MODELLING IMPORT DEMAND SYSTEMS WITH NONSTATIONARY DATA: AN APPLICATION TO THE FRENCH IMPORTS OF VIRGIN OLIVE OIL

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Abstract

This paper aims to provide a flexible methodological framework to estimate import demand models, which explicitly considers the stochastic properties of data and the endogenous/exogenous nature of some variables. The French imports of virgin olive oil have been used as a case study with Spain, Italy and the Rest of the World as main suppliers. The methodological framework starts by the specification a reduced-form VAR. Appropriated exogeneity tests show the exogeneity of Total Real Imports, indicating the appropriateness of estimating a conditional model. Two cointegration relationships have been found. Several restrictions have been tested in order to identify them as AIDS equations. From structural coefficients of the restricted cointegrated vectors expenditure, own- and cross-prices elasticities are computed. Results show the leadership of Spanish exports to the French market. Italian exports compete in the French market with the Spanish exports, being highly dependent on Spanish domestic production conditions.

Key words: Virgin olive oil, France, demand for imports, cointegration, and exogeneity.

1 Introduction

Since the Armington's (1969) seminal work, a number of papers have dealt with the analysis of the existing degree of substitutability or complementarity's among different exporters of a specific product towards a geographical area. The Armington model is based on a weakly separable utility function that assumes a two-stage process in consumers' purchase decisions. In the first step, the total import quantity of a product is determined under the assumption that imports of a specific good are separable from other imports. Also, in most of the studies, it is usually assumed that the demand for the imported good is separable from that coming from domestic production (Honma, 1993; Lin et al. 1991; Agcaoili-Sombilla and Rosegrant, 1994; Muñoz, 1994; and Yang and Koo, 1994; among others). In the second step, total imports of a specific product are allocated among competing import supplies of different sources of origin. Such a model specification is then implemented by assuming that import supplies from different origins are imperfect substitutes among each other. Taking into account this two-step procedure as well as the separability hypothesis, the import demand function of a specific product can be expressed as a function of import prices from the most important supplier countries and total imports of that product.

In this context, it is very useful in step two to specify a demand system. In his original paper, Armington specified and estimated a Linear Expenditure System. This model has been criticised due to its restrictive assumptions: i) unitary elasticities with respect to the total import quantity demanded for the specific product under analysis; and ii) the constant elasticity of substitution. Alston et al. (1990) showed that the imposition of such restrictions could lead to biased elasticities, since some relevant variables had been omitted. Also, they carried out a number of tests (parametric and non-parametric), concluding that the Armington's assumptions were not corroborated by the data. Moreover, from an empirical point of view, most of the studies that have estimated the Armington model econometrically have obtained rather low estimated values of the elasticity of substitution among imported sources of supplies. In their review of the topic, McDaniel and Balisteri (2003) observed that the following robust findings emerge across many reviewed studies: i) long run estimates of the elasticity of substitution are higher than their short run counterparts, ii) the more disaggregated the data sample is, the higher the elasticity of substitution, iii) cross sectional studies generate estimates that are higher than those provided by time series data, and iv) parameter estimates are sensitive to model misspecification (i.e. endogeneity of explanatory variables, underlying theoretical model structure etc.).

Over the last twenty years, a wide range of solutions has been implemented to overcome the weaknesses of the Armington model. To overcome the homotheticity and the constant elasticity of substitution restrictions, authors started to use more general functional forms and/or models that could account for non-homogeneity, and varying elasticities of substitution, simultaneously. Hence, following the seminal paper of Winters (1984), a long list of econometric studies was published, dealing with the estimation of import demand models by geographical sources using flexible functional forms such as AIDS, Rotterdam, translog, generalized Leontief and normalized symmetric quadratic functional forms, etc. Among them, AID system (Deaton and Muellbauer, 1980) has been the most frequently used due to its easy estimation as well as its flexibility in testing all theoretical restrictions (homogeneity and symmetry).

However, an AIDS specification is constructed within a static framework, includes an assumed endogenous-exogenous division of variables and usually employs non-stationary time series for its parameter estimation. When dealing with non-stationary data, failure to establish cointegration often means the non-existence of a steady state relationship among the variables. Hence, estimation results obtained with static AIDS models can be deemed spurious and statistical inference invalid, if the usual assumption of exogenous regressors does not hold and/or no cointegration relationships exist. Thus, there seems to be a risk involved in the estimation of static system with non-stationary data, which regress endogenous variables on assumed exogenous variables, if their statistical validity is not sanctioned with appropriate testing and cointegration analysis.

