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Testing alternative dynamic systems for modelling tourism demand

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Abstract

The goals in this paper are to contribute an empirical study of tourism demand dynamics, and to point out areas where the scrutiny of relationships between theoretical and empirical considerations are likely to produce new insights in this area of research. A flexible general form of a Dynamic Almost Ideal Demand System (DAIDS) is derived to analyse the UK tourism demand for its geographically proximate neighbours France, Spain and Portugal, in the period 1969-1997. 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 to conduct tourism demand analysis in a temporal context. Moreover, the dynamic model offers statistically strong evidence on the inadequacy of the orthodox static AIDS and the other restricted models to reconcile consistently 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 long and short run sensitivity of the 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.

1 Introduction

In what concerns tourism demand analysis, early research efforts (White (1985), O'Hagan and Harrison (1984), Syriopoulos and Sinclair (1993), Papatheodorou (1999), De Mello, Pack and Sinclair (2002)) concentrate on static specifications based on Deaton and Muellbauer's (1980*a*, 1980*b*) 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

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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.

Within a demand analysis context, 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 only be acceptable 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 been generally recognised, specific research in tourism demand using dynamic systems is not abundant and, to the best of our knowledge, only few address these issues in empirical studies (Lyssiotou (2000), Song and Witt (2000), De Mello and Sinclair (2000), De Mello (2001), Li, Song and Witt (2004), De Mello and Nell (2005)). 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. Hence, the objectives of this paper are to contribute an empirical study of tourism demand dynamics and to point out areas where the scrutiny of relationships between theoretical and empirical considerations are likely to produce new insights in this area.

The structure of the paper is as follows. In Section 2, a flexible general dynamic form of an almost ideal demand system (DAIDS) is derived. Nested within this general dynamic form, are several alternative specifications such as the static orthodox AIDS model itself, the partial adjustment (PA) model and the auto-regressive distributed lag (ARDL) model. These models are used to analyse the UK tourism demand for France, Spain and Portugal, in the period 1969-1997. The UK is a major tourism origin of particular importance for the three neighbouring destinations under analysis, and Spain and Portugal are interesting cases as they experienced considerable economic development and major political changes during the sample period in contrast with France, which was a developed and politically steady country over the same period. These features permit the analysis of significant aspects concerning changes in the inter-dependencies and competitive behaviour of the destinations over the sample period. In Section 3, we implement several 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 flexible dynamic system DAIDS proves to be an appropriate and statistically robust form for conducting tourism demand analysis. Section 4 presents the estimation results for long and short run tourism price and budget elasticities, which permit a comparative analysis of the relative magnitude and relevance of the long and short run sensitivity of the UK tourism demand to changes in its determinants. Section 5 concludes.

2 Dynamic AIDS modelling of the UK demand for tourism

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 A.1. For simplicity, we rename

the variables as $w_i = Wi$, $\ln p_j = Pj$, $i, j = 1, \dots, n$, and $\ln(x/P^*) = E$. Hence, equation (A.11) is written as

$$Wi_t = \alpha_i^* + \sum_{j=1}^n \gamma_{ij} Pj_t + \beta_i E_t, \quad (2.1)$$

where Wi represents the expenditure share allocated to the i th destination by UK tourists, Pj 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 (2.1) as the appropriate choice for the steady state structure of the following general dynamic stochastic specification

$$\Delta Wi_t = \sum_{j=1}^n \gamma_{ij}^S \Delta Pj_t + \beta_i^S \Delta E_t + \lambda_i \left(\alpha_i + \sum_{j=1}^n \gamma_{ij}^L Pj_{t-1} + \beta_i^L E_{t-1} - Wi_{t-1} \right) + ui_t, \quad (2.2)$$

where λ_i is the adjustment coefficient of the i th equation, Δ is the first difference operator, subscript $t-1$ indicates the variables' lagged values, 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 and long run responses of the dependent variable to changes in its determinants. Equation (2.2) assumes that to maintain the steady state relationship (2.1), 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 (2.2) 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 are performed to find the most restrictive dynamic specification, which 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 (2.2)

$$\begin{aligned} \Delta Wi_t = & \gamma_{i1}^S \Delta P1_t + \dots + \gamma_{in}^S \Delta Pn_t + \beta_i^S \Delta E_t + \\ & + \lambda \alpha_i + \lambda \gamma_{i1}^L P1_{t-1} + \dots + \lambda \gamma_{in}^L Pn_{t-1} + \lambda \beta_i^L E_{t-1} - \lambda Wi_{t-1} + ui_t. \end{aligned} \quad (2.3)$$

