} \tag{8.5}$$

These measures are all useful statistics for assessing the average quality of a set of forecasts and, provided that competing sets are available, they can be compared. The set of forecasts for which the criterion statistic is smaller will be judged to be better.

In addition to these generally used measures, the predictive failure test can also be a valuable means of complementing the assessment of a forecasting method's accuracy. For the relevant time series, this test evaluates the similarity between observations pertaining to the in-sample estimation period, and those pertaining to the out-of-sample forecasting period. The null hypothesis states that the out-of sample observations used in the forecasting period do not have outlier characteristics when compared with the set of observations of the in-sample estimation period and, statistically, these two sets of observations can be considered to pertain to the same homogeneous sample. If the null is not rejected, both sets can be said to have been withdrawn from the same population. As shown in Pesaran *et al.* (1985), the predictive failure test can also be used as a general specification error test. Under the classical assumptions, this test statistic has an exact F-distribution. Hence, if the observed value of the statistic lies below the relevant critical F value, the null cannot be rejected.[8]

When comparing different sets of forecasts, it is important to assure that the comparative exercise is based on homogeneous sets of information. Hence, it is essential to guarantee that the same time series variable is being forecast and the same prediction horizon is being considered to compare forecasts generated by different methods.

8.3.3. THE FORECASTING PERFORMANCE OF ALTERNATIVE ECONOMETRIC MODELS [9]

8.3.3.1. ARDL models

In chapter 4 we specified single equation dynamic ARDL models of the UK demand for tourism in France, Spain and Portugal. The estimation results obtained with these models using all the observations available are presented in Tables 4.2 (France), 4.4 (Spain) and 4.6 (Portugal). The level of the UK tourism demand for France, Spain and Portugal are the time series variables we are interested in forecasting and these variables are represented in the ARDL models by the log of the UK real per capita expenditure allocated to each destination. Forecasts of these variables are obtained by estimating the ARDL models for the in-sample period 1969-1993, leaving the last four observations for forecasting purposes. Tables 8.1 to 8.3 report the actual and forecasted values of the dependent variable (real per capita expenditure in each destination) and a number of summary statistics for evaluating the accuracy of forecasts accuracy.[10] Table 8.1 presents the forecasting results for France, table 8.2 for Spain and table 8.3 for Portugal.

[8] This test is easily performed on an equation by equation basis when the estimation method is OLS. However, when a system of equations is considered and another estimation method is required, the value of this test statistic is not easy to compute and its statistical distribution is not an exact F-distribution. Therefore, this test is performed only when the models are estimated with OLS on an equation by equation basis.

[9] The forecasts and summary statistics for all the models analysed in this section were computed using Pesaran and Pesaran (1997) Microfit 4.0. If lagged values of the dependent variable are explicitly included as regressors, as in the cases of the ARDL, dynamic AIDS and VAR models, Microfit automatically computes *dynamic* forecasts, otherwise it generates *static* forecasts.

[10] As the ARDL specifications are single equation models which can be estimated using OLS estimation method, the predictive failure test is reported for these models.

Table 8.1: Forecasting results for the log of the UK real per capita expenditure in France

	1994	1995	1996	1997
Actual values	**1.4848**	**1.5080**	**1.4785**	**1.5215**
Forecast	1.4813	1.4966	1.4960	1.5450
Forecast error	0.00357	0.01138	-0.01746	-0.02357
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE	0.015081		0.013995	
MSE	0.0003083		0.0002506	
RMSE	0.017559		0.015831	
Predictive Failure Test: $F(4, 15) = 0.28717(0.882)$				

Table 8.2: Forecasting results for the log of the UK real per capita expenditure in Spain

	1994	1995	1996	1997
Actual values	**1.5515**	**1.5817**	**1.5383**	**1.5477**
Forecasts	1.5092	1.5583	1.5478	1.5597
Forecast error	0.042350	0.023414	-0.0095275	-0.011944
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE	0.024182		0.021809	
MSE	0.0008754		0.0006438	
RMSE	0.029588		0.025373	
Predictive Failure Test: $F(4, 16) = 0.403(0.804)$				

Table 8.3: Forecasting results for the log of the UK real per capita expenditure in Portugal

	1994	1995	1996	1997
Actual values	**0.71954**	**0.76522**	**0.72156**	**0.74808**
Forecasts	0.73205	0.73337	0.71755	0.68552
Forecast error	-0.012517	0.031846	0.004001	0.062560
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE	0.033324		0.027732	
MSE	0.0017025		0.0012752	
RMSE	0.041262		0.035709	
Predictive Failure Test: $F(4, 17) = 0.30196(0.873)$				

8.3.3.2. Static AIDS models

In chapter 5 we specified a static AIDS model for the UK tourism demand using as dependent variable the expenditure share of each destination. We estimated this model unrestricted, under homogeneity and symmetry restrictions [denoted by (H+S)] and under the additional restriction of null cross-price effects between France and Portugal [denoted by $(H+S)^0$]. The estimation results are reported in table 5.1 and show that this additional restriction only improves the precision of the estimates and does not have a significant effect on the magnitudes of the estimated coefficients. Hence, the forecasts obtained with the (H+S) model are not expected to diverge much from the forecasts obtained with the $(H+S)^0$ model.[11] In contrast, the forecasting ability of the unrestricted model is expected to diverge significantly from that of the restricted (H+S) model. Therefore, for comparison purposes, the interesting models to use are the unrestricted model and the (H+S) model.

The forecast ability of these models is compared by estimating the unrestricted (denoted by *) and the restricted (denoted by **) versions for the period 1969-1993, leaving the last four observations for forecasting. Tables 8.4, 8.5 and 8.6 report the actual and forecasts values of the shares for France, Spain and Portugal, respectively, and the same summary statistics used previously for evaluating the accuracy of those forecasts.[12]

[11] Both these models were used to forecast the relevant variable and the results obtained show that the difference between the two sets of forecasts is minimal, never exceeding 0.0003.

[12] The unrestricted AIDS can be estimated using either SURE or OLS on an equation by equation basis, as the two estimation methods are equivalent in this case. However, for the restricted version including cross-equations constraints, the SURE method is applicable. Therefore, we report the results of the predictive failure test only for the unrestricted version.

Table 8.4: Forecasting results for the UK expenditure share of France

	1994	1995	1996	1997
Actual values	**0.39700**	**0.38540**	**0.38967**	**0.40481**
Forecasts**	0.37775	0.38044	0.40188	0.38832
Forecast error**	0.019251	0.004959	-0.012202	0.016484
Forecasts*	0.40027	0.39706	0.42707	0.46015
Forecast error*	-0.003272	-0.011659	-0.037393	-0.055338
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE**	0.018839		0.013224	
MSE**	0.000548		0.000204	
RMSE**	0.023407		0.014281	
MAE*	0.016122		0.026916	
MSE*	0.000450		0.001152	
RMSE*	0.021212		0.033938	
Predictive Failure Test*: F(4, 17) = 0.423(0.79)				

Table 8.5: Forecasting results for the UK expenditure share of Spain

	1994	1995	1996	1997
Actual values	**0.51857**	**0.52625**	**0.52292**	**0.50691**
Forecasts**	0.52613	0.52235	0.49786	0.51260
Forecast error**	-0.007558	0.003893	0.025054	-0.005696
Forecasts*	0.50068	0.50406	0.46910	0.42879
Forecast error*	0.017892	0.022187	0.053812	0.078121
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE**	0.021050		0.010550	
MSE**	0.000637		0.000183	
RMSE**	0.025236		0.013532	
MAE*	0.016681		0.043003	
MSE*	0.000500		0.002453	
RMSE*	0.022362		0.049525	
Predictive Failure Test*: F(4, 17) = 0.593(0.673)				

Table 8.6: Forecasting results for the UK expenditure share of Portugal

	1994	1995	1996	1997
Actual values	**0.084433**	**0.088348**	**0.087411**	**0.088283**
Forecast**	0.096126	0.097201	0.10026	0.099071
Forecast error**	-0.011693	-0.008853	-0.012852	-0.010788
Forecasts*	0.099053	0.098877	0.10383	0.11107
Forecast error*	-0.014620	-0.010529	-0.016420	-0.022785
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE**	0.006626		0.011047	
MSE**	0.000056		0.000124	
RMSE**	0.007508		0.011143	
MAE*	0.006259		0.016088	
MSE*	0.000052		0.000278	
RMSE*	0.007277		0.016684	
Predictive Failure Test*: F(4, 17) = 0.469(0.758)				

8.3.3.3. Dynamic AIDS models

In chapter 6 we estimated a dynamic AIDS model for the UK tourism demand in France, Spain and Portugal unrestricted, under the restrictions of homogeneity and symmetry and null cross-price effects between France and Portugal imposed only on the long-run [denoted as $(H+S)^0L$], and under this same set of restrictions imposed on both the long- and the short-run coefficients [denoted as $(H+S)^0LS$]. We presented the estimation results in table 6.3. For comparison of the elasticities estimates obtained with the dynamic AIDS with those obtained with the AIDS model specified in chapter 5, we used the dynamic model with restrictions imposed only on the long-run coefficients. However, the version of the dynamic AIDS for which the theoretical restrictions are imposed both on the long- and short-run, may be a better representation of the behavioural characteristics of UK tourists than the version for which these restrictions are imposed only on the long-run. Hence, we believe that the former version is able to forecast better than the latter. Hence, the dynamic AIDS model with restrictions imposed only on the long-run and the version imposing the same restrictions both on the long- and short-run are estimated for the period

1969-1993, leaving the four last observations for forecasting purposes and a comparison between the forecasting performances of these two versions will complement the analysis of the predictive ability of this class of models.