When analysing the existence on long-run relationships among non-stationary series, and there are doubts about the exogenous nature of some regressors, one appropriate modelling strategy consists of starting by treating all variables as endogenous within a reduced-form VAR. Next, exogeneity tests for the set of variables in doubt can be carried out. Once the endogenous-exogenous division is established, the reduced rank test can be used to establish the number of cointegrated vectors. Then, estimated of the long-run coefficients can be assessed by imposing exactly-identifying restriction to the VAR. After identifying the VAR structural form, additional restrictions can also be implemented to test its compatibility with specific theories.

We apply this methodological approach to analyse the French imports of virgin olive oil from three main sources: Spain, Italy and the rest of the world. This case study is one of the outcomes obtained from the EU FP6 project “MEDFROL”[1], in which one of the main objectives was to assess the price competitiveness and the export performance of Mediterranean products in the EU. Furthermore, the French market is very appealing as it is the main consumer country of virgin olive oil among non-producer countries within the EU. During the last decade, French imports of virgin olive oil accounted between 12 to 15% of total EU imports, being the second largest EU importer, although still very far from Italy (55 to 60%) (EUROSTAT, 2007).

The paper is structured into the following sections. Section 2 provides some descriptive statistics about the evolution of French imports of virgin olive oil. The methodological approach is developed through sections 3 and 4 together with the discussion of main results. The paper finishes with some concluding remarks.

2 The olive oil market and the French imports

Olive oil is not a homogeneous product (EU Commission, 2004). There are currently several categories of olive oil in the market: virgin oils (mechanically extracted direct from the olives), which comprises the "extra virgin" and "virgin" classes (which are ready for consumption) - and lampante olive oil (which has to be refined); "Composed" olive oil is a blend of refined and "virgin" or "extra virgin" olive oil; and, finally, the olive pomace oil, which consists of a blend of refined olive pomace (residue from the mechanical extraction) oil and "virgin" or "extra virgin" olive oil. Although in this

[1] http://ec.europa.eu/research/fp6/ssp/medfrol_en.htm

section, in some cases, we are going to deal with olive oil, in general, most of our analysis, as well as our empirical work, will concentrate on virgin oils, excluding the lampante oil, that is, only the high quality categories ready for consumption.

The Community is the dominant player on the olive oil market. However, until 1981 it was a net importer as its 425,000 tonnes accounted for only one third of world production. In 1986, after the accession of Greece (1981), Spain and Portugal, the EU became the market reference, averaging 80% of world production. The 1990s saw a rapid rise in EU production as a result of increases in acreages and yields. Compared with harvests in the early 1990s the average production for the last three marketing years doubled in Spain, while Italy and Greece recorded increases of 16% and 18%, respectively. Production in Portugal was fairly stable whereas French production, although very modest in relation to the total for the Community (0.16%), went up slightly. Spain is the world leader producer, accounting for about 35% of world production during the last three marketing seasons. Italy is next, with about 30%, followed by Greece, with around 16%.

As olive oil tends to be consumed in production areas, external trade represents an average of less than 20% of world production. Since mid 1990s, world olive oil exports significantly grew. Italian and Spanish exports - which represent 90% of the total for the EU as a whole – almost doubled. Greek exports, after falling in the mid-1990s, rose by 30%. In terms of categories Greek exports essentially consist of extra virgin olive oil (73% in 2001/02), whereas the figures for Italy and Spain are 45% and 44%, respectively (EU Commission, 2004). In terms of market preparation, all of Greek exports and 91% of Italian exports are in small immediate containers. Exports in bulk represent an appreciable share of Spain's exports (35%), however.

As can be observed in Figure 1, the EU, although being a net exporter, is also one of the world's leading importers of olive oil. Unlike its exports, the EU imports are fairly stable, with specific changes brought about by differences in production. Reduced levels of imports correspond to years in which world output was low or in which the EU production was very high. Conversely, high levels of imports correspond to years in which Community production was relatively small (EU Commission, 2004).