• 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

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

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 model's general form (2.3) can be reduced to an ARDL form as described, under the null hypothesis

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

If H_0 is not rejected, the model reduces to

$$Wi_t = \gamma_{i1} (P1_t - (1 - \lambda) P1_{t-1}) + \dots + \gamma_{in} (Pn_t - (1 - \lambda) Pn_{t-1}) + \beta_i (E_t - (1 - \lambda) E_{t-1}) +$$

$$+(1 - \lambda)Wi_{t-1} + ui_t, \quad (2.4)$$

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_t^* , is a linear function of an explanatory variable, say X_t , such that

$$Y_t^* = \beta_0 + \beta_1 X_t + u_t. \quad (2.5)$$

The partial adjustment hypothesis postulates

$$Y_t = \delta Y_t^* + (1 - \delta)Y_{t-1}. \quad (2.6)$$

Substituting (2.5) in (2.6) leads to

$$Y_t = \delta\beta_0 + \delta\beta_1 X_t + (1 - \delta)Y_{t-1} + \delta u_t. \quad (2.7)$$

Considering now model (2.3), the null hypothesis to be tested is

$$H_0 : \gamma_{ij}^S = \lambda\gamma_{ij}^L = \gamma_{ij} \wedge \beta_i^S = \lambda\beta_i^L = \beta_i, \text{ for all } i, j.$$

If H_0 is not rejected, the model reduces to

$$Wi_t = \lambda\alpha_i + \gamma_{i1}P1_t + \dots + \gamma_{in}Pn_t + \beta_i E_t + (1 - \lambda)Wi_{t-1} + ui_t, \quad (2.8)$$

which is a partial adjustment model similar to the one described in equation (2.7).

• **Static AIDS model**

To test for the static form nested within (2.3) the null hypothesis is

$$H_0 : \gamma_{ij}^S = \gamma_{ij}^L = \gamma_{ij} \wedge \beta_i^S = \lambda\beta_i^L = \beta_i \wedge \lambda = 1, \text{ for all } i, j.$$

If H_0 is not rejected, model (2.3) reduces to

$$Wi_t = \lambda\alpha_i + \gamma_{i1}P1_t + \dots + \gamma_{in}Pn_t + \beta_i E_t + ui_t, \quad (2.9)$$

which is the steady state orthodox AIDS model.

3 Testing the consistency of alternative dynamic models

The methodological strategy in this section develops in the following way. 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 maximisation restrictions on specifications not rejected by the data.

Considering the specific demand context under analysis, the dependent variable of the DAIDS model in equation (2.3) represents the UK tourism demand share Wi 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

$$\begin{aligned} \Delta W i_t = & \lambda \alpha_i + \lambda \gamma_{iP}^L P P_{t-1} + \lambda \gamma_{iS}^L P S_{t-1} + \lambda \gamma_{iF}^L P F_{t-1} + \lambda \beta_i^L E_{t-1} - \lambda W i_{t-1} + \\ & + \gamma_{iP}^S \Delta P P_t + \gamma_{iS}^S \Delta P S_t + \gamma_{iF}^S \Delta P F_t + \beta_i^S \Delta E_t + u i_t. \end{aligned} \quad (3.1)$$

Model (3.1) is the general dynamic structure DAIDS upon which appropriate restrictions described earlier are imposed to obtain equations (2.4), (2.8) and (2.9), respectively, the ARDL, PA and static AIDS models. These models' compatibility with the data is tested below. If the models are compatible, they are further subjected to utility theory constrains and tested. The non-rejection of these constraints indicates theory-consistent models.