However, if comparison of forecasting performances is to be carried out across all the models specified in previous chapters, we must assure that the time series variables which are being predicted are the same. The time series variables we are interested in predicting are the levels of UK tourism shares of France, Spain and Portugal. Yet, in the dynamic AIDS model specified in chapter 6 the dependent variable, representing the destinations' shares, is defined in first differences instead of levels. A simple algebraic manipulation transforms the original first difference dependent variable ($\Delta W_i = W_{it} - W_{it-1}$) in its level form ($W_{it}$), which is the variable of interest for forecasting purposes. This manipulation consists of transferring the variable W_{it-1} from the left-hand side of the equation to its right-hand side. Except for the coefficient of the explanatory variable W_{it-1}, which is now ($1+\lambda$) instead of the original λ, all the regression coefficients remain exactly the same. This shows that the transformation operated is innocuous in terms of the model's estimation results. Therefore, we use this transformed model for obtaining forecasts of the UK tourism demand levels based on the dynamic AIDS model.

Tables 8.7 to 8.9 report the actual and forecasted values of each destination expenditure share level (Wi_t) obtained with the dynamic AIDS model, and the usual summary statistics. The forecast results obtained with the model under the full set of restrictions imposed solely on the long-run coefficients are denoted by •; the forecast results obtained with the model under the full set of restrictions imposed on both the long- and the short-run coefficients are denoted by ••. Table 8.7 presents the forecasting results for France, table 8.8 for Spain and table 8.9 for Portugal.

Table 8.7: Forecasting results for the UK expenditure share of France

	1994	1995	1996	1997
Actual values	**0.39700**	**0.38540**	**0.38967**	**0.40481**
Forecasts•	0.34472	0.30937	0.32640	0.29533
Forecast error•	0.052281	0.076035	0.063271	0.10948
Forecasts••	0.34490	0.32012	0.32887	0.30352
Forecast error••	0.036838	0.038495	0.042336	0.10129
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE•	0.010997		0.075267	
MSE•	0.0001751		0.0061260	
RMSE•	0.013233		0.078269	
MAE••	0.011914		0.069866	
MSE••	0.0001986		0.0052329	
RMSE••	0.014094		0.072339	

Table 8.8: Forecasting results for the UK expenditure share of Spain

	1994	1995	1996	1997
Actual values	**0.51857**	**0.52625**	**0.52292**	**0.50691**
Forecasts•	0.56625	0.60544	0.58724	0.61457
Forecast error•	-0.047682	-0.079192	-0.064320	-0.10766
Forecasts••	0.56555	0.59186	0.58387	0.60660
Forecast error••	-0.046982	-0.065614	-0.060958	-0.099696
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 -1997)	
MAE•	0.010966		0.074713	
MSE•	0.0002060		0.0060682	
RMSE•	0.014353		0.077898	
MAE••	0.012295		0.068312	
MSE••	0.0002572		0.0050419	
RMSE••	0.016037		0.071006	

Table 8.9: Forecasting results for the UK expenditure share of Portugal

	1994	1995	1996	1997
Actual values	**0.084433**	**0.088348**	**0.087411**	**0.088283**
Forecasts•	0.089032	0.085192	0.086362	0.090104
Forecast error•	-0.0045990	0.0031559	0.0010490	-0.0018211
Forecasts••	0.089544	0.088015	0.087253	0.089875
Forecast error••	-0.0051110	0.0003331	0.0001578	-0.0015922
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 –1997)	
MAE•	0.0064103		0.002656	
MSE•	0.0000597		0.0000088	
RMSE•	0.0077283		0.0029802	
MAE••	0.0063505		0.001799	
MSE••	0.0000602		0.0000072	
RMSE••	0.0077599		0.0026830	

8.3.3.4. VAR models

In chapter 7 we modelled the UK demand for tourism using the VAR methodology and the same variables included in the AIDS model. We estimated and forecasted different specifications of the VAR model and came to the conclusion that good forecasts were obtained with the unrestricted reduced-form ('pure') VAR, and the best forecasts were obtained with the structural cointegrated VAR conditioned on exogenous variables under the full set of exactly- and over-identifying restrictions. The forecasting results for the tourism shares of France, Spain and Portugal obtained with the former (denoted by Unrestricted-VAR I), were presented in tables 7.7, 7.8 and 7.9. The forecast results for the same destinations' shares obtained with the latter (denoted by over-VAR III) were presented in tables 7.15, 7.16 and 7.17.

We now report this same information for the expenditure shares of France, Spain and Portugal in tables 8.10, 8.11 and 8.12, respectively. The symbol □ is used to denote the forecasting results obtained with the unrestricted-VAR I model. The symbol □□ is used to denote the forecasting results obtained with the over-VAR III.

Table 8.10: Forecasting results for the UK expenditure share of France

	1994	1995	1996	1997
Actual values	**0.39700**	**0.38540**	**0.38967**	**0.40481**
Forecasts$^{\square}$	0.34422	0.36840	0.39676	0.42782
Forecast error$^{\square}$	0.052780	0.017006	-0.007086	-0.023007
Forecasts$^{\square\square}$	0.39505	0.39260	0.39485	0.39475
Forecast error$^{\square\square}$	0.001942	-0.007194	-0.005178	0.010054
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 –1997)	
MAE$^{\square}$	0.015321		0.024970	
MSE$^{\square}$	0.000315		0.000914	
RMSE$^{\square}$	0.017734		0.030226	
MAE$^{\square\square}$	0.016342		0.006092	
MSE$^{\square\square}$	0.000362		0.000046	
RMSE$^{\square\square}$	0.019016		0.006772	

Table 8.11: Forecasting results for the UK expenditure share of Spain

	1994	1995	1996	1997
Actual values	**0.51857**	**0.52625**	**0.52292**	**0.50691**
Forecasts$^{\square}$	0.57000	0.54972	0.52281	0.49058
Forecast error$^{\square}$	-0.051431	-0.023469	0.000108	0.016328
Forecasts$^{\square\square}$	0.51884	0.52718	0.52568	0.52543
Forecast error$^{\square\square}$	-0.000264	-0.000932	-0.002761	-0.018518
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 –1997)	
MAE$^{\square}$	0.016475		0.022834	
MSE$^{\square}$	0.000301		0.000866	
RMSE$^{\square}$	0.019780		0.029422	
MAE$^{\square\square}$	0.015845		0.005619	
MSE$^{\square\square}$	0.000371		0.000088	
RMSE$^{\square\square}$	0.019268		0.009374	

Table 8.12: Forecasting results for the UK expenditure share of Portugal

	1994	1995	1996	1997
Actual values	**0.08443**	**0.08835**	**0.08741**	**0.08828**
Forecasts$^{\square}$	0.08578	0.08189	0.08043	0.08160
Forecast error$^{\square}$	-0.001349	0.006462	0.006977	0.006679
Forecasts$^{\square\square}$	0.08611	0.08022	0.07947	0.07982
Forecast error$^{\square\square}$	-0.001678	0.008124	0.007939	0.008463
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Estimation period: (1970 -1993)		Forecast period: (1994 –1997)	
MAE$^{\square}$	0.008054		0.005367	
MSE$^{\square}$	0.000088		0.000034	
RMSE$^{\square}$	0.009365		0.005850	
MAE$^{\square\square}$	0.006155		0.006551	
MSE$^{\square\square}$	0.000063		0.000051	
RMSE$^{\square\square}$	0.007914		0.007132	

8.4. FORECASTING PERFORMANCE OF ECONOMETRIC MODELS FOR THE UK TOURISM DEMAND: COMPARATIVE EVALUATION

For each econometric model previously specified, we now compare the forecasts of the relevant dependent variable with its actual values over the forecasting period 1994-1997. The objective is to see how closely the forecasts track its corresponding actual values. The different statistics for the forecast errors reported in tables 8.1 to 8.12 measure the ability of each econometric specification to predict the dependent variable values over the forecasting period. The smaller the value of the statistics, the closer the forecasts are to the observed values of the variable. The quality statistics associated with each model, indicate that the forecasts of the UK demand for tourism in France, Spain and Portugal closely track the actual data and suggest that each of the models has good predictive ability. Yet, however useful these statistics may be to assess the prediction quality of one particular specification, to obtain information about whether better forecasts might be achieved with alternative models, we have to compare their values across all forecasting devices we have available.

We need to go further in our analysis in order to compare the predictive power of different econometric models and to come to a conclusion about which one provides the best forecasts of the UK tourism shares for France, Spain and Portugal. Indeed, we have to contrast the competing sets of forecasts and respective quality statistics obtained with the ARDL models of chapter 4, the AIDS models of chapters 5 and 6 and the VAR models of chapter 7. Yet, for this judgement to be valid, we need to assure that the competing sets of forecasts are homogeneous, in the sense that they represent predicted values of the same time series variable projected for the same time horizon.

The forecasts obtained with each model refer to the same time horizon, which include the four last observations of the out-of-sample forecasting period (1994-1997). However, the variable we are interested in forecasting (UK demand for tourism) is measured in different ways according to the alternative econometric specification underlying the prediction process. For example, while in the AIDS equations of chapter 5 the dependent variable measuring the UK tourism demand for France, Spain and Portugal, is the *share* of the UK tourism expenditure in each destination, in the ARDL models of chapter 4 the dependent variable measuring the same entity, is the log of the UK real per capita expenditure in each destination. Since the forecasts obtained with the AIDS models of chapter 5 and 6 and the VAR models of chapter 7 are those for the level values of the UK expenditure percentage *share* of each destination[13], the only model providing forecasts for a different time series is the ARDL model of chapter 4.