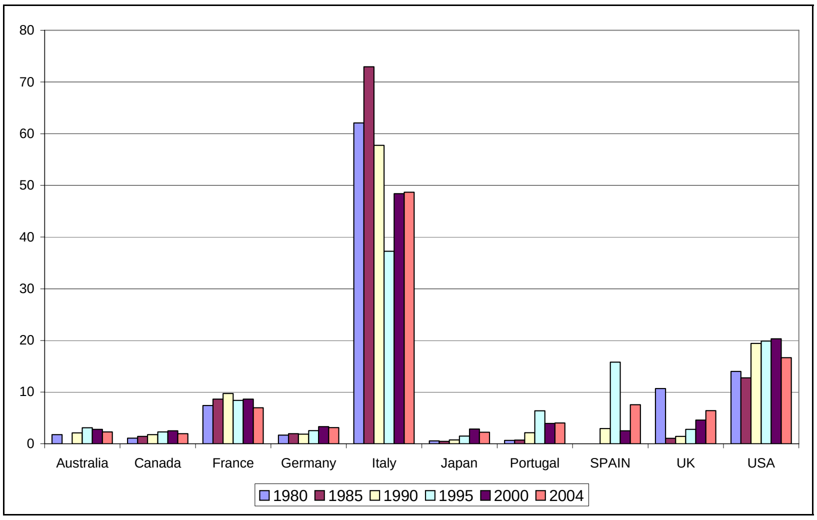

Figure 1. Market share of main olive oil importers in the world (%)

Source: FAOSTAT

Italy tends to account for the bulk of the Community's imports. However, a very interesting point in UE olive oil trade is the inward processing arrangements. Under inward processing arrangements import duty and other commercial policy measures are waived when products are imported from non-member countries for re-exportation in the form of finished products after processing within the Community. Under "by equivalence" inward processing arrangements the importer must export an equivalent quantity of processed olive oil, but not necessarily the actual goods that were processed. They play a major role in the context of Community imports, mainly in the case of Italy, accounting for 60-80% of the total volume of imports.

France is the main importer country within the EU among non-producers, accounting for 12 to 15% of total EU imports of virgin olive oil. French imports mainly come from EU Mediterranean countries (Figure 2). In fact, almost 100% of total imports come from Italy and Spain. However, while the Italy market share has decreased along the last 20 years, the opposite has taken place in the case of Spain. Imports from non-EU countries are marginal. Moreover, Italian prices have been consistently higher than the Spanish ones (Figure 3)

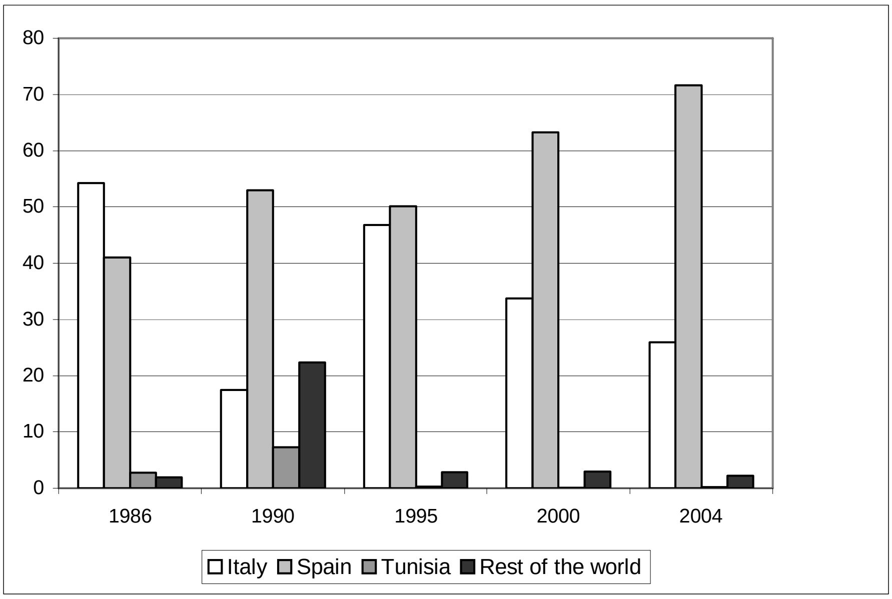

Figure 3. Geographical distribution of French imports of olive oil (%)
Source: FAOSTAT

3 Data sources and preliminary analysis

Data used in this study consists of monthly importation values and quantities to France of virgin olive oil (import unit values as proxies of prices have been obtained by dividing imported values by quantities). The sample period goes from 1995:01 to 2006:12. As mentioned in Figure 2, in the case of France the main sources of imports are Spain and Italy. The other exporting countries have been aggregated under the label “Rest of the World”. Data come from the External Trade Analytical Tables, published by EUROSTAT.

Separability is not an issue here as French production is marginal and almost 100% of olive oil consumed is imported. Moreover, in the case of France the demand for olive oil is independent of that from other vegetable oils (Torres et al., 2004).

