Testing alternative dynamic systems for modelling tourism demand

MARIA M. DE MELLO
Centro de Estudos de Economia Industrial, do Trabalho e da Empresa (CETE), Universidade do Porto, Faculdade de Economia do Porto, Rua Dr. Roberto Frias, 4200-464 Porto, Portugal. E-mail: mmello@fep.up.pt.

NATÉRCIA FORTUNA
Centro de Estudos Macroeconómicos e Previsão (CEMPRE), Universidade do Porto.

This paper presents an empirical study of tourism demand dynamics and identifies areas in which the scrutiny of relationships between theoretical and empirical considerations is likely to produce new insights. A flexible general form of a Dynamic Almost Ideal Demand System (DAIDS) is derived to analyse UK tourism demand for the neighbouring destinations of Portugal, Spain and France during 1969–97. Nested within the general dynamic structure are Deaton and Muellbauer’s static AIDS model itself, the partial adjustment model and the auto-regressive distributed lag model, which are tested against the general dynamic alternative. The empirical results obtained show that DAIDS is a data-coherent and theoretically consistent model, providing evidence of the robustness of this methodology for tourism demand analysis in a temporal context. Moreover, the dynamic model offers statistically strong evidence of the inadequacy of the orthodox static AIDS and other restricted models for the consistent reconciliation of data and theory within their formulations. Estimates for tourism price and expenditure elasticities are obtained, permitting a comparative analysis of the relative magnitudes and statistical relevance of the long-run and short-run sensitivity of UK tourism demand to changes in its determinants.

Keywords: tourism demand; dynamic almost ideal system; partial adjustment system; auto-regressive distributed lag system

JEL classification: C52, D12

The authors thank participants at the conference on Theoretical Advances in Tourism Economics for constructive comments. They are also grateful to two anonymous referees for their useful suggestions. The usual disclaimer applies.

CETE and CEMPRE are supported by Fundação para a Ciência a Tecnologia, Portugal, through the Programa Operacional Ciência, Tecnologia e Inovação (POCTI) of the Quadro Comunitário de Apoio III, which is financed by FEDER and Portuguese funds.
In the study of tourism demand analysis, early research efforts (White, 1985; O’Hagan and Harrison, 1984; Syriopoulos and Sinclair, 1993; Papatheodorou, 1999; De Mello et al., 2002) concentrate on static specifications based on Deaton and Muellbauer’s (1980a; 1980b) Almost Ideal Demand System (AIDS). In purely static specifications, such as the orthodox AIDS approach, consumers are assumed to adjust perfectly and instantaneously to changes in their demand determinants. Yet if features such as habit persistence, unstable preferences, adjustment costs or imperfect information, prevent consumers from adjusting fully every period, an explicit dynamic structure is required to explain demand behaviour and to account for the short-run adjustment process.

In the context of demand analysis, it is realistic to consider that past behaviour alters preferences and, consequently, affects current demand. 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 normally represented by lagged variables. However, there may be little to be learnt from simply adding lagged explanatory variables to an otherwise static model, since the resulting specification may be acceptable only under implausible behavioural hypotheses.

Unless empirical models are appropriately specified and the implications of general theoretical principles are fully integrated, invalid statistical inference may result, inducing research to proceed in less valuable directions.

Although the importance of including explicit dynamic adjustments in demand analysis has generally been recognized, specific research in tourism demand using dynamic systems is not abundant, and only a few authors have addressed these issues in empirical studies. These studies can roughly be divided into two groups according to specification type: vector autoregressive models (for example, Song et al., 2003; Witt et al., 2003; De Mello and Nell, 2005) and dynamic specifications based on Deaton and Muellbauer’s AIDS model (for example, Lyssiotou, 2000; De Mello, 2001; Durbarry and Sinclair, 2003; Li et al., 2004).

The first efforts to derive formally a theoretically consistent dynamic form of Deaton and Muellbauer’s AIDS model for application to tourism demand analysis were made by Lyssiotou (2000) and De Mello (2001). Lyssiotou (2000) introduces expectations within the dynamics of an AIDS model, but does not fully examine the short-run mechanism underlying the adjustment process of tourism demand to its long-run equilibrium levels. De Mello (2001) derives an error correction dynamic form of the AIDS model that explains short-run and long-run tourism demand and accurately forecasts its levels up to four periods ahead. Durbarry and Sinclair (2003) apply this same model of De Mello (2001) to a different set of origins and destinations. They claim that the short-run variables are statistically irrelevant and omit them from the model; thus there is no analysis of the short-run effects, which are the main concern of any dynamic model. The work of Li et al. (2004) constitutes a valuable contribution to the literature on the error correction AIDS approach applied to tourism demand analysis, but the model’s derivation seems to imply that the parameter measuring the adjustment velocity will have ambiguous estimates. Adjustment velocities should be positive and less than one. Adjustment velocities close to unity may indicate the adequacy of a static model. The slower the adjustment process is, the more important short-run dynamic adjustments are. The analysis...
of this important aspect seems to have been omitted. In addition, there appear to be different adjustment velocities across equations, which implies that the parameter estimates are not invariant with the omission of one particular equation (through the adding up propriety). Therefore, if another equation is omitted – say the one relating to Portugal, which represents only 4% of the UK budget, instead of the one relating to ‘other destinations’, representing 31.4% – the estimates could be different, unfolding a set of potentially different conclusions for the same data and model.

Knowing how the allocation of tourism expenditure evolves over time and how tourists adjust their demand behaviour to achieve equilibrium is of considerable interest for tourism business and policy making. The objectives of this paper, therefore, are to contribute an empirical study of tourism demand dynamics and to point out areas in which the scrutiny of relationships between theoretical and empirical considerations is likely to produce new insights. For instance, modelling groups of destinations that allow for similar adjustment velocities within the same system avoids heterogeneity problems that can be difficult to overcome in such specifications. Moreover, identical adjustment velocities allow for invariant results whichever equation one chooses to omit. It is also important to include adequate dummy variables to model structural breaks and historical one-off events. These allow for specifications that, through extensive testing, prove statistically robust and theory-consistent. Their omission can result in misspecified models, with all the well-known undesirable consequences. Specifications that pass all quality tests, outperform competing models and supply plausible and significant estimates to the relevant parameters constitute, in general, reliable means of explaining and forecasting the phenomena under analysis. These are the characteristics that we found in the model derived, tested and estimated in this paper.