Consequently, for forecast comparison purposes, the dependent variable in the ARDL models has to be transformed into a time series representing the level of the UK tourism budget *share* of each destination. The transformation process, although time consuming, is straightforward. Taking the ARDL models and forecasting the original dependent variable for the relevant forecasting horizon, we obtain the forecasts given in section 8.3.3.1. We then compute the antilog of these forecasts and multiply each antilog forecasted value by the appropriate deflator that was originally used in the ARDL model to change nominal expenditure into real expenditure. In this way we obtain the nominal level values of the UK tourism expenditure allocated to each destination. Summing these values

[13] Although the original dependent variables in the AIDS model of chapter 6 are the first differences of the destinations' shares and not their level values as required, we showed in section 8.3.3.3. how to transform the original first difference variables into the desired level variables. In the same section we computed the forecasts for the destinations' shares in *levels* and not in first differences. Hence, these forecasts are readily comparable with those of the static AIDS model of chapter 5 and VAR models of chapter 7.

across destinations for each year of the forecasting period, we obtain the total nominal per capita tourism expenditure allocated to the whole set of destinations. Dividing the expenditure in each country by this total, we obtain the level value forecast of the expenditure share in each destination for each year of the forecasting period. Note that the shares' level values are invariant whether per capita or total expenditure is considered.

By this process we are able to transform the original "real per capita expenditure" forecasts of the ARDL model into the "expenditure *shares*" forecasts, which are the relevant forecasts for comparison purposes as they are the time series predicted with all the other econometric models. Hence, for the ARDL models of chapter 4, tables 8.13, 8.14 and 8.15 present the actual and 'transformed' forecast values of the expenditure *shares* of France, Spain and Portugal, respectively, and the summary statistics previously used for evaluating the accuracy of the forecasts.

Table 8.13: Forecasting results of the ARDL model for the expenditure share of France

	1994	1995	1996	1997
Actual values	**0.3970**	**0.3854**	**0.3897**	**0.4048**
Forecasts	0.4140	0.3926	0.3951	0.4170
Forecast error	-0.0170	-0.0072	-0.0054	-0.0122
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Forecast period: (1994 -1997)			
MAE	0.01045			
MSE	0.0001297			
RMSE	0.011389			

Table 8.14: Forecasting results of the ARDL model for the expenditure share of Spain

	1994	1995	1996	1997
Actual values	**0.5186**	**0.5263**	**0.5229**	**0.5069**
Forecasts	0.4946	0.5215	0.5206	0.5085
Forecast error	0.0240	0.0048	0.0023	-0.0016
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Forecast period: (1994 -1997)			
MAE	0.008175			
MSE	0.0001517			
RMSE	0.012318			

Table 8.15: Forecasting results of the ARDL model for the expenditure share of Portugal

	1994	1995	1996	1997
Actual values	**0.0844**	**0.0883**	**0.0874**	**0.0883**
Forecasts	0.0914	0.0859	0.0843	0.0746
Forecast error	-0.0070	0.0024	0.0031	0.0137
SUMMARY STATISTICS FOR RESIDUAL AND FORECAST ERRORS				
	Forecast period: (1994 -1997)			
MAE	0.006550			
MSE	0.0000630			
RMSE	0.007938			

We can now gather together all previous information, and compare the ability of the ARDL, static AIDS, dynamic AIDS and VAR models to forecast the UK tourism expenditure shares of France, Spain and Portugal, over the period 1994 -1997. Comparison and the choice of the best forecasting model is facilitated by including in tables 8.16, 8.17 and 8.18, the forecasting results for the shares of France, Spain and Portugal, respectively, obtained with all the relevant econometric specifications namely: unrestricted static AIDS model (denoted by static-AIDS); static AIDS under homogeneity, symmetry and null cross-price effects between the equations of France and Portugal [denoted by static-AIDS(H+S)^0]; dynamic AIDS model under homogeneity, symmetry and the same null cross-price effects imposed only on the long-run coefficients [denoted dynamic-AIDS(H+S)^{0}L]; dynamic AIDS model under the same set of restrictions imposed on both the long-run and short-run coefficients [denoted dynamic-AIDS(H+S)^{0}LS]; unrestricted ARDL (denoted by ARDL); unrestricted reduced-form ('pure') VAR model (denoted by VAR I); cointegrated VAR under the full set of exactly- and over-identifying restrictions (denoted by over-VAR III).

Table 8.16: Forecast results for the UK expenditure share of France

FRANCE		1994	1995	1996	1997	MAE	RMSE
Actual values		**0.3970**	**0.3854**	**0.3897**	**0.4048**		
Static-AIDS	Forecast	0.4003	0.3971	0.4271	0.4602	0.0269	0.0339
	Error	-0.0033	-0.0117	-0.0374	-0.0553		
Static-AIDS $(H+S)^0$	Forecast	0.3778	0.3804	0.4019	0.3883	0.0132	0.0143
	Error	0.0193	0.0050	-0.0122	0.0165		
Dynamic-AIDS$(H+S)^0$L	Forecast	0.3447	0.3094	0.3264	0.2953	0.0753	0.0783
	Error	0.0523	0.0760	0.0633	0.1095		
Dynamic-AIDS$(H+S)^0$LS	Forecast	0.3449	0.3201	0.3289	0.3035	0.0699	0.0723
	Error	0.0368	0.0385	0.0423	0.1013		
ARDL	Forecast	0.4140	0.3926	0.3951	0.4170	0.0105	0.0114
	Error	-0.0170	-0.0072	-0.0054	-0.0122		
VAR I	Forecast	0.3442	0.3684	0.3968	0.4278	0.0250	0.0302
	Error	0.0528	0.0170	-0.0071	-0.0230		
Over-VAR III	Forecast	0.3951	0.3926	0.3949	0.3948	0.0061	0.0068
	Error	0.0019	-0.0072	-0.0052	0.0101		

Table 8.17: Forecast results for the UK expenditure share of Spain

SPAIN		1994	1995	1996	1997	MAE	RMSE
Actual values		**0.5186**	**0.5263**	**0.5229**	**0.5069**		
Static-AIDS	Forecast	0.5007	0.5041	0.4691	0.4288	0.0430	0.0495
	Error	0.0179	0.0222	0.0538	0.0781		
Static-AIDS(H+S)^0	Forecast	0.5261	0.5224	0.4979	0.5126	0.0106	0.0135
	Error	-0.0076	0.0039	0.0251	-0.0057		
Dynamic-AIDS(H+S)^{0}L	Forecast	0.5663	0.6054	0.5872	0.6146	0.0747	0.0779
	Error	-0.0477	-0.0792	-0.0643	-0.1077		
Dynamic AIDS(H+S)^{0}LS	Forecast	0.5656	0.5919	0.5839	0.6066	0.0683	0.0710
	Error	-0.0470	-0.0656	-0.0610	-0.0997		
ARDL	Forecast	0.4946	0.5215	0.5206	0.5085	0.0082	0.0123
	Error	0.0240	0.0048	0.0023	-0.0016		
VAR I	Forecast	0.5700	0.5497	0.5228	0.4906	0.0228	0.0294
	Error	-0.0514	-0.0235	0.0001	0.0163		
Over-VAR III	Forecast	0.5188	0.5272	0.5257	0.5254	0.0056	0.0094
	Error	-0.0003	-0.0009	-0.0028	-0.0185		

Table 8.18: Forecast results for the UK expenditure share of Portugal

PORTUGAL		1994	1995	1996	1997	MAE	RMSE
Actual values		**0.0844**	**0.0883**	**0.0874**	**0.0883**		
Static-AIDS	Forecast	0.0991	0.0989	0.1038	0.1111	0.0161	0.0167
	Error	-0.0146	-0.0105	-0.0164	-0.0228		
Static-AIDS(H+S)^0	Forecast	0.0961	0.0972	0.1003	0.0991	0.0110	0.0111
	Error	-0.0117	-0.0089	-0.0129	-0.0108		
Dynamic-AIDS(H+S)^{0}L	Forecast	0.0890	0.0852	0.0864	0.0901	0.0027	0.0030
	Error	-0.0046	0.0032	0.0010	-0.0018		
Dynamic-AIDS(H+S)^{0}LS	Forecast	0.0895	0.0880	0.0873	0.0899	0.0018	0.0027
	Error	-0.0051	0.0003	0.0002	-0.0016		
ARDL	Forecast	0.0914	0.0859	0.0843	0.0746	0.0066	0.0079
	Error	-0.0070	0.0024	0.0031	0.0137		
VAR I	Forecast	0.0858	0.0819	0.0804	0.0816	0.0054	0.0059
	Error	-0.0013	0.0065	0.0070	0.0067		
Over-VAR III	Forecast	0.0861	0.0802	0.0795	0.0798	0.0066	0.0071
	Error	-0.0017	0.0081	0.0079	0.0085		

The first thing to be said about the forecasting performance of the models is that it seems to be fairly accurate overall. In fact, if forecasts had to be obtained using only one of the models without availability of competing sets of forecasts for comparison purposes, any of them would be qualified as a good forecasting device. However, confronted with supplementary information from other sets of forecasts, we can decide which of the econometric models is more reliable for forecasting the UK tourism budget shares of France, Spain and Portugal.