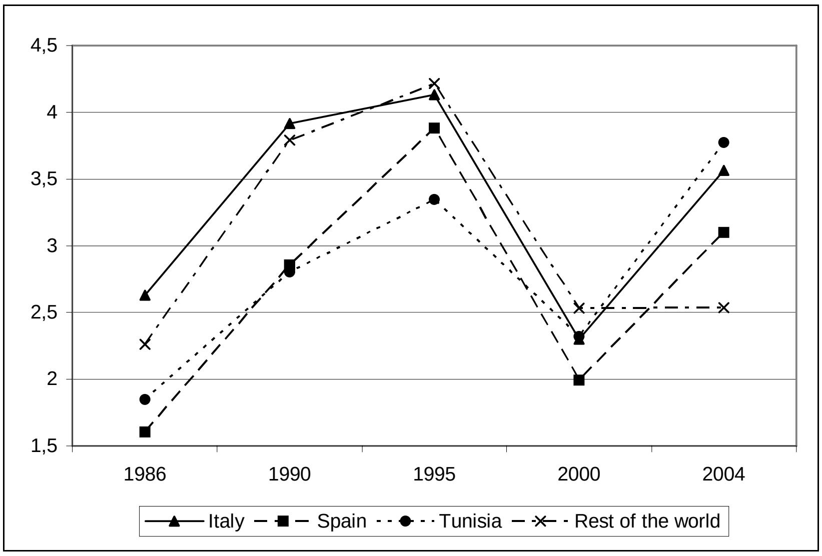

Figure 4. Unit values of French olive oil imports by geographical origin (Euro/Kg)
Source: FAOSTAT

Before proceeding to the specification of the VAR model, it is necessary to analyse the stochastic properties of the data. That is, we must determine if the series are stationary. To identify the number of unit roots for each variable, the more powerful unit root tests, namely, the Modified Generalized-Least-Squares (M-GLS) tests, recently proposed by Perron and NG (1996) and NG and Perron (2001), are used.

Ng and Perron (2001) proposed some modifications of the Phillips and Perron's (1988) (PP) test that are more robust with regards to size distortions when the residual have negative serial correlation. These tests are called M-GLS tests. Like Elliott et al. (1996), who showed that local GLS detrending of the data yielded substantial power gains for the Dickey-Fuller test, Ng and Perron (2001) proposed to follow a similar approach with the PP test. Furthermore, Ng and Perron (2001) showed that the AIC and BIC information criteria were not sufficiently flexible for unit root tests, mainly when there are negative moving-average errors, to select the appropriate number of lags. They therefore suggest the use of a Modified Information Criteria (MIC) that gives better results when we are looking for the appropriate number of lags in the DF-GLS and M-GLS tests. These tests are performed on both levels and first differences of all variables. The inclusion of the trend in the regressions ensures that the unit root hypothesis is not falsely accepted when the series is really trend stationary. Results are presented in Table 1. The unit root tests indicate that all the variables are better characterised as difference rather than trend stationary at the 5% level of significance. For completeness, we have also conducted the Ng and Perron (2001) unit root tests with drift for all variables. In all cases, the null hypothesis of unit root cannot be rejected and the drift terms are significant. Finally, when first differences are used, unit root non-stationarity was rejected at 5% of significance in all cases. This indicates that the level of all variables is non-stationary, i.e. I(1) [2].

Table 1. Results from Ng and Perron (2001) unit root tests of variables included in the import demand system for France

Variables (lags)	$\overline{MZ}_{\alpha}^{GLS}$	$\overline{MSB}_{\alpha}^{GLS}$	ADF^{GLS}
Constant and trend			
RE (1)	-4.6	0.19	-2.82
PS (6)	-11.64	0.20	-2.09
PI (1)	-9.12	0.22	-2.14
PO (1)	-4.52	0.31	-1.49
WS (1)	-7.82	0.23	-2.13
WI (4)	-10.49	0.21	-2.26
WO (4)	-9.02	0.23	-2.11
Critical values (5%)	-17.30	0.168	-2.91
Constant			
RE (6)	-3.57	0.27	-1.75
PS (6)	-2.26	0.46	-1.06
PI (1)	-5.54	0.29	-1.87
PO (1)	-3.27	0.39	-1.30
WS (2)	-1.16	0.47	-0.60
WI (2)	-1.18	0.48	-0.60
WO (5)	-37.86	0.11	-4.27
Critical values (5%)	-8.10	0.23	-1.98

Notes: RE represents the Real Total French Imports of virgin olive oil; PS, PI and PO indicate the corresponding price in Spain, Italy and the rest of the world, respectively. WS, WI and WO denote the market share of Spain, Italy and the Rest of the World, respectively. See Ng and Perron (2001) for a description of the different test statistics.