The structure of the paper is as follows. In the next section, a flexible general form of a Dynamic Almost Ideal Demand System (DAIDS) is derived. Nested within this general form are several alternative specifications, such as the static orthodox AIDS model itself, the partial adjustment (PA) and the auto-regressive distributed lag (ARDL) models. These models are used to analyse UK tourism demand for Portugal, Spain and France during 1969–97. The UK is a major tourist origin country that is of particular importance for the three neighbouring destinations under analysis. Spain and Portugal are interesting cases in that they experienced considerable economic development and major political change during the sample period, in contrast to France, which was already well-developed and remained politically stable over the same period. These features enable analysis of significant changes in the interdependencies and competitive behaviour of the destinations over the sample period. In the subsequent section, we implement statistical tests to assess the consistency between the alternative models and the principles of consumer demand theory. The results show that the restricted models nested within the general specification are theoretically inconsistent, while the dynamic system DAIDS proves to be appropriate and statistically robust for conducting tourism demand analysis. We then present the estimation results for long-run and short-run tourism price and budget elasticities. These permit a comparative analysis of the relative magnitude and relevance of the long-run and short-run sensitivity of UK tourism demand to changes in its determinants. Finally, we offer our conclusions.
Dynamic AIDS modelling of UK tourism demand

In what follows, a flexible dynamic structure for the AIDS model is derived, based on the work of Anderson and Blundell (1983; 1984). Nested within the general structure are the static model itself and dynamic specifications such as the PA and the ARDL models, which can be tested against the general DAIDS.

Consider the orthodox static AIDS model described in Appendix 1. For simplicity, we rename the variables $w_i = W_i$, $\ln p_j = P_j$, $i, j = 1, \ldots, n$, and $\ln x/P^* = E$. Hence, Equation A9 in Appendix 1 is written as:

$$Wi = \alpha^* + \sum_{j=1}^{n} \gamma_j P_j + \beta_i E, \quad t = 1, \ldots, T,$$

(1)

where $W_i$ represents the expenditure share allocated to the $i^{th}$ destination by UK tourists, $P_j$ stands for the effective price of tourism in destination $j$, and $E$ represents the UK real per capita tourism budget allocated to all destinations under analysis.

Consider Equation (1) as the appropriate choice for the steady-state structure of the following general dynamic stochastic specification:

$$\Delta Wi = \sum_{j=1}^{n} \gamma_j^{s} \Delta P_j + \beta_i^{s} \Delta E + \lambda_i \left( \alpha + \sum_{j=1}^{n} \gamma_j^{l} P_{j,t-1} + \beta_i^{l} E_{t-1} - Wi_{t-1} \right) + ui_t,$$

(2)

where $\lambda_i$ is the adjustment coefficient of the $i^{th}$ equation, $\Delta$ is the first difference operator, subscript $t-1$ indicates the variables’ lagged values, and $ui_t$ is the $i^{th}$ disturbance term assumed to be characterized by a singular, independent and identical distribution over time. The parameters with superscripts $S$ and $L$ can be interpreted, respectively, as the short-run and long-run responses of the dependent variable to changes in its determinants. Equation (2) assumes that, to maintain the steady-state relationship (Equation 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 disequilibrium observed in the previous period.

Although Equation (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 a ‘shaky’ statistical inference procedure. Hence, a sequence of tests is performed to find the most restrictive specification that is consistent with the particular set of $T$ observations available. Examples of such specifications are provided below.

Consider the following equivalent form of the general Equation (2) where $\lambda_i = \lambda$:

$$\Delta Wi = \gamma_i^{s} \Delta P_i + \ldots + \gamma_i^{s} \Delta P_n + \beta_i^{s} \Delta E + +\lambda \alpha + \lambda \gamma_i^{l} P_{i,t-1} + \ldots + \lambda \gamma_i^{l} P_{n,t-1} + \lambda \beta_i^{l} E_{t-1} - \lambda W_{i,t-1} + ui_t,$$

(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:
Dynamic systems for modelling tourism demand

which is an ARDL model of order \( m \). This general form may be reduced to a parsimonious one by applying several criteria (Hendry and Richard, 1983) which include the definition of \( m \) as a small number. The general form of model (3) can be reduced to an ARDL form as described, under the null hypothesis:

\[
H_0 : \lambda \alpha_i = 0 \land \gamma^L_{ij} = \gamma^R_{ij} \land \beta^L_i = \beta^S_i = \beta_i, \quad \text{for all } i, j.
\]

If \( H_0 \) cannot be rejected, then the model reduces to

\[
W_i = \gamma^L_{1i}(P_{1i} - (1 - \lambda)P_{1i-1}) + \ldots + \gamma^L_{ni}(P_{ni} - (1 - \lambda)P_{ni-1}) + \beta_i(E_i - (1 - \lambda)E_{i-1}) + (1 - \lambda)W_{i-1} + u_i,
\]

which is a first-order ARDL model.

Partial adjustment (PA) model

Consider the flexible accelerator model of economic theory, which assumes that the equilibrium level of a dependent variable, say \( Y^* \), is a linear function of an explanatory variable, say \( X \), such that:

\[
Y^* = \beta_0 + \beta_1 X + u_i.
\]

The partial adjustment hypothesis postulates:

\[
Y_i = \delta Y^*_i + (1 - \delta)Y_{i-1}.
\]

Substituting (5) in (6) leads to:

\[
Y_i = \delta \beta_0 + \delta \beta_1 X_i + (1 - \delta)Y_{i-1} + \delta u_i.
\]

Considering now model (3), the null hypothesis to be tested is:

\[
H_0 : \gamma^S_{ij} = \lambda \gamma^L_{ij} = \gamma^R_{ij} \land \beta^S_i = \lambda \beta^L_i = \beta_i, \quad \text{for all } i, j.
\]

If \( H_0 \) cannot be rejected, then the model reduces to:

\[
W_i = \lambda \alpha_i + \gamma^L_{1i} P_{1i} + \ldots + \gamma^L_{ni} P_{ni} + \beta_i E_i + (1 - \lambda)W_{i-1} + u_i,
\]

which is a partial adjustment model similar to the one described in Equation (7).

Static AIDS model

To test for the static form nested within Equation (3) the null hypothesis is:

\[
H_0 : \gamma^S_{ij} = \lambda \gamma^L_{ij} = \gamma^R_{ij} \land \beta^S_i = \lambda \beta^L_i = \beta_i \land \lambda = 1, \quad \text{for all } i, j.
\]

If \( H_0 \) cannot be rejected, then the model reduces to:

\[
W_i = \alpha_i + \gamma^L_{1i} P_{1i} + \ldots + \gamma^L_{ni} P_{ni} + \beta_i E_i + u_i,
\]

which is the steady-state orthodox AIDS model.
Testing the consistency of alternative dynamic models

The methodological strategy in this section develops as follows. First, we define a general dynamic structure for the expenditure adjustment process of UK tourism consumers. Second, we test for more restrictive specifications believed to be consistent with the data. Finally, we test the utility maximization restrictions on specifications not rejected by the data.