Analysis is facilitated by interpretation of both the MSE and RMSE statistics in terms of the original units of measurement of the variable forecasted. This variable is each destination's share of the UK tourism budget and is measured in percentage points (pp). Hence, the statistics' values represent the average deviations of the forecasts from the actual values, measured in percentage points. For instance, the values of MSE (0.0066) and RMSE (0.0071) for the over-VAR III model forecasting the share of Portugal may be interpreted as the amount by which the forecasts of this variable deviate from the corresponding true values of the series. In this case, then, the forecasts deviate from the variable's actual value by an (absolute) error of less than one percentage point (0.7 pp) on average.

Another complementary measure of accuracy can be derived by comparing the average forecast errors with the average actual values of the series. This measure can be crucial for the correct interpretation and comparison of the models' forecasting performance. Indeed, the interpretation of a computed average (absolute) error of less than 1pp for a series of observed shares averaging 8%, is quite different from that of the same error for a series averaging 50%. This is so because a statistic value in the interval (0.000; 0.005), which means a maximum average (absolute) error of half of a percentage point, leads to different (relative) error intervals depending on the average size of the share's actual values. For example, considering the forecasting period 1994-1997, a statistic value of half of a percentage point (0.005) corresponds to a 5.7% (relative) error of the average share for Portugal (0.0871), as 0.005/0.0871=0.057. Yet, the same statistic value means less than 1% error relative to the average share for Spain (0.5187), as 0.005/0.5187=0.0096. We think that the importance of these aspects justifies the transformation of average (absolute) error intervals given by the quality statistics, into average (relative) error intervals which take into consideration the weight of the errors relative to the actual series. Since, overall, both the MSE and RMSE quality criteria give similar information, we concentrate attention on the RMSE statistic hereafter.

For the purpose of comparing absolute errors with relative errors, we present table 8.19 which shows, for the destinations' average share values in the period 1994-1997 (denoted by $\overline{W}_{94-97}$), the absolute error range of the RMSE statistic, the corresponding relative error range and the models for which the forecast errors lay within these ranges.

Table 8.19: Absolute and relative forecast error range for the destination shares

RMSE absolute error range	Destination	$\overline{W}_{94-97}$	RMSE Relative error range	Models and respective RMSE absolute and relative errors
(0.000; 0.005)	France	0.3942	(0% - 1.5%)	
	Spain	0.5185	(0% - 1%)	
	Portugal	0.0871	(0% - 6%)	Dynamic-AIDS(H+S)^{0}L: 0.0030 (3.4%) AIDS(H+S)^{0}LS: 0.0027 (3.1%)
(0.005; 0.010)	France	0.3942	(1.5% - 2.5%)	Over-VAR III: 0.0068 (1.7%)
	Spain	0.5185	(1% - 2%)	Over-VAR III: 0.0094 (1.8%)
	Portugal	0.0871	(6% - 12%)	VAR I: 0.0069 (7.9%) Over-VAR III: 0.0071 (8.2%) ARDL: 0.0079 (9.1%)
(0.010; 0.015)	France	0.3942	(2.5% - 4%)	ARDL: 0.0114 (2.9%) Static-AIDS(H+S)^0: 0.0143 (3.6%)
	Spain	0.5185	(2% -3%)	ARDL: 0.0123 (2.4%) Static-AIDS(H+S)^0: 0.0135 (2.6%)
	Portugal	0.0871	(12% - 18%)	Static-AIDS(H+S)^0: 0.0111 (12.7%)
(0.015; 0.020)	France	0.3942	(4% - 6%)	
	Spain	0.5185	(3% - 4%)	
	Portugal	0.0871	(18% - 23%)	Static-AIDS: 0.0167 (19.2%)

Table 8.19 shows that most models produce forecasts which, on average, differ from the actual series by less than 1.5 percentage points. This low average error indicates these models as good forecasting devices. However, while an absolute error below 1.5 percentage points means an error below 4% relative to the average shares of France or Spain, the same absolute error can mean an error of up to 18% relative to the average share of Portugal. This indicates that models within the same quality range for all destination shares, such as the static-AIDS(H+S)^0 model with an average absolute errors ranging from 1 to 1.5

percentage points whatever the destination, produces average relative errors ranging from 2.5% to 4% of the share for France, 2% to 3% of the share for Spain and 12% to 18% of the share for Portugal. On the other hand, the ARDL has an absolute error range between 1 to 1.5 percentage points for the shares of Spain and France and, for the share of Portugal, has the lower rage of 0.5 to 1 pp. This model gives for the former a relative error of less than 4%, while for the latter the relative error is 9.1%. Hence, the forecasting performance of models can differ considerably in the relative errors they produce, even when they are qualified in the same absolute error range. As a result, when interpreting the accuracy of the quality statistics, a researcher must be aware of both the absolute and relative average errors for a more reliable ranking of the models.

We now turn to the analysis of the results shown in table 8.18. An immediate indication given by the results of the forecast quality statistics for all destinations' shares, is that unrestricted or partially restricted models perform worse than those taking full account of the theoretical bounds of consumers behaviour. The performance difference between the unrestricted and fully restricted specifications of the static AIDS and VAR models is remarkable. The performance difference between the partially restricted and the fully restricted versions of the dynamic AIDS model is not as significant. This is so because, while the unrestricted version of the static AIDS and the VAR models does not consider any of the constraints suggested by economic theory, the partially restricted dynamic model, although not considering them in the short-run, does so in the long-run. Thus, the results seem to indicate that the imposition of appropriate theoretical restrictions on the coefficients of econometric models leads to better forecasting performance. In econometrics, as in physics where researchers often say that 'what nature does not forbid is compulsory in theory', we say that 'what is not (statistically) forbidden, is (modelling) compulsory'. Therefore, models which do not reject appropriate theoretical constraints must be modelled under these hypotheses, allowing them better to describe and forecast the underlying economic reality.

Another straightforward indication given by the forecasting results reported is that of the best and worst forecasting system of equations model. The best is the cointegrated VAR model under the full set of exactly- and over-identifying restrictions, which produces average absolute errors below 1 percentage point for all destination shares, and average relative errors below 2% of the average shares for France and Spain and around 8% of the average share for Portugal. The worst, is the unrestricted static AIDS model, which produces average absolute errors ranging 1.5 to 5 percentage points for all destination shares, and average relative errors exceeding 8% of the average share for France, 9% of the average share of Spain and 19% of the average share for Portugal. Hence, for the remainder of this forecasting comparison exercise, we exclude the unrestricted static AIDS model from the comparative analysis.

The second best forecasting system seems to be the restricted static AIDS model. This model supplies forecasts with an average absolute error not exceeding 1.5 percentage points for all destination shares, and average relative errors below 4% of the average share for France, 3% of the average share for Spain and 13% of the average share of Portugal. However, if we analyse the systems on an equation-by-equation basis, the ability of the static AIDS model to forecast the shares of Portugal, Spain and France is out-performed by that of all the other models in the case of Portugal, and by the ARDL model in the cases of France and Spain. This single equation model, produces average absolute errors below 1 pp for the share of Portugal and below 1.5 pp for the shares of France and Spain. Furthermore, the corresponding average relative errors of this model's forecasts are 2.9% of the average share for France, 2.6% of the average share for Spain and 9.1% of the average share for Portugal.

Nevertheless, the analysis of the predictive ability of a system should be carried out on a system basis, rather than an equation-by-equation basis, as that of a single equation model should be carried out on a separate equation basis and not as a system. Indeed, there are queries which can only be answered by considering all equations in a system as a whole entity. For instance, aside from the fact that the best and the worst forecasting models can be readily pointed out from the information reported earlier, there seems to be no easy way of distinguishing between the other models in their relative predictive power. Hence, there seems to be space for some originality in conceiving a method of ranking the models according to the accuracy of their forecasts.

The forecasting performance of the seven models is ranked by awarding, for each destination share, a maximum of 7 marks to the specification with the lowest average absolute error, 6 marks to the next lowest and so on. This is done in such a way that the model with the highest average absolute error gets only 1 mark. Then, summing the marks of each model across destination shares, we obtain a total which can be used to compare the forecasting accuracy of the models. For example, while the dynamic VAR system of equations under the full set of exactly and over-identifying restrictions (denoted by over-

VAR III) gets 7 marks in the share equation for France, 7 marks in the share equation for Spain and 4 in the share equation for Portugal, summing a total of 18, the static unrestricted AIDS system (denoted by static-AIDS) obtains only 3+3+1=7 marks. Figure 8.1 shows the forecasting models according to their specification features and associated total sum of marks in brackets.

Figure 8.1: Models' specification features and rank.

$$\text{MODEL}\begin{cases}\text{SYSTEM}\begin{cases}\text{STATIC}\begin{cases}\text{UNRESTRICTED}\Rightarrow\text{AIDS}\ \mathbf{(7)}\\ \text{RESTRICTED}\Rightarrow\text{AIDS}(\text{H}+\text{S})\ \mathbf{(12)}\end{cases}\\ \text{DYNAMIC}\begin{cases}\text{UNRESTRICTED}\Rightarrow\text{VAR I}\ \mathbf{(13)}\\ \text{PARTIALLY RESTRICTED}\Rightarrow\text{AIDS}(\text{H}+\text{S})\text{L}\ \mathbf{(8)}\\ \text{RESTRICTED}\Rightarrow\begin{cases}\text{AIDS}(\text{H}+\text{S})\text{LS}\ \mathbf{(11)}\\ \text{Over}-\text{VAR III}\ \mathbf{(18)}\end{cases}\end{cases}\end{cases}\\ \text{SINGLE}\begin{cases}\text{STATIC}\\ \text{DYNAMIC}\left\{\text{UNRESTRICTED}\Rightarrow\text{ARDL}\ \mathbf{(15)}\right.\end{cases}\end{cases}$$

With this unorthodox method of ranking the prediction quality of different models, we can easily establish that systems of equations are better forecasters than single equation models, dynamic models out-perform static models, and fully restricted specifications provide more accurate forecasts than unrestricted or partially restricted models. We can also readily spot the best forecaster – the over-VAR III model with 18 marks – and the worst forecaster – the unrestricted static-AIDS model with 7.