The Wald test performed on the alternative models provided the following results. The orthodox static AIDS hypothesis presents a statistic value of $\chi^2(10) = 46.10$, which lies above the 1% critical value, implying its rejection. The partial adjustment hypothesis against the general DAIDS presents a statistic value of $\chi^2(9) = 22.63$, which lies above the 1% critical value, implying its rejection. The ARDL hypothesis against the general DAIDS presents a statistic value of $\chi^2(11) = 103.49$, which lies above the 1% critical value, implying its rejection.¹ Only the dynamic structure (3.1) reveals itself fully compatible with the data and with the assumptions of consumer demand theory. Therefore, the remainder of this section focuses on this model.

The analysis of the data indicates the possible presence of a structural break in the coefficient of the real expenditure variable E , dividing the sample into two sub-periods: 1969-1979 and 1980-1997. The analysis of the data also indicates the possible relevance of an intercept-dummy D in the share equations for France, Spain and Portugal, over the period 1974-1981. This information is integrated in specification (3.1) in the following way. We assume that the structural break is only relevant in the long run. Hence, we add to equation (3.1) a new variable $S E_{t-1}$. This variable is constructed using a step-dummy variable S for the year 1979. S is then multiplied by E_{t-1} giving rise to variable $S E_{t-1}$ which is included in equation (3.1). We assume that the intercept-dummy variable D may have significant effects both in the long run and in the short run. Therefore, we add to equation (3.1) variables D_{t-1} and ΔD_t . Assessment on the statistical significance of added variables $S E_{t-1}$, D_{t-1} and ΔD_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 results (p-values in brackets).

Hypotheses under test	Wald statistic
H_0 : Non-significance of ΔD_t	
U : $PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} W i_{t-1} \Delta PP_t \Delta PS_t \Delta PF_t \Delta E_t \Delta D_t D_{t-1} S E_{t-1}$	$\chi^2(2) = 1.78 (0.41)$
R : $PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} W i_{t-1} \Delta PP_t \Delta PS_t \Delta PF_t \Delta E_t D_{t-1} S E_{t-1}$	Not rejected
H_0 : Joint non-significance of D_{t-1} and $S E_{t-1}$	
U : $PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} W i_{t-1} \Delta PP_t \Delta PS_t \Delta PF_t \Delta E_t D_{t-1} S E_{t-1}$	$\chi^2(4) = 18.06 (0.00)$
R : $PP_{t-1} PS_{t-1} PF_{t-1} E_{t-1} W i_{t-1} \Delta PP_t \Delta PS_t \Delta PF_t \Delta E_t$	Rejected

Table 1: Tests for the individual and joint significance of variables $S E_{t-1}$, D_{t-1} and ΔD_t

In Table 1, the test indicates that D is only relevant in the long run, since the non-significance of

¹When the ARDL specification is tested without the null intercepts restriction, the model is rejected with a statistic value of $\chi^2(9) = 27.00$.

ΔD_t cannot be rejected. The joint significance of D_{t-1} and SE_{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 λ_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 considerations above, the i th equation of the general DAIDS model of UK tourism demand for destination i is

$$\Delta W i_t = a_{i1} + a_{i2} P P_{t-1} + a_{i3} P S_{t-1} + a_{i4} P F_{t-1} + a_{i5} F E_{t-1} + a_{i6} S E_{t-1} + a_{i7} D_{t-1} - \lambda W i_{t-1} + a_{i8} \Delta P P_t + a_{i9} \Delta P S_t + a_{i10} \Delta P F_t + a_{i11} \Delta E_t + u i_t. \quad (3.2)$$

The structural break that divides the expenditure observations into two sub-periods 1969-1979 and 1980-1997 is included in equation (3.2) 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 $F E_{t-1}$ and $S E_{t-1}$.²

Hypotheses under test ³	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 France and Portugal	$\chi^2(5) = 2.56$ (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
H_0 : Long run homogeneity, symmetry and null cross-price effects and short run homogeneity and symmetry	$\chi^2(8) = 10.35$ (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	$\chi^2(9) = 10.59$ (0.31) Not rejected

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

The general DAIDS (3.2) is now tested against further constrained models under the restrictions of homogeneity, symmetry and null cross-price effects between the equations for France and Portugal, both in the long and short run.⁴ Table 2 presents these test results.

²Model (3.2) including $F E_{t-1}$ and $S E_{t-1}$ 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 (respectively, a_5 and a_6), whereas in the latter, the information for the second period has to be obtained by summing E_{t-1} and SE_{t-1} coefficients.