Considering the specific demand context under analysis, the dependent variable of the DAIDS model in Equation (3) represents the UK tourism demand share $W_i$ of destination $i$, where $i = P$ (Portugal), $S$ (Spain), $F$ (France). Its explanatory variables include the effective tourism price in Portugal, Spain and France (respectively, $PP$, $PS$ and $PF$), and the UK real per capita tourism expenditure $E$ such that:

$$\Delta W_i = \lambda \alpha_i + \lambda \gamma^P_i PP_{t-1} + \lambda \gamma^S_i PS_{t-1} + \lambda \gamma^F_i PF_{t-1} + \lambda \beta_i E_{t-1} - \lambda W_{i,t-1} +$$

$$+ \gamma^P_i \Delta PP_i + \gamma^S_i \Delta PS_i + \gamma^F_i \Delta PF_i + \beta_i \Delta E_i + u_i.$$  \hspace{1cm} (10)

Model (10) is the general dynamic structure DAIDS on which appropriate restrictions described earlier are imposed to obtain Equations (4), (8) and (9) – the ARDL, PA and static AIDS models, respectively. The compatibility of these models with the data is tested below. If the models are compatible, they are further subjected to utility theory constraints and tested. The non-rejection of these constraints indicates theory-consistent models.

The orthodox static AIDS, PA and ARDL models are rejected against the general DAIDS model at the 1% significance level, with the Wald test statistic values, respectively, $\chi^2 (10) = 46.10$, $\chi^2 (9) = 22.63$ and $\chi^2 (11) = 103.49$. Only the dynamic structure (10) reveals itself fully compatible with the data.

![Real expenditure in 1969–1997](image)

**Figure 1.** UK real per capita tourism budget allocated to Portugal, Spain and France.
and with the assumptions of consumer demand theory. Therefore, the remainder of this section focuses on this model.

As illustrated in Figure 1, the data analysis indicates the possible presence of a structural break in the coefficient of the real expenditure variable $E$, dividing the sample into two sub-periods: 1969–79 and 1980–97.

The data analysis (see Figure 2) also indicates the possible relevance of an intercept dummy $D$ in the share equations for Portugal, Spain and France, over the period 1974–81. This information is integrated in specification (10) in the following way. We assume that the structural break is relevant only in the long run. Hence, we add to Equation (10) a new variable $SE_{t-1}$, constructed using a step dummy variable $S$ for the year 1979. $S$ is then multiplied by $E_{t-1}$, giving rise to variable $SE_{t-1}$ which is included in (10). We assume that the intercept dummy $D$ may have significant effects both in the long run and in the short run. Therefore, we add to (10) variables $D_{t-1}$ and $\Delta D_t$.

Assessment of the statistical significance of the added variables $SE_{t-1}$, $D_{t-1}$ and $\Delta D_t$ is undertaken by testing an unrestricted $U$ model, which includes these variables, against a restricted $R$ model, which excludes them, using the Wald statistic. Table 1 presents the text results ($P$-values in brackets), which indicate that $D$ is relevant only in the long run, since the non-significance of $\Delta D_t$ cannot
Table 1. Tests for the individual and joint significance of variables $SE_{t-1}$, $D_{t-1}$ and $\Delta D_r$.

| Hypotheses under test | Wald statistic |
|------------------------|---------------|
| $H_0$: Non-significance of $\Delta D_r$ | $\chi^2 (2) = 1.78 (0.41)$ |
| $U$: $PP_{t-1}PS_{t-1}PF_{t-1}E_{t-1}W_{t-1} \Delta PP \Delta PS \Delta PF \Delta E$, $SE_{t-1} D_{t-1} \Delta D_r$ | Not rejected |
| $R$: $PP_{t-1}PS_{t-1}PF_{t-1}E_{t-1}W_{t-1} \Delta PP \Delta PS \Delta PF \Delta E$, $SE_{t-1} D_{t-1} \Delta D_r$ | $\chi^2 (4) = 18.06 (0.00)$ |

be rejected. The joint significance of $SE_{t-1}$ and $D_{t-1}$ is not rejected. Therefore, these variables are included in the model.

There is no evident reason to believe that different velocities of adjustment should exist among the share equations. Indeed, UK tourists have fairly similar information about the destinations under analysis, implying that tourists are expected to adjust with similar speed to changes in their demand determinants. Consequently, the assumption of equal adjustment coefficients for all share equations in the dynamic system is also tested. The null for this test restricts the adjustment coefficients $\lambda_i$ to be equal across equations, so that $\lambda_i = \lambda$ for all $i$. The Wald statistic value $\chi^2 (1) = 0.03$, does not reject this hypothesis. Hence, the constraint of equal adjustment coefficients across equations is integrated in all subsequent restricted models derived from the general DAIDS.

Given the above considerations, the $i^{th}$ equation of the general DAIDS model of UK tourism demand for destination $i$ is:

$$
\Delta W_i = a_{i1} + a_{i2} PP_{t-1} + a_{i3} PS_{t-1} + a_{i4} PF_{t-1} + a_{i5} FE_{t-1} + a_{i6} SE_{t-1} + a_{i7} D_{t-1} - \\
- \lambda W_{i-1} + a_{i8} \Delta PP + a_{i9} \Delta PS + a_{i10} \Delta PF + a_{i11} \Delta E_i + \epsilon_i.
$$

(11)

The structural break that divides the expenditure observations into two sub-periods, 1969–79 and 1980–97 is included in Equation (11) by the use of two dummy variables, $F$ and $S$, which assume the 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 $SE_{t-1}$.

The general DAIDS (11) is now tested against further constrained models under the restrictions of homogeneity, symmetry and null cross-price effects between the equations for Portugal and France, in both the long run and short run. The results are presented in Table 2, and confirm that none of the hypotheses is rejected by the data. Hence, long-run and short-run homogeneity and symmetry cannot be rejected at the 5% significance level. This is also true for the hypothesis of long-run and short-run null cross-price effects between the equations for Portugal and France. These results have important implications for the modelling and prediction of consumer behaviour. They suggest that knowledge of the way 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, to obtain comprehensive information on the error-correction mechanism triggering the process of adjustment, it may not be sufficient simply to introduce trend factors in the usual static AIDS.
Dynamic systems for modelling tourism demand

Table 2. Tests for utility theory restrictions on the long-run and short-run coefficients.

| Hypotheses under test                                                                 | Wald statistic |
|---------------------------------------------------------------------------------------|----------------|
| $H_0$: Long-run homogeneity                                                           | $\chi^2 (3) = 0.54 (0.91)$ Not rejected |
| $H_0$: Long-run homogeneity and symmetry                                              | $\chi^2 (4) = 2.16 (0.71)$ Not rejected |
| $H_0$: Long-run homogeneity, symmetry and null cross-price effects between the share equations of Portugal and France | $\chi^2 (5) = 2.56 (0.77)$ Not rejected |
| $H_0$: Long-run homogeneity, symmetry and null cross-price effects and short-run homogeneity | $\chi^2 (7) = 8.72 (0.27)$ Not rejected |
| Long-run homogeneity, symmetry and null cross-price effects and short-run homogeneity, symmetry and null cross-price effects | $\chi^2 (8) = 10.35 (0.24)$ Not rejected |
| Long-run homogeneity, symmetry and null cross-price effects and short-run homogeneity, symmetry and null cross-price effects | $\chi^2 (9) = 10.59 (0.31)$ Not rejected |

Note: The constraint of equal adjustment coefficients across equations is integrated in all subsequent restricted models derived from the general structure. Therefore, in all the hypotheses tested this constraint holds previously. Consequently, the number of degrees of freedom for the statistic includes this first restriction.

formulations. Nor does it appear sufficient to choose dynamic models representing specific theories of short-run correction. A more general dynamic structure seems to be required to match data and theory in a consistent way.