4 Cointegration analysis

4.1 The VAR model specification and Cointegration rank

Taking into account that all variables in the AID system are I(1), the Johansen's (1988) procedure is used in order to check the possible existence of stationary equilibrium relationships among them. The base-line econometric specification for multivariate cointegration is a VAR(p) representation of a k-dimensional time series vector Y_t reparametrized as a Vector Error Correction Model (VECM):

$$\Delta Y_t = \mu D_t + \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} - \Pi Y_{t-1} + e_t \qquad (1)$$

[2] The KPSS (Kwiatkowski et al., 1992) test also gives the same results.

where, Y_t =[WS_t, WI_t, PS_t, PI_t, PO_t, RE_t]' is a (6x1) column vector of variables[3] (RE represents the Real Total French Imports of virgin olive oil; PS, PI and PO indicate the corresponding price in Spain, Italy and the rest of the world, respectively; and WS, WI and WO denote the market share of Spain, Italy and the Rest of the World, respectively); D_t is a vector of deterministic variables (intercepts, trend...), being μ is the matrix of parameters associated with D_t; Γ_i are (6×6) matrices of short-run parameters (i=1,...,p-1), where p is the number of lags; Π is a (6×6) matrix of long-run parameters and e_t is the vector of disturbances niid(0,Σ).

In empirical applications, the choice of the cointegration rank (r) is frequently sensitive to: i) the deterministic terms included in the system (such as a constant and/or a trend) and the way in which such components interact with the error correction term; and ii) the appropriate lag length to ensure that the residuals are close to being serially uncorrelated. System (1) has been specified with the constant term restricted to the cointegration space[4]. Moreover, based on the Akaike Information Criterion (AIC) and the Likelihood Ratio test (Tiao-Box, 1981), a VECM with four lags is chosen. As a first check of the statistical adequacy of the model, multivariate misspecification tests have been computed. The Godfrey (1988) multivariate autocorrelation and the Doornik and Hansen (1994) multivariate normality tests indicated that the model with two lags had normality problems due to excess Kurtosis. This result is a clear indication that the model is misspecified. This result is a clear indication that the model is misspecified due to the presence of structural breaks. The univariate unit root analysis does not tell us anything about the potential relationship between variables. One cannot use the break dates in the unit root analysis to determine potential breaks in the equilibrium relationships between the variables. Two variables may have individual structural breaks that are important for their univariate behaviour but, if these breaks are common to both variables, they may not be relevant in the functioning of the long-run equilibrium. In this case, it would be better to analyse how the structural change in the variables affects the equilibrium relationships of the system. To this end, we have used the stability tests proposed by Hansen and Johansen (1999).

For VEC models without parameter restrictions and without exogenous variables, the eigenvalues from a reduced rank regression, which are also used in the cointegration rank tests, can be computed recursively by the Johansen method. Hansen and Johansen (1999) propose recursive statistics for stability analysis of VECM. The recursive eigenvalue can be used as the basis for formal tests of parameter constancy. Results of the application of the recursive statistics indicate the existence of a break point in month 2001:02. Then, system (1) has been re-estimated by introducing a restricted dummy variable in the long run, as suggested by Johansen et al (2000):

$$\Delta Y_t = \alpha(\beta', \gamma_1', \gamma_2') \begin{pmatrix} Y_{t-1} \\ D_{1t} \\ D_{2t} \end{pmatrix} + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \mu D_t + \sum_{i=1}^{p} \delta_i d_{t-i} + \varepsilon_t \tag{2}$$

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[3] As three different import sources are considered, only two market share equations are considered due to the adding-up restriction. We omit the share equation for the other of the World (WO).

[4] According to the results in Table 1, the hypothesis that E[ΔY_t]=0 cannot be rejected for all market shares and total imports, indicating that there is no evidence of a linear trend in the data. In any case, following Harris (1995) several tests (subject to the rank restriction on the long-run matrix Π) have been conducted to empirically select the deterministic component introduced in the model. Results clearly indicate that a model with a restricted constant is statistically preferred.

where d_t is a dummy variable which takes the value one for t=2001:02 and zero otherwise, D_t is given by $D_t' = (D_{1t}', D_{2t}')$, where D_{1t}=1 for $\mathrm{p}+1 \leq \mathrm{t} \leq 2001:02$ and zero otherwise and D_{2t}=1 for $2001:02+\mathrm{p}+1 \leq \mathrm{t} \leq 2006:12$ and zero, otherwise.