This schematic way of ranking the models also permits the identification of aspects that may appear as inconsistencies. For instance, if systems of equations are better forecasters than single equation models why does the ARDL model out-perform the AIDS system of equations? The explanation resides in the fact that the dynamic aspects of the UK demand for tourism seem to be more relevant, on an equation-by-equation basis, than the potential interactions existing among destinations are on a system of equations basis. Consequently, an econometric specification such as the ARDL model that takes into account the short-run dynamics of the adjustment process, although not contemplating equations interactions, is expected to forecast better than a model which does not take explicit account of dynamics. These are important characteristics of UK tourists' behaviour which, once included within a model's structure, can make it a powerful and reliable forecasting device. However, the single equation form of the ARDL model does not allow for testing the theories of consumer demand behaviour and this limitation can affect its qualities.

On the other hand, the ARDL specification differs substantially from all the other models in the way it portrays the UK tourism demand behaviour towards the destinations considered. This specification models the UK tourism budget allocated to each destination as a function of prices and income. It does not model destination shares as a function of prices and tourism expenditure. Hence, although we could "transform" the original ARDL forecasts of tourism budgets (measured in pounds sterling) into tourism shares (measured in percentage points), we did not "transform" the model itself and we cannot directly compare models with different dependent variables. Apart from this "transformation" which made comparison of forecasts possible, the ARDL model remains a different approach from all the other econometric models. Unless the same variables entering the AIDS and VAR systems can be used to model the UK demand for tourism within an ARDL framework the comparisons between the former and the latter are limited. The ARDL modelling with the same variables included in the AIDS and VAR models is possible, but it is a matter for future research.

Nevertheless, an ambiguity needing clarification still remains in the model ranking scheme displayed above. If dynamic specifications are expected to forecast better than static ones, why does the restricted static AIDS model out-perform the restricted dynamic AIDS? Given the inherent dynamic nature of tourism demand, one would expect a general dynamic AIDS specification to forecast better than a static one, irrespective of the destination. Yet, a comparison of the restricted forms of the static and dynamic AIDS models seems to favour the static rather than the dynamic in the cases of France and Spain, while in the case of Portugal the opposite conclusion can be drawn. Two sets of reasons can be suggested to explain these findings: one can be qualified as statistical; the other embeds theoretical and empirical arguments.

The static AIDS model specified in chapter 5 appears to be statistically more robust than the dynamic AIDS model of chapter 6. Although both models respond positively to all the diagnostic tests performed for evaluating their statistical quality, the static AIDS model seems to do so more convincingly than the dynamic model. For instance, several of the coefficients of the equations in the dynamic system appear to be statistically insignificant. Yet, a solution to improve this model is not the simply deletion of the insignificant variables as this would jeopardise the dynamic *general* structure we wanted to implement in the first place, and disturb the invariance of the coefficients' estimates conferred by the adding-up property within the SUR estimation method. On the other hand, retaining the insignificant variables may lead to imprecision of the estimates which may take away some of the model's accuracy. These considerations seem to indicate that the more statistically robust a model is, the more reliable it appears to be for forecasting purposes.

However, if we compare the estimation results of the equations for Portugal in the static and in the dynamic AIDS models, the static specification indicates statistically better estimation results than the dynamic and yet, the dynamic AIDS model still provides better forecasts of the UK demand for Portugal than the static model. It seems, then, that the explanation must reside in more areas than just the statistical robustness of the models. The explanation may be found in two major aspects of the short-run dynamic adjustment underlying tourism demand behaviour for each destination: the factors that are mainly responsible for dragging the dependent variable towards its equilibrium path, and the econometric tools being used to depict this adjustment mechanism. These will now be considered.

One key aspect of the particular static AIDS model specified in chapter 5 is that it is not an orthodox AIDS model. The Deaton and Muellbauer (1981) orthodox AIDS model, regresses destinations' shares on prices and per capita expenditure. The "unorthodox" form of the AIDS model adopted in chapter 5, regresses the same dependent variable on prices, expenditure and a time trend variable, and models a structural break for the coefficient of the income variable. Recall that the orthodox version of the AIDS model was convincingly rejected by the data, either when tested against the unorthodox form of chapter 5, or against the general dynamic form specified in chapter 6. Therefore, we conclude that on the one hand, a trend variable must be included in an AIDS specification which does not explicitly account for the inherent short-run dynamics of tourism demand and, on the other hand, when the equilibrium adjustment mechanism is explicitly modelled, the trend variable becomes superfluous and the orthodox AIDS model is rejected.

In the AIDS model of chapter 5, the trend is a highly significant explanatory variable in the equations for France and Spain but not so in the equation for Portugal. In the equations for France and Spain, the trend variable seems to be a key factor in mimicking the inherent dynamic nature of tourism demand behaviour, while in the equation for Portugal it seems to fail this objective. Apparently, the dynamic process by which tourists' behaviour adjusts to their long-run equilibrium in the share equation for Portugal, is different from those for France and Spain. This may be a reason why the AIDS model of chapter 5 forecasts better than the dynamic AIDS in the cases of France and Spain but not so in the case of Portugal.

A static model is, by definition, a steady-state long-run equilibrium model and its coefficients give information about the long-run impacts that changes in the explanatory variables have on the dependent variable. Therefore, if for any particular destination the long-run effects happen to be more important than the short-run ones, a steady state parsimonious static AIDS model with trend is likely to be more adequate than the complex and over-parameterised structure of the general dynamic AIDS. If, in addition, the trend variable inserted in the steady-state model can capture features of the demand behaviour, like habit persistence, which are not accommodated by other regressors, we have a main reason for believing that for destinations for which these features are particularly relevant, a static model with a trend is able to describe better the underlying demand behaviour than its over-parameterised dynamic counterpart. We believe this to be the case for France and Spain. For Portugal, however, the relevant dynamic factors seem to be linked with other reasons, and habit persistence features do not appear to be a characteristic of the UK tourism demand for this country. Hence, in this case, an explicit lag structure of the regressors might be able to capture better the demand dynamic behaviour. This seems to be confirmed by the insignificance of the trend variable in the equation for Portugal in the static AIDS model, and by the superior forecasting performance of the general dynamic form of this model for this particular destination.

However, as noted previously, the comparative analysis of the predictive ability of econometric systems should be conducted considering each system as a whole, rather than in an equation-by-equation perspective. From this point of view, the forecasting ability of the unrestricted VAR model out-performs that of all AIDS specifications, being second best only to the cointegrated fully restricted VAR. Given its formal simplicity and ease of estimation, the reduced-form unrestricted (pure) VAR has all the desirable qualities to make its choice an easy one for forecasting purposes. Yet, if the most accurate forecasts of the destinations' shares are required, then the predictive ability of the fully restricted VAR overshadows that of all other models. This is not surprising since the fully restricted VAR incorporates the three main features of a quality forecasting model: it takes explicit account of short-run dynamics, includes theoretical hypotheses accounting for the rational behaviour of consumers and uses a formal specification matching theory and data in a quantitative framework which is both empirically plausible and statistically robust.

8.5. CONCLUSION

Kierkegaard says that life must be lived forwards, but can only be understood backwards. This applies also to economics, where the reality of a particular phenomenon may only become apparent by understanding its past behaviour, and the ways in which it evolves can only be 'lived' in the future. Yet, if economic behaviour could be foreseen in the present, it would permit a more knowledgeable approach to its future consequences. This is the reason why accurate forecasts are so important. They are fundamental inputs in decision-making processes and economic policy strategy.

An appropriately specified econometric model can be projected into the future to forecast the behaviour of the underlying economic reality it portrays. Yet, econometric models are more than simple forecasting devices. These models are formal quantitative relationships, linking theory and observed data, which allow us to understand economic behaviour, interpret economic facts, test economic theories and evaluate economic policy. Forecasting, however, has a crucial role in economic analysis and is an integral part of the evaluation process of econometric modelling. Therefore, the goal of this chapter was to obtain forecasts from the econometric models specified previously and to address the key problem of evaluating their statistical quality by comparing their forecasting performance.

Many methods can be used to obtain time series forecasts. However, irrespective of the method selected, there are aspects which have to be taken into account in order to ensure reliability in the method employed and reasonable accuracy in the forecasts produced. First, the time path of the phenomenon we want to predict has to embed certain 'regularities' which self-reproduce in the future. Then, these regularities and the features of their self-reproducing process must be captured and adequately depicted by the selected forecasting method. The former is a characteristic of any economic time series. The latter depends on the specific forecasting model adopted to project the series into the future.

The empirical evidence gathered in previous chapters indicates that the quantitative frameworks chosen for modelling the UK demand for tourism were theoretically coherent, empirically plausible and statistically robust representations of the regularities underlying the economic phenomenon under investigation. The ARDL model specified in chapter 4, the AIDS systems estimated in chapters 5 and 6 and the VAR approach used in chapter 7 were considered to be, in their own right, appropriate means of anticipating the level values of the UK tourism demand for France, Spain and Portugal. Under this assumption, we used those models, estimated over the in-sample period 1969-1993, to obtain point forecasts of the destinations' shares in the out-of-sample period 1994-1997.