Empirical results and their interpretation

Estimation results and diagnostic tests

Table 3 presents the estimation results, obtained with the SUR method, for the following three models: the general DAIDS model (11), under the sole restriction of equal adjustment coefficients; the same model under the additional restrictions of homogeneity $H$, symmetry $S$ and null cross-price effects between the equations of Portugal and France $N$, applied only in the long run $L$, denoted as $HSN^L$; and the same model under the restrictions of homogeneity, symmetry and null cross-price effects, applied both in the long run and short run $LS$, denoted as $HSN^{LS}$. The table shows the coefficient estimates (asymptotic $t$-values in brackets) for the share equations of Portugal (PT), Spain (SP) and France (FR). Goodness-of-fit indicators, such as the residual sum of squares (RSS), equation log-likelihood (ELL) and system log-likelihood (SLL) values, are also presented.

Table 4 shows 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 DAIDS model ($P$-values in brackets). The diagnostic tests indicate that the equations of the DAIDS model are well-defined and statistically robust.

Interpretation of elasticity estimates

A thorough analysis of the dynamic coefficients, accompanied by a comparison of long-run and short-run estimates is worthwhile. This analysis is carried out...
| Explanatory variable | Unrestricted | \( HSN^L \) | \( HSN^S \) |
|----------------------|--------------|-------------|-------------|
|                      | PT           | SP          | FR          | PT           | SP          | FR          | PT           | SP          | FR          |
| Constant             | 0.0995       | 0.2330      | 0.4599      | 0.0695       | 0.2618      | 0.4265      | 0.0671       | 0.2821      | 0.4299      |
|                      | (4.50)*      | (2.66)*     | (7.41)*     | (5.41)*      | (3.25)*     | (7.94)*     | (5.93)*      | (4.07)*     | (7.49)*     |
| \( PP_{t-1} \)       | -0.0238      | -0.0479     | 0.0717      | -0.0494      | 0.0494      | 0.0000      | -0.0472      | 0.0472      | 0.0000      |
|                      | (-0.82)      | (-0.63)     | (0.97)      | (2.43)*      | (2.43)*     | (---)       | (-2.52)*     | (2.52)*     | (---)       |
| \( PS_{t-1} \)       | 0.0899       | -0.5036     | 0.4137      | 0.0494       | -0.4486     | 0.3992      | 0.0472       | -0.4545     | 0.4074      |
|                      | (2.69)*      | (-5.73)*    | (5.16)*     | (2.43)*      | (-5.72)*    | (5.38)*     | (2.52)*      | (-5.06)*    | (4.79)*     |
| \( PF_{t-1} \)       | -0.0568      | 0.5450      | -0.4882     | 0.0000       | 0.3992      | -0.3992     | 0.0000       | 0.4074      | -0.4074     |
|                      | (-1.22)      | (4.08)*     | (-3.79)*    | (---)        | (5.38)*     | (-5.39)*    | (---)        | (4.79)*     | (-4.79)*    |
| \( FE_{t-1} \)       | -0.0163      | 0.0644      | -0.0481     | -0.0038      | 0.0359      | -0.0322     | -0.0034      | 0.0476      | -0.044      |
|                      | (-1.47)      | (2.23)*     | (-1.81)*    | (-0.50)      | (1.63)      | (-1.60)     | (-0.46)      | (2.60)*     | (-2.71)*    |
| \( SE_{t-1} \)       | -0.0039      | 0.0295      | -0.0256     | 0.0018       | 0.0188      | -0.0207     | 0.0026       | 0.0177      | -0.0204     |
|                      | (-0.82)      | (2.41)*     | (-2.30)*    | (0.67)       | (1.81)*     | (-2.13)*    | (1.04)       | (1.96)*     | (-2.38)*    |
| \( D_{t-1} \)        | -0.0152      | -0.0172     | 0.0324      | -0.0145      | -0.0118     | 0.0263      | -0.0136      | -0.0271     | 0.0407      |
|                      | (-2.34)*     | (-1.15)     | (2.24)*     | (-2.37)*     | (-0.87)     | (2.17)*     | (-2.28)*     | (-1.77)*    | (2.92)*     |
| \( WI_{t-1} \)       | 0.792        | 0.792       | 0.792       | 0.758        | 0.758       | 0.758       | 0.779        | 0.779       | 0.779       |
|                      | (7.45)*      | (7.45)*     | (7.45)*     | (7.94)*      | (7.94)*     | (7.94)*     | (7.78)*      | (7.78)*     | (7.78)*     |
| \( \Delta PP_t \)    | -0.1013      | 0.1010      | 0.0003      | -0.0700      | 0.0512      | 0.0188      | -0.0569      | 0.0569      | 0.0000      |
|                      | (-2.10)*     | (0.77)      | (0.00)      | (-1.58)      | (0.39)      | (0.16)      | (-1.55)      | (1.55)      | (---)       |
| \( \Delta PS_t \)    | 0.1152       | -0.1506     | 0.0354      | 0.0416       | 0.0286      | -0.0702     | 0.0569       | -0.1761     | 0.1192      |
|                      | (1.82)*      | (-0.92)     | (0.23)      | (0.96)       | (0.23)      | (-0.62)     | (1.55)       | (-1.80)*    | (1.33)      |
| \( \Delta PF_t \)    | -0.0261      | 0.2557      | -0.2296     | 0.0076       | 0.1500      | -0.1576     | 0.0000       | 0.1192      | -0.1192     |
|                      | (-0.61)      | (2.25)*     | (-2.13)*    | (0.22)       | (1.76)*     | (-2.04)*    | (---)        | (1.33)      | (-1.33)     |
| \( \Delta E_t \)     | 0.004        | 0.1057      | -0.1017     | -0.0069      | 0.1201      | -0.1132     | 0.0020       | 0.0636      | -0.0656     |
|                      | (0.19)       | (1.89)*     | (-2.04)*    | (-0.33)      | (2.18)*     | (-2.31)*    | (-0.12)      | (1.34)      | (-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      | 167.04       | 167.04      | 167.04      | 167.04       | 167.04      | 167.04      |

Key: *, † and ‡ represent, respectively, the 1%, 5% and 10% significance levels.
Table 4. Diagnostic tests for the equations of the DAIDS model.

| Eqn | Serial correlation | Functional form | Normality | Heteroscedasticity |
|-----|--------------------|-----------------|-----------|--------------------|
|     | LM     | F    | LM | F | LM | LM | F |
| PT  | 0.05 (0.83) | 0.02 (0.88) | 0.00 (0.97) | 0.00 (0.98) | 0.92 (0.63) | 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) | 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) | 0.02 (0.88) | 0.02 (0.88) |

on an equation-by-equation basis, leaving comparison across equations to be dealt with later, when interpreting the elasticity estimates. The interpretation of the estimation results focuses on the model denoted $HSN^L$ for which constraints are imposed only on the long run.