³The 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 χ^2 statistic includes this first restriction.

⁴Null cross-price effects between France and Portugal 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).

The results in Table 2 confirm that none of the hypotheses is rejected by the data. Hence, 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 for 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, 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 formulations. Neither 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.

Explanatory variable	Unrestricted			HSN^L			HSN^{LS}		
	PT	SP	FR	PT	SP	FR	PT	SP	FR
constant	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) [◦]	0.0000 (—)	-0.0472 (-2.52) [◦]	0.0472 (2.52) [◦]	0.0000 (—)
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) [•]	0.0000 (—)	0.3992 (5.38) [•]	-0.3992 (-5.39) [•]	0.0000 (—)	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) [•]	0.792 (7.45) [•]	0.792 (7.45) [•]	0.758 (7.94) [•]	0.758 (7.94) [•]	0.758 (7.94) [•]	0.779 (7.78) [•]	0.779 (7.78) [•]	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)	0.0000 (—)
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) [*]	0.0000 (—)	0.1192 (1.33)	-0.1192 (-1.33)
E_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		

Table 3: Estimation results for the unrestricted dynamic, HSN^L and HSN^{LS} models

4 Empirical results and their interpretation

Table 3 presents the estimation results, obtained with SUR method, for the following three models: first, the general DAIDS model (3.2), under the sole restriction of equal adjustment coefficients; second,

the same model under the additional restrictions of homogeneity H , symmetry S and null cross-price effects between the equations of France and Portugal N , applied only in the long run L , denoted as HSN^L and third, 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} .⁵ Table 3 shows the coefficient estimates (asymptotic t-values in brackets) for the share equations of Portugal (PT), Spain (SP) and France (FR). Symbols \bullet , \circ and \star 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.

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.

Equation	Serial correlation		Functional form		Normality		Heteroscedasticity	
	LM	F	LM	F	LM	F	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)

Table 4: Diagnostic tests for the equations of the DAIDS model

Interpretation of the elasticity estimates

A thorough analysis of the dynamic 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 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 estimate suggests a rapid adjustment of UK tourism demand to equilibrium, after changes in its determinants. Indeed, 76% of that adjustment are attained in the current period, and only 24% are 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, it should be noted that, to obtain the actual estimates of the long run coefficients, the coefficients of the lagged variables have to be divided by the estimate of λ . 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 .

Due to the log-linear form of the model, elasticities 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 and short run and

⁵Given the system singularity, estimation was carried out by deleting one of the three equations. Since results are invariant whichever equation is deleted, we choose to omit the share equation for Portugal. The coefficient estimates of this equation were later retrieve from the coefficient estimates of the other two.

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 the UK 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.⁶ 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 on the UK tourism demand for Spain and Portugal and a positive effect for France.

For all equations, the short run coefficients are, in general, 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 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.

A more interesting analysis of the results requires the relevant elasticities 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 shares values given in Appendix A. Table 5 shows these elasticity estimates and respective t-values in brackets, for the DAIDS model denoted HSN^L .

Country	Model notation	Expenditure elasticities		Own-price elasticities		Cross-price elasticities					
		first period	second period	first period	second period	PP		PS		PF	
						first period	second period	first period	second period	first period	second period
PT	HSN^L long run	0.913 (5.22)	1.027 (26.22)	-2.128 (-4.61)	-1.729 (-5.93)			1.176 (2.62)	0.714 (2.39)	0.039 (0.49)	-0.011 (-0.69)
	HSN^L short run	0.911 (3.35)		-1.895 (-3.40)				0.573 (1.01)		0.138 (0.28)	
SP	HSN^L 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)			0.807 (5.40)	0.983 (5.38)
	HSN^L short run	1.213 (12.38)		-1.048 (-4.56)		0.071 (0.32)				0.171 (1.02)	
FR	HSN^L 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)		
	HSN^L short run	0.685 (5.02)		-1.298 (-5.45)		0.082 (0.25)		-0.050 (1.96)			

Table 5: Expenditure and uncompensated own-price and cross-price elasticities

The long run elasticities obtained from the DAIDS are similar in magnitude and signs to those obtained from De Mello et al.'s (2002) AIDS model. Therefore, the discussion and comments about these elasticity estimates provided there 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

⁶The 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.

periods. This should be expected as the De Mello et al.’s model is a “seemingly-dynamic” specification which 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 in Section 2. Nevertheless, examination of the short run elasticity estimates and a comparative analysis of short and long run demand behaviour cannot be assessed with these type 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 more recent behaviour of the UK demand.