Multivariate tests for autocorrelation and normality, in this case, indicated the statistical adequacy of the model (33.88 and 14.73, with critical values of 50.71 and 21.03, respectively, at the 5 per cent level); thus allowing us to apply reduced rank tests.

The structure of the AIDS model, to analyse the French imports of virgin olive oil, specifies the shares WS, WI and WO, as the only endogenous variables. Changes in these variables are explained by a set of assumed exogenous regressors including import prices (PS, PI and PO) and Total Imports (RE). However, there seems to be no obvious theoretical or empirical basis challenging the multi-stage budgeting process underlying the rationality of an AIDS expenditure share system, which sets variable RE as an exogenous determinant of the demand shares. The reasons are as follows. In a VAR, all variables are assumed to endogenous implying that a bi-directional cause-effect relationship between RE and shares should exist. In the import demand context, however, even if it reasonable to consider that changes in RE affect import shares, it does not seem realistic to expect that changes in these shares influence the way in which French consumers allocate their budgets. In addition, the estimation results of the RE_t equation in the Unrestricted VAR model indicated that only their own lags were significant while the rest of the variables remained not to be statistically significant[5]. Thus, in the following we are going to assume RE to be exogenous.

When exogenous variables are considered, the Y_t vector can be partitioned as $Y_t = (Z_t,', X_t ')'$, where Z_t,=[WS,WI,PS,PI,PO]' is an (5x1) vector of endogenous variables and X_t =RE is the exogenous variable, which can be considered as the "long-run forcing" variables in the system, that is, changes in X_t have a direct influence on the variables Z_t, while they are not affected either by the changes in the equilibrium relationships nor by past changes in Z_t. This is equivalent to the notion that the set of variables Z_t do not Granger-cause X_t.

If the variables in X_t are not cointegrated, Pesaran et al. (2000) show that the k-variable system defined in (2) can be decomposed to following two subsystems:[6].

Conditional subsystem: $$\Delta Z_t = \Lambda X_t + \alpha(\beta', \gamma_1', \gamma_2') \begin{pmatrix} Y_{t-1} \\ D_{1t} \\ D_{2t} \end{pmatrix} + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \mu D_t + \sum_{i=1}^{p} \delta_i d_{t-i} + \varepsilon_t \qquad (3)$$

Marginal subsystem: $$\Delta X_t = \mu_x + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta X_{t-i} + \varepsilon_{xt} \qquad (4)$$

[5] Results are not presented due to space limitations but are available from authors upon request.

[6] Under this decomposition, variables in X_t are assumed to be weakly exogenous with respect to the cointegration space. Moreover, if the variables in Z_t do not Granger-cause X_t, then these variables are assumed to be strongly exogenous with respect to the cointegration space, that is, they would be only explained by their own past in the marginal subsystem.

“Variables cannot be exogenous per se” (Hendry, 1995). A variable can only be exogenous with respect to a set of parameters of interest. Hence, if the variables X_t are deemed to be exogenous with respect to parameters in (3), the marginal model (4) can be neglected and the conditional model (3) is complete and sufficient to sustain valid inference. Hence, the knowledge of the marginal model will not significantly improve the statistical or forecasting performance of the conditional model. Following this line of reasoning, the conditional model is used to test for cointegration, which is equivalent to testing for the Rank (r) of matrix Π_z.

To supply further empirical support for our claim that feedback effect might be absent in the relationships between X_t and Z_t, we use the causality concept proposed by Granger (1969). The block Granger non-causality test is a multivariate generalisation of the Granger causality test that can be used to establish if one or more variables should or should not integrate the set of endogenous variables in the VAR. The LR statistic to test the null of zero-value coefficients of Z_t,=[WS,WI,PS,PI,PO]’ in the RE equation was $\chi^2(20) = 12.59$, indicating that the null cannot be rejected at the 5% level. As mentioned, this result reinforced our assumption of excluding RE from the set of endogenous variables in the VECM and the conditional model (3) is used to test for cointegration.

Once the correct model has been specified, the next step consists of determining the number of cointegrated vectors using cointegration rank tests. Table 2 shows the results from the trace cointegration statistic. As can be observed, we cannot reject the presence of two cointegrating relationships, which is exactly the number of equations we have to estimate. However, as many previous simulation studies have demonstrated, asymptotic distributions are often poor approximations for small sample distributions. Thus, the results from both tests have to be interpreted with some caution. Juselius (1999) proposes an alternative approach in which she tests the significance of the adjustment coefficients for the r-th cointegrating vector (α_{ir}). If all α_{ir} coefficients are non-significant, then the cointegration rank should be reduced to (r-1). In our case, all estimated adjustment coefficients for the second cointegrating vector were significant, confirming the presence of two long-run relationships among the five variables.