The adjustment velocity estimate is 0.76 for all share equations. This suggests a rapid adjustment of UK tourism demand to equilibrium after changes in its determinants. Indeed, 76% of the 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, on issues that may influence their decision to visit Portugal, Spain or France.

Since, by construction, the adjustment velocity parameter $\lambda$ is multiplied by the intercept and long-run parameters of all share equations, it should be noted that to obtain the actual long-run estimates the coefficients of the lagged variables have to be divided by the estimate of $\lambda$. As a result, the actual long-run coefficient estimates of, say, $D_{t-1}$ in the equations for Portugal, Spain and France are, respectively, $-0.019$, $-0.016$ and $0.035$.

Due to the model’s log-linear form, the elasticity values cannot be directly assessed from the coefficient estimates. However, their signs and magnitudes can provide information on both the direction and intensity of tourism demand to changes in its determinants. In general, all coefficient 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 run 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 UK demand for Spain to changes in the UK tourism budget. Conversely, in the share equation for France, these coefficients are all negative, indicating an inelastic response of UK demand for France to changes in the UK tourism budget. In the case of Portugal, none of these coefficients is statistically significant. Moreover, when significant, both the long-run 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 competing rather than complementary. The results also indicate that the political and economic events of 1974–81 had a negative effect on UK tourism demand for Portugal and Spain and a positive effect on the demand for France.

For all equations, the short-run coefficients are, in general, statistically insignificant. This may indicate that the effects on 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 dynamic model in spite of its short-run insignificance, the consistency of the long-run estimates, and the high adjustment velocity of UK demand to changes in tourism budget and prices.
Table 5. Expenditure and uncompensated own-price and cross-price elasticities.

| Country | Model notation | Expenditure elasticities | Own-price and cross-price elasticities |
|---------|----------------|--------------------------|-----------------------------------------|
|         | First period   | Second period            | First period   | Second period | First period   | Second period | First period   | Second period |
| PT      | HSN^L          | 0.913                    | 1.027         | –2.128        | –1.729        | 1.176         | 0.714         | 0.039         | –0.011        |
|         | long run       | (5.22)                   | (26.22)       | (–4.61)       | (–5.93)       | (2.62)        | (2.39)        | (0.49)        | (–0.69)       |
|         | HSN^L          | 0.911                    | –1.895        | 0.573         | 0.138         |               |               |               |               |
|         | short run      | (3.35)                   | (–3.40)       | (1.01)        | (0.28)        |               |               |               |               |
| SP      | HSN^L          | 1.076                    | 1.047         | 0.097         | 0.120         | –1.980        | –2.150        | 0.807         | 0.983         |
|         | long run       | (21.39)                  | (36.27)       | (2.30)        | (2.41)        | (–10.82)      | (–10.12)      | (5.40)        | (5.38)        |
|         | HSN^L          | 1.213                    | 0.071         | –1.048        | 0.171         |               |               |               |               |
|         | short run      | (12.38)                  | (0.32)        | (–4.56)       | (1.02)        |               |               |               |               |
| FR      | HSN^L          | 0.865                    | 0.929         | 0.012         | 0.007         | 1.730         | 1.400         | –2.608        | –2.336        |
|         | long run       | (9.65)                   | (25.34)       | (1.50)        | (1.93)        | (4.98)        | (5.05)        | (–8.76)       | (–9.36)       |
|         | HSN^L          | 0.685                    | 0.082         | –0.050        | –1.298        |               |               |               |               |
|         | short run      | (5.02)                   | (0.25)        | (–0.02)       | (–5.45)       |               |               |               |               |

A more interesting analysis of the results requires the relevant elasticity values. Given the model log-linear form, the uncompensated expenditure and price elasticities have to be computed using the coefficient estimates of the dynamic specification, and the formulae and share values given in Appendix 2. Table 5 shows these elasticity estimates and respective \( t \)-values in brackets for the DAIDS model denoted \( HSN^L \).

The long-run elasticities obtained from the DAIDS are similar in magnitude and signs to those obtained from the AIDS model of De Mello et al. (2002). Therefore, the discussion and comments about these elasticity estimates provided in that study apply, in general terms, to the long-run elasticity estimates obtained from the DAIDS. Indeed, the latter estimates not only present similar values 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 De Mello et al. model is a ‘seemingly-dynamic’ specification that seems to allow for some correction of omitted temporal factors through the addition of a trend variable and the consideration of a non-constant expenditure coefficient. However, if comparisons 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 conclusion is drawn from the rejection of the orthodox AIDS when tested against the ‘unorthodox’ form of De Mello et al., and is further supported by its rejection against the general DAIDS specified earlier. Nevertheless, examination of the short-run elasticity estimates and a comparative analysis of short-run and long-run demand behaviour cannot be assessed with these types of models, even if they are ‘seemingly-dynamic’. To carry out such analysis, we need a general dynamic framework such as the one underlying the DAIDS model. Thus we focus on the model denoted \( HSN^L \), and on the second period (the last two decades), given that it relates to the more recent behaviour of UK demand.
While both short-run and long-run estimates of the expenditure elasticities in the equation for Portugal are close to unity, the corresponding estimates for Spain and France present significant differences in their short-run and long-run magnitudes. The long-run expenditure response of UK demand for France is close to unity but, in the short run, it is clearly inelastic. In the equation for Spain, the long-run expenditure response is also close to unity but, in the short run, it is clearly elastic.\(^7\)

The short-run own-price elasticity for Spain has the lowest value of all, and the one for Portugal the largest. This indicates that UK tourists seem to be less sensitive to short-term price changes in Spain than in Portugal or France. This information, supplemented by the fact that Spain presents the highest estimate for the short-run expenditure elasticity, suggests that Spain is a preferential destination for UK tourists in the short run. Hence, Spain has a comparative advantage in relation to its neighbouring competitors, and wider scope for manoeuvre in policies involving short-term prices. However, for all destinations long-run elasticity estimates are close to \(-2\), indicating that UK tourists are highly sensitive to long-term price changes and will penalize, in a similar way, any of the destinations for an increasing price policy. Therefore, price policies should clearly separate long-run and short-run decisions and, in the particular case of Spain, 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 Portugal and France.