The statistics used to measure the accuracy of the forecasts indicate that most specifications, namely the ARDL model, the unrestricted and fully restricted VAR specifications and the restricted static AIDS system, produce forecasts with average absolute errors below 1.5 percentage points, which can be viewed as reasonably accurate. Beyond this range, for one or more destination shares, are the unrestricted static and the dynamic AIDS specifications. The results also indicate that dynamic models perform better than static ones, fully restricted models generate more precise forecasts than their unrestricted or partially restricted counterparts, and the cointegrated VAR specification under the full set of restrictions is the best forecasting model.

Therefore, if the goal is to obtain the most accurate forecasts possible of the UK tourism budget shares for France, Spain and Portugal, the appropriate econometric modelling has to take into account dynamics and the imposition of appropriate theoretical restrictions. Furthermore, the quantitative framework within which the share variables are modelled should preferably involve a system of equations where the variables are not *a priori* assumed as exogenously given, and an appropriate account of qualitative events which significantly affect the path of the dependent variables over the sample period. An econometric model with these characteristics is the cointegrated VAR specification, of which the structural form is exactly-identified with restrictions matching those used to identify the share equations of a steady-state AIDS system.

However, if a cost-benefit perspective is important and less accurate forecasts are not a problem, the unrestricted reduced-form ('pure')VAR model has to be considered the best predictor. The model's remarkable qualities of simplicity of form, ease of estimation and minimal assumptions alongside its fairly accurate forecasting ability, give it prime place relative to the other econometric models. Yet, if a sensible economic interpretation of the long-run relationships is also required, the estimation of this model's structural form does not help. Rather, its coefficients are revealed as theoretically inconsistent, empirically implausible and statistically insignificant estimates of the corresponding long-run parameters. In contrast, the estimation of the fully restricted cointegrated VAR gives coherent, plausible and significant estimates of the structural long-run relationships linking the variables within the model. Furthermore, these estimates are similar to those obtained with the restricted AIDS system providing additional support for this model to be considered as a reliable means of explaining the behaviour of the destination tourism shares over the period 1969-1997.

CHAPTER 9

CONCLUSION

The key role that tourism can play in assisting economic growth in destinations has been broadly acknowledged. Accordingly, research on the economics of tourism has been a growing area of interest, and tourism demand modelling and forecasting studies have been reported more frequently in economics literature. However, the majority of these studies fail to apply theoretical and methodological tools which are fundamental to the construction of accurate and reliable models for explaining and predicting tourism phenomena. The recognition of these models' limitations has recently demonstrated the need for studies that, on the one hand, clearly identify the areas that require improvement and, on the other hand, provide methods which can ameliorate the reliability of the results. This new approach to tourism demand analysis consists of both extending the theoretical and empirical content of previous quantitative models and of presenting alternative formal specifications that can validly describe and predict tourism demand behaviour. These specifications must be based on sound theoretical grounds and subjected to extensive testing to substantiate, or otherwise, their empirical findings.

In this thesis, we examine the methodologies applied in the existing literature on tourism demand analysis and point out their limitations. Then, we discuss alternative methodological approaches which contribute both to enlarging the theoretical basis of currently used models, and implementing recent advances in econometric modelling, quality evaluation, hypothesis testing and forecasting procedures. Throughout the study we show how models based on sound theoretical grounds and constructed under the appropriate econometric rules can withstand rigorous statistical scrutiny and provide a consistent and reliable description of tourism demand and accurate predictions of its behaviour.

This study makes several important contributions. The thesis analyses the UK demand for tourism in its geographically proximate neighbours, France, Spain and Portugal, using data for the period 1969-1997. The choice of the countries involved took into account the fact that the UK is a major tourism origin which is of particular importance to France, Spain and Portugal as destinations. Moreover, Spain and Portugal are interesting cases for analysis as they were experiencing considerable economic development during the sample period. At the beginning of the period, they demonstrated many characteristics of underdeveloped countries but by the final year of the sample, they could be included alongside the more developed European economies. In contrast, France was a developed country over the whole sample period, allowing for comparison between it and its poorer neighbours. Furthermore, Portugal and Spain underwent major political and economic changes in the 1970s and 1980s (the Portuguese revolution in 1974, the change from dictatorial regimes to democracy in the mid 1970s and the integration in the EU in 1986). It was important to take account of these changes in the models. In addition, the concept of neighbourhood and the different development stages attained by Spain and Portugal over time permitted the analysis of interesting aspects of the changes in the competitive behaviour of destinations and their inter-dependencies, by means of dynamic modelling frameworks.

The majority of early empirical studies of tourism demand have relied on static single equation models. These *ad hoc* models are not based on demand theory and, generally, fail to take account of dynamics, nonstationarity of data, and potential simultaneity bias resulting in invalid statistical inference and poor forecasting ability. In chapter 3, the specification deficiencies, theoretical flaws, and technical inadequacies associated with traditional tourism demand modelling were examined and the required solutions were pointed out. In the following chapters, these models' inadequacies were overcome by the construction and estimation of alternative econometric specifications which were shown to be both theoretically consistent and empirically plausible. These specifications are derived from economic theory, incorporate dynamics; consider the interrelationships among destinations and test utility theory hypotheses within system of equations frameworks; question pre-assumed endogenous/exogenous constraints and overcome simultaneity bias using a vector autoregressive approach. The models estimated in this study were subject to rigorous quality scrutiny using both the traditionally quality criteria and more recent econometric methodologies such as structural constancy testing, causality and exogeneity testing and cointegration analysis. The results showed statistical evidence of the existence of a long-run equilibrium relationship between the UK tourism demand and its determinants and, although some of the models are better forecasters than others, all present fairly good predictive quality.

The theoretical, methodological and econometric approaches introduced in each chapter proceed from the simple to the more complex, providing an expanding structure which incorporates, step by step, a number of new concepts and analytical tools towards a comprehensive examination of the UK tourism demand for its neighbouring destination countries. Yet, however wide and detailed the coverage of theoretical and practical issues is in this study, there are important matters that, for one reason or another, had to be ignored or addressed only superficially. The direction of this study's inclusive structure had to be decided at the beginning, and the choice of one direction automatically excludes others which may be as fruitful and interesting as the one explored. Moreover, the wide range of quantitative specifications suitable for approximating consumer demand behaviour are such that a completely comprehensive analysis of them would require more than one thesis on the subject. The lack of consideration of such aspects in this study by no means diminishes their importance but rather enhances the need for their inclusion in future research.

The objective of this concluding chapter is, therefore, to identify the main contributions of previous chapters in enlarging the scope of the theoretical and econometric approaches currently applied in tourism demand analysis, supplying a rigorous and original analysis of the economics of UK tourism demand for its southern neighbours based on the statistical results obtained, and pointing out directions in which some of the issues require further research.

Chapter 2 explains, by means of basic statistics, graphs and tables, the evolution of the UK tourism demand for its southern neighbouring countries over the sample period 1969-1997. The analysis focuses on the significance of UK tourists' preferences relative to other origins for France, Spain and Portugal, and on the importance of these countries relative to other tourism destinations for UK tourists. This analysis provided guidance for the quantitative approach adopted in subsequent chapters since the appropriate characterisation of variables and the definition of the quantitative relationships linking them, are generally grounded on a good understanding of trends, features and facts affecting the behaviour of the time series under study. A thorough examination of the available information is, therefore, of paramount importance in quantitative analysis. Yet, in tourism research contexts, detailed and reliable information in suitable quantities is difficult to gather. The main reason is that the relevant time series of tourism data are generally available on a yearly basis, at very aggregate levels and for relatively short periods of time. This implies that the quantitative approaches describing demand can seldom do other than treat tourism as a commodity "*which is generally undifferentiated other than by destination and which is purchased by consumers who are unspecified other than by nationality*" (Sinclair and Stabler, 1997, p. 215). Yet, the consideration of the microeconomic features of consumers' behaviour and the application of more powerful methodological approaches to tourism demand analysis depend on the availability of data at the household disaggregated level, which can link specific types of demand with specific types of tourists. Future research concerning these issues is of paramount importance for identifying the preferences structure in market niches which differentiate tourists by such criteria as income level, social group, age or gender, and the tourism product by its different components and mixes of characteristics. Nevertheless, aggregate data permit other levels of analysis, which are also of great importance for economic and policy examination as is demonstrated by the empirical findings provided in this thesis.

Chapter 3 begins with an overview of early empirical research on tourism analysis, explaining and critically evaluating the econometric specifications that have been used to estimate demand. The chapter examines the functional forms usually adopted, the type of variables included and the estimation results obtained from different static single equation models. The results show that slightly different specifications can produce considerably different estimates, thereby providing inconsistent results upon which no reliable conclusions can be based. The disparities seem to emerge from the lack of a sound empirical methodology and/or a prevalent theoretical framework within which plausible consumer behaviour hypotheses can be fully integrated and tested. Static single equation models of tourism demand neglect interdependencies among competing destinations, overlook dynamics and lack a sound theoretical basis within which consumers' preferences can be modelled. Empirical specifications constrained by these methodological faults are bound to produce biased and inconsistent estimation results. Perhaps, as Leamer (1987) suggests, 'elimination altogether' of the static single equation model is unavoidable if research is to provide more reliable indications for economic analysis and policy purposes.