Table 2. Results from the cointegration trace statistic

r	p-r	LR	p-value	Critical value (95%)[a]	Critical value (90%)[a]
0	5	124.38	0	88.48	84.25
1	4	77.01	0.0019	63.66	60.06
2	3	40.98	0.0761	42.74	39.80
3	2	22.61	0.1205	25.50	23.26
4	1	9.90	0.1167	11.59	10.22

a Critical values with only a break in levels. The response surface is generated according to Trenkler (2004).

As mentioned previously, the estimation of the VECM, subject to the rank restriction on the long-run matrix Π, does not generally determine a unique set of cointegrating relationships. Moreover, the two long-run relationships have not the form of AIDS equations as both market shares are included in the model. Then, to identify the long-run equilibrium relationships and theory-compatible AIDS equations, several restrictions should be introduced.

The general expression of the long-run relationships is given by

$$\beta' Y_t = \begin{pmatrix} \beta_{11} & \beta_{12} & \beta_{13} & \beta_{14} & \beta_{15} & \beta_{16} & \beta_{17} & \beta_{18} \\ \beta_{21} & \beta_{22} & \beta_{23} & \beta_{24} & \beta_{25} & \beta_{26} & \beta_{27} & \beta_{28} \end{pmatrix} \begin{pmatrix} WS_{t-1} \\ WI_{t-1} \\ PS_{t-1} \\ PI_{t-1} \\ PO_{t-1} \\ RE_{t-1} \\ D_{1t} \\ D_{2t} \end{pmatrix} \tag{5}$$

If in (5) we introduce the following exact-identifying restrictions $\beta_{11} = 1; \beta_{22} = 1$ and $\beta_{12} = \beta_{21} = 0$ we get the share equations of the AIDS model. Table 3 reports the estimated β parameters as well as their standard errors.

Table 3. Estimated β^a matrices under long-run structural identification

	WS	WI	PS	PI	PO	RE	D_{1t}.	D_{2t}
$\hat{\beta}'_1$	1	0	0.535 (0.295)	-0.710 (0.316)	0.243 (0.122)	-0.131 (0.003)	-0.016 (0.003)	0.0095 (0.003)
$\hat{\beta}'_2$	0	1	-0.609 (0.271)	0.770 (0.290)	0.248 (0.112)	0.105 (0.004)	0.0164 (0.002)	-0.007 (0.002)

[a.] Standard errors for β are given in parentheses

At the 5% level of significance, all coefficients are significant and present the expected signs and magnitudes. These two cointegrating relationships can be interpreted as long-run AIDS equations. In the next step, demand theory restrictions (homogeneity and symmetry) can be directly tested as over-identifying restrictions. Taking into account (5), the homogeneity hypothesis can be formulated as follows:

$$\begin{aligned} &\beta_{13} + \beta_{14} + \beta_{15} = 0, \text{ and} \\ &\beta_{23} + \beta_{24} + \beta_{25} = 0 \end{aligned} \tag{6}$$

The Likelihood Ratio (LR) statistic for testing the three over-identifying restrictions is 4.12, which is below the 5% critical value of χ^2 (2) (=5.99). Thus, the homogeneity hypothesis cannot be rejected. Second, and in order to be consistent with economic theory, we have carried out a joint test of the symmetry and homogeneity restrictions. The symmetry hypothesis requires the introduction of the following cross-equation restrictions:

$$\beta_{14} = \beta_{23} \tag{7}$$

The LR statistic for jointly testing the restrictions imposed[7] in (6) and (7), which is asymptotically distributed as $\chi^2(3)$ (the critical value is 7.81 at the 5% significance level), was 4.23 suggesting that both theoretical restrictions cannot be rejected. Table 4 reports the estimated long-run parameters with the restrictions of homogeneity and symmetry being imposed restrictions (asymptotic standard errors in parenthesis).

Table 4. Estimated β^a matrices under long-run structural identification

	WS	WI	PS	PI	PO	RE	D_{1t}.	D_{2t}
$\hat{\beta}'_1$	1	0	0.340	-0.323	0.017	-0.101	-0.547	0.065
			(0.045)	(0.023)	(0.058)	(0.057)	(0.099)	(0.029)
$\hat{\beta}'_2$	0	1	-0.323	0.317	0.006	0.095	-0.570	-0.60
			(0.013)	(0.015)	(0.056)	(.054)	(0.077)	(0.028)

[a.] Standard errors for β are given in parentheses

4.2 Elasticities

From estimated parameters in Table 4, we have calculated the corresponding elasticities. However, in this case, as a transitory dummy has been introduced, the calculated elasticities are period specific. We have assumed that the structural change only affects the intercepts, that is, the changing behaviour after 2001 only has modified the average market shares. On the other hand, the effects of changes in prices or changes in the total imported value have remained more or less stable over time.