While both short 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 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.

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 France or Portugal. This information, supplemented with 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 a wider scope for manoeuvre concerning 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 penalise, in a similar way, any of the destinations for an increasing price policy. Therefore, price policies should clearly separate long and short run decisions and, in the particular case of Spain, 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 inference concerning the cross-price effects drawn from De Mello et al.’s study apply, in general terms, to the long run estimates of our 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 DAIDS model permits the analysis of short run cross-price elasticities, not possible with De Mello et al. AIDS approach. 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% significance level. In contrast, their long run counterparts are all significant 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.

5 Conclusion

With most economic models, and particularly with demand systems, the analysis is usually centered on the long run coefficients, and is independent from 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 rulling out theoretically plausible features involving short run dynamics), and inappropriate dynamic structures lead to dynamic misspecifications that may give rise to unreliable estimates, invalid inference, and inaccurate forecasts. Flawed estimates produce misleading analysis that may induce inadequate policy measures.

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

De Mello et al.'s (2002) AIDS system allows for dynamic-like elements in its share equations. Consequently, and in contrast with the orthodox static AIDS, that "seemingly-dynamic" model is consistent with the utility maximisation assumptions of consumer theory. However, information about the short run adjustment process cannot be accessed with this approach, and a clear separation between short and long run effects cannot be made. Hence, an appropriate dynamic specification does matter when modelling demand systems, and it is the correct means to obtain reliable estimates of both long and short run responses of tourism demand to changes in its determinants.

In this paper, we specify a flexible general dynamic form of the AIDS system whos estimation results provide 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 De Mello et al.'s "seemingly-dynamic" system to provide reliable long run information, and on the inadequacy of the orthodox AIDS and other specific dynamic models, to reconcile consistently data and theory within their formulations. The tests carried out prove that the general dynamic AIDS encompasses specific formulations such as the ARDL and the PA models, is consistent with the postulates of utility theory, and provides robust and empirically plausible estimates.

The estimation results also indicate directions for future research. For instance, an interesting feature uncovered by the estimation of the dynamic AIDS is the fact that the utility theory hypotheses hold, both in the long and in the short run. At this point, it is opportune to call upon the findings of Anderson and Blundell (1984) who, when confronted with a similar situation say: "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 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. Given that, for the general dynamic structure estimated here, both long and short run homogeneity and symmetry hold, the inherent implication is that the rationality of utility maximisation postulates

is observed for both the long and short run behaviour of UK tourism demand. A plausible explanation for this fact is that the general dynamic model, being sufficiently robust to track accurately the UK demand behaviour over the sample period, provides statistically reliable information. Being so, the fact that it indicates that UK tourists adjust very fast to changes in their demand determinants, implies that short run coefficients should be either non-significant or have irrelevant magnitudes. Indeed, this is the general indication of the estimates obtained with the dynamic AIDS. Hence, the faster consumers adjust their demand behaviour, the less significant 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.

The empirical results provided show how the estimation of a dynamic AIDS system provides new information about tourism demand behaviour. This approach allows for intertemporal rationality of consumer behaviour by explicitly considering the underlying short run adjustment mechanism. Reliable estimates for price and budget elasticities were obtained permitting a comparative analysis of the magnitudes and statistical relevance of long and short run sensitivity of UK tourism demand to changes in its determinants.

Nevertheless, there are theoretical and empirical issues which still have to be addressed to endorse AIDS models as dependable specifications. An important matter in quality evaluation of econometric models is their forecasting ability. Statistical models can be good means to describe short and 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 a following up research work.