The critics of the static single equation modelling argue that the most significant weakness of this approach is its general failure to aptly integrate the dynamics of tourism demand. An appropriate approach within the single equation framework would have to consider a flexible dynamic functional form, involving all the relevant determinants of tourism demand and allowing for the separate estimation of short- and long-run coefficients. The construction of such a dynamic model was implemented in chapter 4, in accordance to the "general to specific" methodology. Thus, in chapter 4, a more reliable single equation approach was considered, by the derivation of ARDL error-correction models which incorporates the dynamic dimension of tourism demand behaviour absent from the static version.

Prior to the estimation of the parsimonious ARDL models, we tested the order of integration of all time series involved and obtained indications that all were nonstationary. Nonstationary time series can give rise to spurious regressions, unless it can be shown that the variables are cointegrated. From the ARDL models of chapter 4, we derived the long-run equilibrium regressions and tested for stationarity of the residuals. The tests indicated stationarity for each demand equation's residuals and, hence, the existence of cointegrated relationships between the UK demand for tourism in each destination and its determinants. The statistical validity of the estimation results obtained from both the long-run and the short-run specifications was further confirmed by a battery of diagnostic tests which provided sufficient evidence to classify the ARDL models as robust, structurally stable and well-defined specifications. Therefore, the ARDL models of chapter 4 proved to be reliable instruments for the investigation of the UK tourism demand and reasonably accurate forecasters of its future behaviour. These findings show that, even within a single equation structure, the incorporation of dynamics into the functional forms can make a huge difference in the reliability of the estimation results, demonstrating that appropriate modelling and testing are fundamental to the accuracy of econometric models.

The dynamic ARDL single equation models have been claimed to be more reliable means of estimating long-run equilibrium relationships than the two-stages method of Engle and Granger (1987). Yet, tourism demand studies applying the former are practically non-existent in the literature. An interesting aspect to be analysed in future research would consist of comparing the statistical robustness and forecasting performance of the ARDL models as derived in chapter 4, with those of single equation specifications derived with the two-stages Engle-Granger method.

Nevertheless, the equation-by-equation nature of these models inhibits the formal modelling of interdependencies among destinations, and precludes testing for restrictions involving cross-equation coefficients. Hence, the models' structure must change to embrace a system of demand equations within which these features can be included. In chapter 5 we examined the UK demand for tourism within a system of equations approach based on the AIDS model of Deaton and Muellbauer (1980). The orthodox version of the AIDS system specifies each destination's share of the UK tourism budget as a function solely of tourism prices and real per capita expenditure. This version, however, was clearly rejected by the data, against the 'unorthodox' AIDS model adopted in chapter 5. This model includes a trend variable and the consideration of a non-constant coefficient for the expenditure explanatory variable. The trend variable accounts for changes in UK tourists' tastes not taken into account by the other explanatory variables. The structural break in the coefficient of the expenditure variable accounts for changes in the allocation of tourists' expenditure which occurred over time, due to factors that modified the political and economic relationships between the origin and destinations and among the destinations themselves. Unlike earlier studies using the orthodox static AIDS approach which rejected utility theory hypotheses and, more often than not, revealed serial correlation and other mis-specification bias, the "unorthodox" AIDS model of chapter 5 was shown to be statistically well-defined, data-coherent and consistent with the utility theory postulates. The AIDS system of chapter 5 has specification features which allow for the incorporation of dynamic-like elements in its equations and, in contrast with the orthodox system used in previous research, this 'seemingly-dynamic' model was consistent with the data and with the utility maximisation assumptions of consumer theory. Hence, this model was considered to provide reliable information about the long-run behaviour of the UK tourism demand for France, Spain and Portugal.

However, the AIDS approach adopted in chapter 5 does not explicitly consider the dynamic features of tourism demand behaviour and cannot account for the short-run correction mechanism which underlies the demand adjustment process. Thus, a clear separation between short- and long-run effects cannot be assessed. Although, in many instances, the focus of interest of economic research is to uncover the long-run structural relationships between dependent and independent variables, the short-run dynamics underlying the steady-state equilibrium model are also of importance, particularly in cases where the short-run effects have relevant magnitudes. Consequently, an explicit dynamic specification was found to be the appropriate means of obtaining reliable information about both the long- and short-run responses of the UK tourism demand to changes in its determinants, and the incorporation of short-run dynamics within the AIDS system was considered to be the logical next step of the investigation.

In chapter 6 we estimated a flexible general dynamic form of the AIDS system. The estimation results showed this model to be data coherent and theoretically consistent, providing empirical evidence of the robustness of the AIDS methodology for undertaking tourism demand analysis in a temporal context. In addition, the dynamic model provided statistically robust evidence on both the capacity of the 'seemingly-dynamic' AIDS of chapter 5 to supply reliable long-run information, and on the inadequacy of the orthodox AIDS and other restricted models nested in the more general specification, to conciliate consistently data and theory within their quantitative formulations. Indeed, the long-run estimates obtained from the general dynamic AIDS model were similar to those obtained from the steady-state version estimated in chapter 5, which seems to support the latter as an adequate model for describing the long-run behaviour of the UK tourism demand. On the other hand, the orthodox AIDS model and specific models nested in the more general form of the dynamic AIDS were unmistakably rejected against the more general form. In addition, the general dynamic AIDS system conformed with all quality criteria and hypotheses tests, providing estimation results which were statistically robust, empirically plausible and theoretically consistent. These results are encouraging and indicate directions for future research. For instance, they show that an appropriate dynamic specification does matter when modelling systems of equations, and can have considerable impact on the results of tests concerning the validity of theoretical hypotheses.

The utility theory constraints in dynamic specifications are generally tested for the long-run coefficients. The motivation for testing theory restrictions, preferably in the steady state, is based on the assumption that, in the short-run, consumers may not have fully adjusted to changing circumstances and, hence, homogeneity and symmetry may not be observed in short-run behaviour. The results obtained from the general dynamic AIDS model suggest that tourists adjust their behaviour with a lag. Hence, homogeneity and symmetry are not expected to hold in the short-run. However, when tested, these hypotheses could not be rejected, meaning that the utility maximisation postulates are observed both in the long- and in the short-run. This result is not completely unexpected. Given that the estimates of the general dynamic model suggested that UK tourists adjust very fast to changes in their demand determinants, non-significant and/or irrelevant magnitudes for the short-run coefficients should be expected. Indeed, that was the general indication of the estimates provided by the model. With insignificant short-run estimates, the statistical process by which constraints are imposed on the coefficients generally leads to an 'easier' non-rejection (a possible under-rejection) of the null. Therefore, the faster consumers adjust their demand behaviour, the less significant short-run coefficients are, and the likelier is the non-rejection of utility theory postulates imposed on them. This statement requires, of course, further empirical support, which can only be delivered in the context of future research.

The similarity of the long-run estimates obtained from the general dynamic AIDS model with those obtained from the steady-state version estimated in chapter 5 and the statistical robustness of both models, seemed to support the existence of a structural equilibrium relationship between the UK tourism demand and its determinants. Yet, there were still theoretical and empirical issues that had to be addressed to endorse the AIDS model as a quality specification. One important issue regards the spurious regression problem. The AIDS system includes nonstationary time series, so that its estimation results can be spurious unless the variables are cointegrated. Other issues are related with the *a priori* division of endogenous/exogenous variables assumed by the AIDS approach. Since the AIDS system assumes that the current levels of its explanatory variables are exogenously determined, unless this is the case, mis-specification bias can occur, invalidating its estimation results and statistical inference.

Therefore, when data series are nonstationary, the estimation of econometric models which regress endogenous variables on several assumed exogenous variables, without sanctioning their statistical validity with cointegration analysis may be questionable. Hence, given non-stationary data and potential feedback effects, an efficient econometric approach for estimating long-run relationship(s), is a system of equations permitting all variables to be treated as endogenous and appropriate cointegration analysis. The vector autoregressive approach is an appropriate solution to overcome these problems.

The vector autoregressive approach establishes a model specification which is a valid alternative to both the dynamic single equation and the traditional structural multi-equation approaches. The Johansen (1988) procedure permits cointegration analysis and establishes an efficient method of identifying the number of cointegrating vectors and estimating the structural parameters. This methodology is applied to system of equations and avoids the problems of spurious regression, simultaneous equation bias, *a priori* division of endogenous-exogenous variables and unfounded imposition of zero restrictions.

The AIDS model specified in chapter 5 is a system of equations which incorporates nonstationary data series and assumes exogeneity for their right-hand side variables. The implications of these features could undermine the statistical validity of this model if no cointegrating relationships were found linking its variables. Hence, the objective of chapter 7 was to find empirical evidence of cointegration in order to support, or otherwise, the inference procedures and estimation results obtained with the AIDS specification. For this purpose, we specified a reduced-form unrestricted VAR system including the same variables as the AIDS model, and used pertinent procedures to establish the lag-length, deterministic components and the endogenous/exogenous division of the. Once the appropriate form of the VAR was in place, the Johansen cointegrated rank test was used to determine the number of cointegrating vectors. The test provided statistical support for the economic principles underlying the AIDS model, which predicts the existence of exactly (n-1) long-run (cointegrated) relationships in a system of n expenditure share equations. In addition, the structural parameters of the cointegrated vectors in the VAR model were exactly-identified with restrictions matching those of the normalisation process used to identify the tourism share equations of an AIDS system. Finally, the resulting structural form of the cointegrated VAR, identical to that of the AIDS specification, was tested under additional over-identifying restrictions such as homogeneity, symmetry and null cross-price effects and these hypotheses were not rejected. In addition, the elasticities estimates obtained with the structural coefficients of the cointegrated VAR under the full set of theoretical restrictions proved to be fairly similar to the corresponding elasticities obtained from the AIDS models estimated previously. Consequently, strong evidence was obtained to support the AIDS model's ability to portray accurately the theoretical predictions underlying the rationale of the UK tourism demand behaviour and the destinations' competitive conduct.