Table 5 shows the calculated import demand elasticities before and after the structural change. As can be observed elasticities in both periods are rather similar and can be jointly discussed. In both periods, expenditure elasticities of the two main exporters (Spain and Italy) are positive and significant. However, only in the case of imports coming from Spain is the elasticity higher than one. This predicts that Spain will increase its market share if EU imports of virgin olive oil also increase. Since the structural change only affects the intercepts, any differences in expenditure elasticities between the two periods are due to changes in market shares. In any case, a formal test has been carried out to check if the elasticities before and after the structural change are statistically different. Results indicate that no significant differences have been found.

Table 5. Long-run Demand Elasticities for French imports of virgin olive oil:

	Expenditure Elasticities	Price Elasticities (t-ratios)[a]		
		Spain	Italy	Other
1995-2000				
Spain	1.140*	-1.57*	0.70*	0.16
Italy	0.629*	1.98*	-2.14*	-0.34
Other	0.727	1.49	-0.016	-1.73
2001-2005				
Spain	1.152*	-1.61*	0.79*	0.177

[7] In this case, the restricted model subject to non-linear restrictions is estimated using the non-linear switching algorithm in PcFiml version 9.0.

Italy	0.696*	1.69*	-1.92*	-0.27
Other	0.760	1.34	0.072	-1.65

[a] Marshallian own-price and Hicksian cross-price elasticities.

Own-price elasticities are negative in both periods and elastic with the singular exception being imports coming from the rest of world, which are not significantly different from zero. These results indicate that Spain and Italy could gain market share in the UE through competitive prices. For those countries included in the study, the situation has not changed substantially before and after the structural change (no significant differences have been found).

Compensated cross-price elasticities are also shown in Table 5. Imports coming from Spain and Italy are highly substitutive and compete in prices. The corresponding elasticity is positive and significantly different from zero. Production conditions in both countries are main determinants of their competitiveness in French markets. Exports from the Rest of the World have a marginal importance and they behave as independent in relation to those coming from the two exporters.

5 Concluding remarks

During the last two decades, a large number of papers analysing import demand models have used flexible demand systems, being the AID system the most commonly used. However, most of the studies has neglected the stochastic properties of series involved. If series are non-stationary, the estimation of AIDS static models can lead to spurious results and invalid statistical inference. In some cases, dynamics have been considered by specifying a partial adjustment framework. On the other hand, another set of papers, using the cointegration framework, have not accounted for the possibility of some variables in the AID system to be exogenous. In this paper, we provide a flexible methodological framework to estimate import demand models, which explicitly considers the stochastic properties of data and the endogenous/exogenous nature of some variables.

The French imports of virgin olive oil have been used as a case study. In any case, from an empirical point of view, this is an interesting case as, among non-producers, France is the larger consumer country as it is the most important EU importer. The French market can be considered an adequate test for analysing the price competitiveness of exports coming from traditional producers (i.e. Spain and Italy). Taking into accounts the pattern of French imports of olive oil, three main sources have been considered: Spain, Italy and the rest of the World.

Results from unit root tests have indicated that all variables are non-stationary. Then, a reduced-form VAR with two lags has been specified. Appropriated exogeneity tests have indicated that Total Real Imports can be considered as an exogenous variable. Thus, a conditional subsystem has been estimated. Moreover, as the Italian market share showed a significant decrease during the second half of the sample period (more precisely, since 2000), one transitory impulse dummy has been restrictedly introduced in the long run. To long run relationships have been found from which several tests have been performed to identify a theory-consistent AIDS system. Finally, the structural coefficients of the restricted cointegrated vectors were used to compute the expenditure, own- and cross-prices elasticities. As a transitory dummy has been introduced, the calculated elasticities are period specific. In any case, the effects of changes in prices or changes in the total imported value have remained more or less stable over time.

Results also show the leadership of Spanish exports to the French market. Spanish exports increase (decrease) more than proportional to French imports of virgin olive oil. Imports from third countries are almost irrelevant. Italian exports compete in the French market with the Spanish exports, being highly dependent on Spanish domestic production conditions.

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