The inference concerning the cross-price effects drawn from the study of De Mello et al.\(^\) study applies, in general terms, to the long-run estimates of our dynamic model. Indeed, the lack of sensitivity of UK demand for tourism in Portugal (France) to price changes in France (Portugal), the decreasing response of the UK demand for tourism in Portugal and France to price changes in Spain, and the increasing response of the demand for Spain to price changes in Portugal or France are common features of the estimation results provided by both models. However, the DAIDS model permits analysis of short-run cross-price elasticities that is not possible with the De Mello et al. AIDS approach. An interesting feature of the short-run cross-price elasticities when compared with their corresponding values in the long run is particularly worth noting: none of the short-run elasticities is statistically significant at the 5% significance level, while their long-run counterparts are all significant except, of course, for the case of Portugal versus France. 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 accordingly, with significant effects for the destinations in question. Thus destinations are more likely to retain their tourism receipts if they are able to avoid long-run increases in their own prices and if competitors keep their price cuts within the short run.
Conclusion

Obtaining reliable information about how tourists allocate expenditure and how they adjust to equilibrium is of considerable interest for tourism analysis and policy making in the area of tourism demand. As consumers, in general, do not immediately adjust to changes in their demand determinants, appropriate dynamic systems are essential before plausible behavioural hypotheses can be tested. Even so, these problems have been largely ignored in tourism demand research, as suitable dynamic generalizations of demand systems are still a rare feature in empirical studies.

With most economic models, and particularly with demand systems, the analysis is usually centred on the long-run coefficients. This analysis is independent of any short-run dynamics fitted to the data. Yet as the dynamic specification may be affected by the restrictions imposed on the long-run parameters, the use of inappropriate short-run dynamics may, in turn, affect the outcome of tests conducted on the long run. Hence, both static specifications (typically ruling out theoretically plausible features involving short-run dynamics) and inappropriate dynamic structures lead to misspecifications that may give rise to unreliable estimates, invalid inference and inaccurate forecasts. Flawed estimates produce misleading analysis that may induce inadequate policy measures.

The main objective of this paper has been to show that carefully built models do matter for the progress of research in tourism demand analysis. To make valuable progress in research, we need dynamic specifications based on theory that establishes the main variables to be modelled, suggests the mathematical formulations to be derived and proposes the fundamental restrictions to be tested. We also need to decipher the data meticulously (not ‘torture’ them) in order to incorporate fundamental dummy variables that adequately translate existing structural breaks and one-off historical events. Such dynamic specifications allow us to build empirical applications that, with no statistical ambiguities, will give sound and reliable answers to the genuine problems which are so often faced by business managers and policy makers.

In this paper the estimation results of the flexible general dynamic form of the AIDS system provide empirical evidence of the robustness of this methodology for conducting tourism demand analysis in a temporal context. This dynamic model offers reliable evidence on the inadequacy of the orthodox AIDS to reconcile data and theory consistently within its formulation. Moreover, the tests carried out prove that the DAIDS model, besides encompassing specific dynamic formulations such as the ARDL and the PA, is consistent with the postulates of utility theory and provides robust and empirically plausible estimates for both the long run and the short run.

If a dynamic formulation is required to describe the behaviour of an origin country’s tourism demand adequately, then any static formulation of the phenomenon would be a misspecification. However, the long-run elasticity estimates of the De Mello et al (2002) static AIDS are similar to those of the DAIDS. Why is this so? The main reasons are as follows. First, the AIDS system of De Mello et al includes dynamic-like elements (such as the expenditure structural break and a trend variable) which, if ignored, would result unambiguously in a misspecified model. Consequently, and in contrast with an
orthodox static AIDS, the De Mello et al 'seemingly dynamic' model is consistent with the utility maximization assumptions of consumer theory and supplies plausible and significant long-run estimates. Second, in the particular case of UK tourism demand, the short-run adjustment process is very fast. Given the estimated adjustment velocity, more than three-quarters of the total impact on the dependent variable, due to sustainable unit changes of the explanatory variables, takes place within the current period. This being the case, the 'unorthodox' AIDS model of De Mello et al can be thought as an acceptable solution for the estimation of long-run effects. However, if this were not the case, there would not be any 'seemingly dynamic' elements capable of saving the De Mello et al formulation from obvious misspecification. Indeed, with a slower adjustment velocity, the short-run correction process would be of significant importance, and the choice of De Mello et al would be an unrealistic approach to the estimation of the long-run impacts on tourism demand. That is why dynamic specifications should always be a first choice, although not necessarily the last.

Therefore, an appropriate dynamic specification should always be implemented first when modelling demand systems, as this is the correct means of obtaining reliable estimates of both long-run and short-run responses. If, after the estimation of the appropriate dynamic form, indications are given of the possible irrelevance of the short-run adjustment process, then and only then can a static formulation be considered within the spirit of 'general-to-specific' modelling. Imposing dynamics where dynamics is not mandatory seems to be an unjustifiable attempt to comply with the most recent trends in econometric modelling rather than considering the requirements of the data.

The estimation of the DAIDS model, besides defining the conditions under which a static AIDS system can provide reliable long-run information, also indicates directions for future research. An interesting feature uncovered by the estimation results is the fact that the utility theory hypotheses hold in both the long run and the short run. In dynamic specifications utility theory constraints are, generally, tested for the long run and not for the short run. The motivation rests on the idea that consumers may not have fully adjusted to changing circumstances in the short run and, consequently, homogeneity and symmetry may not be observed. At this point, it is opportune to call on the findings of Anderson and Blundell (1984, p 40) who, confronted with a similar situation, state:

\[ \Pi, \] 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.

Given that, for the general dynamic structure estimated here, both long-run and short-run homogeneity and symmetry hold, the inherent implication is that the rationality of utility maximization postulates is observed for both the long-run and short-run behaviour of UK tourism demand. A plausible explanation for this is that the general dynamic model, being sufficiently robust to track UK demand behaviour accurately over the sample period, indicates that UK tourists adjust very rapidly to changes in their demand determinants. This rapid adjustment process implies that most of the short-run coefficients either should
be insignificant or should have irrelevant magnitudes. Indeed, this is the general indication of the estimates obtained. Hence, the more rapidly consumers adjust their demand behaviour, the less significant the short-run effects should be, 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.

Endnotes

1. According to Durbarry and Sinclair (2003, p 934), 'a full dynamic specification of the system of equations model is provided in De Mello and Sinclair (2000a, 2000b) and De Mello (2001)'. Therefore, the work of Durbarry and Sinclair (2003) does not constitute 'the first attempt to use the error correction AIDS approach in tourism demand modelling and forecasting' as, erroneously, stated by Li et al (2004, p 142).