The theoretical and empirical consistency of the cointegrated VAR under the full set of restrictions implied that the model's predictive ability would be good. This was, indeed, the case as the quality criteria measuring the accuracy of this model's forecasts indicated it as the most precise forecasting device, among all VAR specifications used in chapter 7. The criteria suggested that the general reduced-form (unrestricted) VAR was the next best forecasting model. This finding gave empirical support to the claimed competence of VAR models for forecasting purposes, and endorsed the qualities of modelling simplicity and estimation ease of the "pure" unrestricted VAR. Given these results, we conclude that if the main goal of a multivariate modelling exercise is only to provide fairly accurate forecasts, rather than to estimate the structural relationships among the variables, the reduced-form VAR model may constitute the best choice.

Given the results obtained in chapter 7, we believe we have contributed enough empirical evidence for, on the one hand, considering the AIDS approach as a theoretically consistent and statistically robust means of producing valid and reliable estimates of the long-run equilibrium parameters underlying the relationships between destinations' tourism shares and its determinants and, on the other hand, for confirming the competence of the VAR model, either in its more general unrestricted form or under the full set of theoretical restrictions, to provide accurate forecasts of the destinations' tourism shares.

The forecasting ability of econometric models is a fundamental aspect of their quality evaluation. Therefore, chapter 8 evaluates the models of chapters 4 to 7, by comparing their relative forecasting ability over the out-of-sample period 1994-1997. The results showed that most of the specifications, namely the fully restricted VAR specification, the ARDL model and the restricted 'unorthodox' AIDS system, produce forecasts with average absolute errors below 1.5 percentage points, which can be viewed as accurate. For all destinations, the fully restricted VAR produces forecasts with average absolute errors ranging between 0.6 and 1 percentage points; the ARDL models, between 0.7 and 1.3 percentage points, and the restricted 'unorthodox' AIDS model, between 1.1 and 1.5 percentage points. Only the reduced-form VAR and the unrestricted AIDS models produce errors outside this range, for one or more destinations. The results indicate that the fully restricted models generate more precise forecasts than their unrestricted or partially restricted counterparts, dynamic models perform better than static ones, and the cointegrated VAR under the full set of restrictions is the out-performs all the other models. The conclusion drawn from these results is, that appropriately specified econometric models are excellent forecasters. Moreover, claims that the performance of univariate models is superior to econometric models are unlikely to hold, if these models are tested against well-defined, robust and consistent econometric specifications.

A good econometric model provides a formal quantitative framework which can be validly used to understand economic activities, interpret economic relationships, test economic theories, evaluate economic policy and predict economic behaviour. The main conclusion that can be drawn from the extensive modelling exercise carried out in this study is that reliable and consistent results can be derived from appropriately specified models which are based on theory postulates and integrate knowledge of specific features, facts and events that affected the economic relationships between origin and destinations, and among destinations themselves. The results obtained from the different specifications of chapters 4 to 7 are consistent across the models, in contrast to the *ad hoc* single equation approach of chapter 3. The consistency of the results provided by the former derives from their foundation of pertinent economic principles and sound econometric methodologies. Thus, these models constitute the necessary and reliable basis for quantifying, understanding and predicting the long-run behaviour of the UK tourism demand.

Each econometric specification provides information about particular aspects of demand behaviour but, inevitably, leaves others unattended. For instance, the ARDL single equation models permit the incorporation of dynamics, cointegration analysis and give direct information about own-price and income elasticities. With this specification, we were able to assess the short- and long-run impacts on UK tourism demand induced by changes in such variables as destinations' prices, UK tourism expenditure in neighbouring destinations and UK per capita income. Yet, its equation-by-equation structure does not allow for the incorporation of theoretical constraints imposed on cross-equation parameters.

The AIDS system of equations allows for the incorporation of these constraints as well as for the modelling of a dynamic structure. The empirical information it provides concerns the own- and cross-price elasticities and, instead of income elasticity, it supplies estimates of tourism expenditure elasticities. The specific form of the models adopted in chapters 5 and 6 includes a structural break in the expenditure variable coefficient, which allowed for the examination of the changing sensitivity, over time, of the UK tourism demand elasticities. Nevertheless, the assumptions underlying the AIDS model include *a priori* division of exogenous/endogenous variables, which may not hold, and variables that are nonstationary. The validation of its results requires that the variables in the model are cointegrated, forming a number of cointegrating vectors as predicted in the AIDS system; that is, a system of n equations must have n-1 cointegrating vectors. In addition, the identification of their structure must match that underlying the AIDS equations.

These issues can be examined using the VAR approach, which treats all variables as endogenous and allows for identification of the cointegrating vectors and, hence, for the long-run structural coefficients which are the main interest of the analysis. The VAR approach can sanction the validity of the AIDS model and provides the same type of information. However, the modelling, testing and inference procedures involved in the

VAR approach are more complex than those used in the AIDS models. Nevertheless, it should be noted that the AIDS system can only be sanctioned through cointegration analysis which requires the procedures undertaken to estimate and test a VAR model.

The main conclusion is, therefore, that a comprehensive analysis of tourism demand must include a wide range of appropriately specified models which can provide different insights about the main features of interest and supply information about the consistency and validity of the results obtained from the different approaches. If the models comply with all quality criteria and their results are consistent, we can be confident of their accuracy and validity for economic analysis and policy purposes. This was the case for the empirical findings in this study. Hence, we can use the results to draw conclusions about the UK tourism demand behaviour for France, Spain and Portugal.

The results show that the long-run expenditure elasticities, although close to unity for all destination countries, tend to increase for France and Portugal and to decrease for Spain, in the more recent years of the sample. In contrast, Spain displays the smallest short-run expenditure elasticity compared with those of its neighbours. This may indicate that Spain has a comparative advantage, relative to its neighbours, in terms of short-run increases in the UK tourism budget but seems to be losing ground to France and Portugal in the long-run preferences of UK tourists.

In the more recent years of the sample, the estimates of the sensitivity of the UK demand to own-price changes provided by the static AIDS model and by the fully restricted cointegrated VAR, are similar for France and Portugal, indicating a long-run value close to –2. However, in the equation for Spain, the former model indicates an own-price elasticity close to –1, while the latter indicates a value close to –2. Given that the VAR model complies with all quality criteria as well as being the best forecaster, we assume its estimation results as more reliable. The information given for the UK demand reaction to short-run own-price changes indicates the demand for Spain to be less sensitive than that for its neighbours. Hence, Spain seems to have a comparative advantage relative to its neighbours in the context of price increases, both in the long- and in the short-run. However, as the UK demand sensitivity to price-changes in Portugal and France shows a tendency to diminish over time, while that to price-changes in Spain seems to be increasing, the comparative advantage of Spain may be short-lived.

The information given by the AIDS and VAR models concerning the competitive conduct of the destinations is remarkably consistent. The long-run estimates support our conjecture that the geographically closer destinations are, the likelier is to find evidence of significant competitive behaviour between them. Indeed, in all cases, the results indicate that France and Portugal are competitive destinations relative to Spain in UK tourists' preferences, but no significant link was found between France and Portugal. However, the short-run estimates suggest, in all cases, no significant reaction of the UK tourism demand for one destination to price-changes in its neighbours. This may indicate that in the short-run, decreasing prices in one destination are not likely to affect significantly its level of UK demand and, therefore, seems to be a means of producing net gains from UK tourists. The results also show that, in the long-run, the UK demand for Portugal or France is more sensitive to price changes in Spain, than the demand for Spain is to price changes in Portugal or France. This indicates that the UK demand for Spain may be considered relatively insensitive to long-term 'cut-throat' price-decreasing policies from its neighbours. Yet again, the results indicate, for the more recent years of the sample, a diminishing sensitivity of the UK demand for France and Portugal to price changes in Spain, and an increasing sensitivity of the demand for Spain to price changes in its neighbours. Therefore, once again, the UK tourists' secular preference for Spain seems to be changing, slowly but progressively, towards favouring its neighbouring competitors.

Although the context of the analysis carried out in this study concerns the tourism demand of one origin for three destination countries, any of the models specified in chapters 4 to 7 can be readily extended to a larger number of destinations and origins without any loss of generality. Furthermore, the AIDS and VAR systems of equations can be applied in contexts other than the demand for an undifferentiated tourism product at the national level. Indeed, we can use these models to explain and predict demand for particular regions within one country, specific resorts within one region and even individual items, such as type of accommodation or tourism attractions, within one local area.

As long as tourists' spending on diversified products can be divided into expenditure shares, the relationships between the dependent variable and its determinants can be modelled as a system of share equations within the AIDS framework. The VAR models are even more general, as they do not necessitate the use of share equations but accept any other form of measuring tourism demand. The only problem with VAR specifications and AIDS models is that they are data-intensive models and, in tourism contexts, data are relatively scarce.

The ARDL models, as single equation specifications are less data-intensive than the former and do not erode degrees of freedom to the same extent as the other models, when additional regressors are included. Hence, they can be useful for incorporating other variables of interest, such as transport costs or advertising expenditure, when relevant, without losing precision in their estimates. All these extensions constitute interesting issues for future research.

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