A Appendix

A.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 nonnegative quantities q_i at given prices p_i , $i = 1, \dots, n$. Let $q = (q_1 \ q_2 \ \dots \ q_n)$ be the quantities vector of the n goods purchased, and $p = (p_1 \ p_2 \ \dots \ p_n)$ 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 maximise the utility, subject to the budget constraint

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

The solution for this maximisation problem leads to the Marshallian (uncompensated) demand functions $q_i = g_i(p, x)$. Alternatively, consumer's problem can be defined as the minimum total expenditure necessary to attain a specific utility level u^* , at given prices

$$\min \sum_{i=1}^n p_i q_i, \text{ subject to } u(q) = u^*. \quad (\text{A.2})$$

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

$$C(p, u) = \sum_{i=1}^n p_i h_i(p, u) = x. \quad (\text{A.3})$$

Given total expenditure x and prices p , the utility level u^* is derived from the problem solution stated in (A.1). Solving (A.3) for u , an indirect utility function is obtained such that $u = v(p, x)$.

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 summarised 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 (A.3) for u , the indirect utility function is obtained, such that $u = v(p, x)$; and finally, $h_i(p, v(p, x)) = g_i(p, x)$ is retrieved stating the Hicksian and the Marshallian demand functions as equivalent.

The Hicksian and Marshallian demand functions have the following properties:

- adding-up: $\sum_{i=1}^n p_i h_i(p, u) = \sum_{i=1}^n p_i g_i(p, x)$ (all budget shares sum to unity);
- homogeneity: $h_i(p, u) = h_i(\theta p, u) = g_i(p, x) = g_i(\theta p, \theta x)$, for all $\theta > 0$ (a proportional change in all prices and expenditure has no effect on the quantities purchased);
- symmetry: $\partial h_i(p, u)/\partial p_j = \partial h_j(p, u)/\partial p_i$ for all $i \neq j$ (consumer's choices are consistent);
- negativity: the $(n \times n)$ matrix of elements $\partial h_i(p, u)/\partial p_j$ is negative semi-definite, that is, for any n vector ξ , the quadratic form $\sum_{i=1}^n \sum_{j=1}^n \xi_i \xi_j \partial h_i(p, u)/\partial p_j \leq 0$, meaning that a rise in prices results in a fall in demand as required for normal goods.

The AIDS model specify the cost function

$$\ln C(p, u) = a(p) + ub(p), \quad (\text{A.4})$$

where $a(p) = \alpha_0 + \sum_{i=1}^n \alpha_i \ln p_i + 2^{-1} \sum_{i=1}^n \sum_{j=1}^n \gamma_{ij} \ln p_i \ln p_j$ and $b(p) = \beta_0 \prod_{i=1}^n p_i^{\beta_i}$. The derivative of (A.4) with respect to $\ln p_i$ is

$$\frac{\partial \ln C(p, u)}{\partial \ln p_i} = \alpha_i + \sum_{j=1}^n \gamma_{ij} \ln p_j + u \beta_i \beta_0 \prod_{i=1}^n p_i^{\beta_i}. \quad (\text{A.5})$$

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

$$\ln x = a(p) + ub(p). \quad (\text{A.6})$$

Solving (A.6) for u , we obtain

$$u = \frac{\ln x - a(p)}{b(p)}. \quad (\text{A.7})$$

Substituting (A.7) in (A.5), we have

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

Defining a price index P such that $\ln P = a(p)$, we have

$$\frac{\partial \ln C(p, u)}{\partial \ln p_i} = \alpha_i + \sum_{j=1}^n \gamma_{ij} \ln p_j + \beta_i (\ln x - \ln P)$$

or

$$w_i = \alpha_i + \sum_{j=1}^n \gamma_{ij} \ln p_j + \beta_i \ln \left(\frac{x}{P} \right), \quad (\text{A.8})$$

where

$$\ln P = \alpha_0 + \sum_{k=1}^n \alpha_k \ln p_k + 2^{-1} \sum_{k=1}^n \sum_{l=1}^n \gamma_{kl} \ln p_k \ln p_l. \quad (\text{A.9})$$

The equations (A.8) and (A.9) 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 tourism budget allocated to the set of n destinations. This share variability is explained by tourism prices p in i and alternative destinations j and by the per capita expenditure x allocated to the set of destinations, deflated by price index P . The model has the following properties:

- the adding-up restriction requiring that all budget shares sum up to unity: $\sum_{i=1}^n \alpha_i = 1$, $\sum_{i=1}^n \beta_i = 0$, $\sum_{i=1}^n \gamma_{ij} = 0$, for all j ;
- the homogeneity restriction requiring that a proportional change in all prices and expenditure has no effect on the quantities purchased: $\sum_{j=1}^n \gamma_{ij} = 0$, for all i ;
- the symmetry restriction requiring consumer consistent choices: $\gamma_{ij} = \gamma_{ji}$, for all i, j ;
- the negativity restriction requiring that a rise in prices result in a fall in demand, that is, 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 (A.9) 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 (A.9) can be substituted by the Stone price index $\ln P^*$ such that

$$\ln P^* = \sum_{i=1}^n w_i^B \ln p_i, \quad (\text{A.10})$$

where w_i^B is the budget share of destination i in the base year. With this simplification for P , the system of equation (A.8) can be rewritten and estimated in the following form

$$w_i = \alpha_i^* + \sum_{j=1}^n \gamma_{ij} \ln p_j + \beta_i \ln \left(\frac{x}{P^*} \right). \quad (\text{A.11})$$

A.2 Expenditure, own-price and cross-price elasticities

Expenditure and price elasticities cannot be directly accessed in (A.11), given its linear-log form. Nevertheless, the elasticities values can be retrieved from the coefficients in (A.11), using the following formulae:

- expenditure elasticity

$$\epsilon_i = \frac{1}{\bar{w}_i} \frac{\partial w_i}{\partial \ln x} + 1 = \frac{\beta_i}{\bar{w}_i} + 1,$$

- uncompensated own-price elasticity

$$\epsilon_{ii} = \frac{1}{\bar{w}_i} \frac{\partial w_i}{\partial \ln p_i} - 1 = \frac{\gamma_{ii}}{\bar{w}_i} - \beta_i \frac{w_i^B}{\bar{w}_i} - 1,$$

- uncompensated cross-price elasticity

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

- compensated own-price elasticity

$$\epsilon_{ii}^\bullet = \epsilon_{ii} + w_i^B \epsilon_i = \frac{\gamma_{ii}}{\bar{w}_i} + w_i^B - 1,$$

- compensated cross-price elasticity

$$\epsilon_{ij}^\bullet = \epsilon_{ij} + w_i^B \epsilon_i = \frac{\gamma_{ij}}{\bar{w}_i} + w_j^B,$$

where \bar{w}_i is the sample average share of destination i ($i = 1, \dots, n$) and w_j^B is the share of destination j ($j = 1, \dots, n$) in the base year.

A.3 The AIDS model of the UK tourism demand for France, Spain and Portugal

The AIDS model assumes that consumers allocate their budget 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 France, Spain and Portugal 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 budget to tourism and other goods; then to tourism in France, Spain and Portugal and other destinations; finally they decide between France, Spain or Portugal. The AIDS system is applied to this last stage using the following form

$$\begin{cases} WF_t = \alpha_F + \gamma_{FP}PP_t + \gamma_{FS}PS_t + \gamma_{FF}PF_t + \beta_F E_t + uF_t \\ WS_t = \alpha_S + \gamma_{SP}PP_t + \gamma_{SS}PS_t + \gamma_{SF}PF_t + \beta_S E_t + uS_t \\ WP_t = \alpha_P + \gamma_{PP}PP_t + \gamma_{PS}PS_t + \gamma_{PF}PF_t + \beta_P E_t + uP_t \end{cases} .$$

A.4 Variables definition

The variables integrating the AIDS model of UK tourism demand for France, Spain and Portugal 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 , $i = F$ (France), S (Spain) and P (Portugal), is defined as

$$Wi = \frac{EXP_i}{EXP_F + EXP_S + EXP_P},$$

where 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

$$Pi = \ln \left(\frac{CPI_i}{CPI_{UK} R_i} \right),$$

where CPI_i is destination i consumer price index, CPI_{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

$$E = \ln \left(\frac{\sum_{i=1}^n EXP_i}{UKP P^*} \right),$$

where UKP is UK population and $\ln P^*$ is the Stone index defined in equation (A.10).

A.5 Data sources

The data for UK tourism expenditure, disaggregated by destinations and measured in £ million sterling, were obtained from the Business Monitor MA6 (1970-1993), continued as Travel Trends (1994-1998). Data on the UK population, price indexes and exchange rates were obtained from the International Financial Statistics Yearbooks, International Monetary Fund (1980-1998).

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