2. When the ARDL specification is tested without the null intercepts restriction, the model is rejected with a statistic value of $\chi^2 (9) = 27$.

3. Model (11), including $FE_{t-1}$ and $SE_{t-1}$, is an equivalent form of a model including $E_{t-1}$ and $SE_{t-1}$. The former has the advantage of giving straightforward information on the coefficients of variable $E_{t-1}$ in the first and second periods ($a_5$ and $a_6$ respectively), whereas in the latter the information for the second period has to be obtained by summing $E_{t-1}$ and $SE_{t-1}$ coefficients.

4. Null cross-price effects between Portugal and France are imposed because we assume that price changes in Portugal do not affect UK tourism demand for France and vice versa. See De Mello et al (2002).

5. Given the system singularity, estimation was carried out by deleting one of the three equations. Since the results are invariant whichever equation is deleted, we choose to omit the share equation for Portugal. The coefficient estimates of this equation were later retrieved from the coefficient estimates of the other two, using the adding-up property explained in Appendix 1.

6. 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.

7. If the average UK tourist (medium income, medium education, medium age) receives a salary increase or a productivity premium, or is lucky at the races, and acquired an extra couple of thousand pounds, he or she would want to take the family on the first flight to Malaga, Costa del Sol or Benidorm (cheaper than the train to Paris) for a short relaxing vacation on the beach, where food and accommodation are good and cheap. He or she would not want to visit the cold north of France or busy and expensive Paris. If the decision is to go south and head to a beach (as it is in the majority of cases), this type of British tourist prefers to go all the way to the South of Spain. Tourists who go to France instead have a different habit framework. They tend to be upper middle class, have a higher level of education and are, probably, older. They do not change much when their budget fluctuates a little. Every other month they take the train to Paris to offer their partner a romantic dinner at Chez Pierre, they do not miss the Cézanne exhibition at the Pompidou Centre, they love the quiet Brittany beaches, they absolutely cannot do without the famous bouillabaisse of Madame Letice in Marseille, they love the Riviera off-season, and they do not miss one Cannes Festival. British tourists heading for Spain who correspond to this latter set of characteristics are as few as those who fulfil the former are many. Therefore, the expenditure response to changes in UK tourists’ budgets will be elastic for Spain, but inelastic for France.

References

Anderson, G., and Blundell, R. (1983), 'Testing restrictions in a flexible dynamic demand system: an application to consumers' expenditure in Canada', Review of Economic Studies, Vol 50, No 3, pp 397–410.

Anderson, G., and Blundell, R. (1984), 'Consumer non-durables in the UK: a dynamic demand system', Economic Journal, Vol 94, No 376a, pp 35–44.

De Mello, M. (2001), Theoretical and Empirical Issues in Tourism Demand Analysis, PhD thesis, University of Nottingham, Nottingham.
Appendix 1

Derivation of Deaton and Muellbauer’s (1980a; 1980b) AIDS model

Let $x$ be the exogenous budget or total expenditure which is to be spent within a given period on some or all of $n$ goods. These goods can be bought in non-negative quantities $q_i$ at given prices $p_i$, $i = 1, \ldots, n$. Let $q = (q_1, q_2, \ldots, q_n)'$ be the quantities vector of the $n$ goods purchased, and $p = (p_1, p_2, \ldots, p_n)'$ be the price vector. The budget constraint of a representative consumer is $\sum_{i=1}^{n} p_i q_i = x$. Defining the utility function as $u(q)$, the consumer's aim is to maximize the utility, subject to the budget constraint:

$$\max u(q), \text{ subject to } \sum_{i=1}^{n} p_i q_i = x. \quad (A1)$$

The solution for this maximization problem leads to the Marshallian (uncompensated) demand functions $q_i = g_i(p, x)$. Alternatively, the consumer's problem can be defined as the minimum total expenditure necessary to attain a specific utility level $u^*$, min
\[ \sum_{i=1}^{n} p_i q_i, \text{ subject to } u(q) = u^*. \]

The solution for this minimization problem leads to the Hicksian (compensated) demand functions \( q_i = h_i(p, u). \) Therefore, a cost function can be defined as:

\[ c(p, u) = \sum_{i=1}^{n} p_i h_i(p, u) = x. \] (A2)

Given total expenditure \( x \) and prices \( p, \) the utility level \( u^* \) is derived from the problem solution stated in Equation (A1). Solving Equation (A2) for \( u \) an indirect utility function is obtained such that \( u = u(p, x). \)

The AIDS model specifies a cost function, which is used to derive the demand functions for the commodities under analysis. The derivation process can be summarized in the following three steps: first, \( \partial c(p, u)/\partial p_i = h_i(p, u) \) is derived establishing the Hicksian demand functions; second, solving (A2) for \( u \) the indirect utility function is obtained, such that \( u = u(p, x); \) and, finally, \( b(p, u(p, x)) = g(p, x) \) is retrieved stating the Hicksian and the Marshallian demand functions as equivalent. The cost function of AIDS model is:

\[ \ln c(p, u) = a(p) + ub(p). \] (A3)

where \( a(p) = \alpha_0 + \sum_{i=1}^{n} \alpha_i \ln p_i + 2^{-1} \sum_{i=1}^{n} \sum_{j=1}^{n} \gamma_{ij} \ln p_i \ln p_j \text{ and } b(p) = \beta_0 \prod_{i=1}^{n} p_i^{\beta_i}. \)

The derivative of Equation (A3) with respect to \( \ln p_i \) is

\[ \frac{\partial \ln c(p, u)}{\partial \ln p_i} = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln p_j + u\beta_i \prod_{i=1}^{n} p_i^{\beta_i}. \] (A4)

As \( a(p, u) = x, \) then \( \ln x = a(p) + ub(p) \) and, solving for \( u, \)

\[ u = \frac{\ln x - a(p)}{b(p)}. \] (A5)

Substituting (A5) in (A4), we have:

\[ \frac{\partial \ln c(p)}{\partial \ln p_i} = \frac{\partial c(p)}{\partial p_i} \frac{p_i}{c(p)} = h_i(p) \frac{p_i}{c(p)} = \frac{p_i q_i}{c(p)} = w_i = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln p_j + \beta_i (\ln x - a(p)). \]

Defining a price index \( P \) such that \( \ln P = a(p), \) we further obtain:

\[ w_i = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln p_j + \beta_i \ln \left( \frac{x}{P} \right). \] (A6)

where

\[ \ln P = \alpha_0 + \sum_{k=1}^{n} \alpha_k \ln p_k + 2^{-1} \sum_{k=1}^{n} \sum_{l=1}^{n} \gamma_{kl} \ln p_k \ln p_l. \] (A7)

Equations (A6) and (A7) are the basic equations of the AIDS model.

In a tourism demand context, \( i \) is a destination country among a group of \( n \) alternative destinations demanded by tourists of a given 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 tourism prices \( p_i \) in \( i \) and alternative destinations \( j \) and by the per capita expenditure \( x \) allocated to the set of destinations, deflated by price index \( P. \) By using properties of Hicksian and Marshallian demand functions, the model satisfies the following:

- **adding-up restriction** requiring that all budget shares sum up to unity: \( \sum_{i=1}^{n} \alpha_i = 1, \sum_{i=1}^{n} \beta_i = 0, \sum_{i=1}^{n} \gamma_{ij} = 0, \text{ for all } j; \)
- **homogeneity restriction** requiring that a proportional change in all prices and expenditure has no effect on the quantities purchased: \( \sum_{j=1}^{n} \gamma_{ij} = 0, \text{ for all } i; \)
Dynamic systems for modelling tourism demand

- symmetry restriction requiring consumer consistent choices: $\gamma_{ij} = \gamma_{ji}$ for all $i \neq j$;

- negativity restriction requiring that a rise in prices results in a fall in demand — that is, the condition of negative own-price elasticities for all destinations.

The restrictions imposed on $\alpha$ and $\gamma$ comply with these assumptions and ensure that Equation (A7) 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 and Muellbauer, 1980b, p 76). Hence, the deflator $P$ in Equation (A7) can be substituted by the Stone price index $\ln P^*$ such that:

$$\ln P^* = \sum_{j=1}^{n} w_j^B \ln p_j$$

(A8)

where $w_i^B$ is the budget share of destination $i$ in the base year. With this simplification for $P$, the system of Equation (A6) can be rewritten and estimated in the following form:

$$w_i = \alpha_i^* + \sum_{j=1}^{n} \gamma_{ij} \ln p_j + \beta_i \ln \left( \frac{x}{P^*} \right)$$

(A9)

Appendix 2

Expenditure, own-price and cross-price elasticities

Expenditure and price elasticities cannot be directly accessed in Equation (A9), given its linear-log form. Nevertheless, the elasticities values can be retrieved from the coefficients in (A9), using the following formulae:

- **expenditure elasticity**
  $$\epsilon_i = \frac{1}{\bar{w}_i} \frac{\partial \bar{w}_i}{\partial \ln x} + 1 = \frac{\beta_i}{\bar{w}_i} + 1,$$

- **uncompensated own-price elasticity**
  $$\epsilon_{ui} = \frac{1}{\bar{w}_i} \frac{\partial w_i}{\partial \ln p_i} - 1 = \frac{\gamma_{ii}}{\bar{w}_i} - \beta_i \frac{w_i^B}{\bar{w}_i} - 1,$$

- **uncompensated cross-price elasticity**
  $$\epsilon_{ij} = \frac{1}{\bar{w}_i} \frac{\partial w_i}{\partial \ln p_j} - \frac{\gamma_{ij}}{\bar{w}_i} = \beta_i \frac{w_j^B}{\bar{w}_i} - \beta_i \frac{w_j^B}{\bar{w}_i},$$

- **compensated own-price elasticity**
  $$\epsilon_{ui}^* = \epsilon_{ui} + w_i^B \epsilon_i = \frac{\gamma_{ii}}{\bar{w}_i} + w_i^B - 1,$$

- **compensated cross-price elasticity**
  $$\epsilon_{ij}^* = \epsilon_{ij} + w_i^B \epsilon_i = \frac{\gamma_{ij}}{\bar{w}_i} + w_j^B,$$

where $\bar{w}_i$ is the sample’s average share of destination $i$, $i = 1,\ldots,n$, and $w_j^B$ is the share of destination $j$, $j = 1,\ldots,n$, in the base year.
Appendix 3

AIDS model of UK tourism demand for Portugal, Spain and France

The AIDS model assumes that consumers allocate their budgets to commodities in a multi-stage budgeting process implying independent preferences. Thus for the UK tourism demand AIDS model, it is assumed that the UK tourism expenditure allocated to Portugal, Spain and France is separable from that allocated to other destinations and the decision to spend money in those countries is made in several stages. First, UK tourists allocate their budgets to tourism and other goods; second, they allocate them to tourism in Portugal, Spain and France and other destinations; and, finally, they decide between Portugal, Spain or France. The AIDS system is applied to this last stage using the following form:

\[
\begin{align*}
WP_i &= \alpha_p + \gamma_{pp}PP_i + \gamma_{ps}PS_i + \gamma_{pf}PF_i + \beta_pE_i + u_P \\
WS_i &= \alpha_s + \gamma_{sp}PP_i + \gamma_{ss}PS_i + \gamma_{sf}PF_i + \beta_sE_i + u_S \\
WF_i &= \alpha_f + \gamma_{fp}PP_i + \gamma_{fs}PS_i + \gamma_{ff}PF_i + \beta_fE_i + u_F
\end{align*}
\]

Appendix 4

Definition of variables

The variables integrating the AIDS model of UK tourism demand for Portugal, Spain and France are the shares of UK tourism budget allocated to these destinations, \( WP, WS \) and \( WF \); destination tourism prices, \( PP, PS, PF \); and UK real per capita tourism budget, \( E \). Each share \( Wi, i = P \) (Portugal), \( S \) (Spain) and \( F \) (France), is defined as:

\[
Wi = \frac{\text{EXP}_i}{\text{EXP}_P + \text{EXP}_S + \text{EXP}_F},
\]

where \( \text{EXP}_i \) is the nominal tourism expenditure allocated by UK tourists to destination \( i \). The effective price of tourism in destination \( i \) is defined as

\[
\pi_i = \ln \left( \frac{\text{CPI}_i}{\text{CPI}_\text{UK} R_i} \right),
\]

where \( \text{CPI}_i \) is destination \( i \) consumer price index, \( \text{CPI}_\text{UK} \) is UK consumer price index, \( R_i \) is the exchange rate between \( i \) and the UK, defined as number of units of country \( i \) currency per unit of the UK currency. The per capita UK real tourism expenditure allocated to all destinations is:

\[
E = \ln \left( \frac{\sum_{i=1}^{n} \text{EXP}_i}{\text{UKP} \pi'} \right),
\]

where \( \text{UKP} \) is UK population and \( \ln\pi' \) is the Stone index defined in Equation (A8).
Appendix 5

Data sources

The data for UK tourism expenditure, disaggregated by destinations and measured in million pounds sterling, were obtained from the Business Monitor MA6 (1970–93), continued as Travel Trends (1994–98). Data on the UK population, price indexes and exchange rates were obtained from the International Financial Statistics Yearbooks of the International Monetary Fund (1980–98).