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The EU demand for imports of virgin olive oil

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Abstract

This paper has analysed the import demand for virgin olive oil in the EU and more precisely in the Italian market, as it concentrates more than 80% of EU imports, aiming to determine the relative position of Mediterranean EU and non-EU countries exports and their degree of substitutability or complementarity. The methodology used is based on the specification of a Threshold Almost Ideal Demand System in which special attention has been paid to the stochastic properties of the series involved. In an empirical context, the paper has aimed to provide a set of import demand elasticities that can be useful in trade models. Results point to Spain as the leader in the Italian virgin olive oil market. It is expected that this position will be maintained in the future. Greece has improved its relative position after its accession into the EU. However, imports coming from Greece are highly dependent on the situation in Spain. Tunisia has good potential for future exports development as a consequence of new perspectives of trade liberalisation taking into account its relative position in the Italian market, in spite that its exports are currently constrained due to existing quotas.

Key words: Olive oil, Italy, elasticities, imports, TAIDS

The EU demand for imports of virgin olive oil

1. Introduction

Olive trees have a long tradition in Mediterranean countries as we have noticed that in the Roman Empire olive-growing was common practice. Nowadays, the Mediterranean basin concentrates around 98% of the world's olive trees and accounts for the bulk of world olive oil production. The EU uses to be not only the world's largest market for Mediterranean products but it also remains the prime outlet for the southern Mediterranean countries' exports on which their national economies depend to a great extent. Imports of olives and olive oil, however, are not very substantial primarily because of the leading position that Spain, Italy and Greece hold within the EU market. However, production in non EU Mediterranean countries is still high. Tunisia and Turkey are the world's fourth and fifth largest olive producing countries respectively, while Tunisia is a prime olive oil supplier for the EU.

The EU has conceded a number of trade privileges either directly to certain countries or to the whole area (Barcelona Agreement in 1995 for the creation of a Free Trade Zone in the Mediterranean basin); thereby allowing for closer trade cooperation with these countries. This Trade Zone, free from duties and other import barriers will affect both the non-EU Mediterranean countries, that will be allowed easier entrance to a large market, and the southern EU Member States (mainly Greece, Italy and Spain) that produce similar products and will be faced with increased competition that might lead to lesser market shares in a previously 'exclusive' market.

The main objective of this paper is to assess, by conducting import demand analyses, the price competitiveness and the export performance of the olive oil sectors in the Mediterranean regions. Moreover, it aims to providing results that may prove helpful for decision-making at the Community level at a time that the changing environment, both intra-EU (CAP reform, EU enlargement) as well as extra-EU (WTO negotiations, ongoing globalisation and liberalisation of international markets) urges for detailed knowledge regarding potentials and future market trends.

The literature dealing with the estimation of import demand elasticities covers quite a gamut of perspectives. Some of the studies adopt a macro-economic point of view where the object is to forecast the evolution of principal trade flows and to evaluate the impact of exchange rate changes on the trade balance (Sarris, 1981). The data used in these studies usually includes the value of total imports from a specific country. However, when the focus of the study is to analyse the existing degree of substitutability or complementarity among different exporters of a specific product towards a geographical area, more specificity is needed.

Towards this end, one of the seminal and more popular models was based on Armington (1969). 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. It is assumed 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 of different origin 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 to specify a flexible full demand system for the model in step two. Among them, the AIDS system, applied for the first time by Winters (1984) to import demand analyses, has been one of the most frequently used due to its easy estimation as well as its flexibility in testing all theoretical restrictions (homogeneity, symmetry and negativity). This is the system that we are going to use in this paper.

However, from a methodological point of view, this paper presents two main novelties. First, it paper constitutes one of the first attempts to explicitly consider the stochastic properties of data, that is, if series are non-stationary and cointegrated. And second, up to know, the empirical literature has assumed linear adjustments of imports to price changes. However, this is not necessarily true and non-linearities may be present depending on the behaviour of domestic prices. In other words, trade patterns may be different depending on price differentials between domestic and foreign markets. To tackle with this issue, in this paper a Threshold Almost Ideal Demand System (TAIDS) is specified and estimated, which is the main contribution to the existing literature on import demand models.

To achieve this objective, the paper has been structured into the following sections. Section 2 presents some descriptive statistics about olive oil trade patterns in the Mediterranean basin. The theoretical background is presented in Section 3. Section 4 describes the data series used as well as their stochastic properties, which have ultimately determined the econometric approach followed in this paper. The TAIDS model is specified and estimated in Section 5, as well as the calculated elasticities. The paper finishes with some concluding remarks.

2. Olive oil trade patterns in the EU

The olive oil, however, 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 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 production of olive oil is heavily conditioned by both agronomic and climatic conditions. As much of the cultivated surface is not irrigated, drought periods are particularly harmful for olive trees (i.e. Spain in 1995/96). Moreover, production is determined by alternate bearing, a characteristic of olive trees whereby bumper crops tend to be followed by lower production the following year.

In the last few decades olive oil production has featured periods of growth followed by stagnation (EU Commission, 2004). At the beginning of the 1980s world production was about 1.8 million tonnes, 40% up on the figure recorded in the mid-1960s. After a relatively stable period production again showed an upturn in the second half of the 1990s, to reach 2.5 million tonnes. Average world production for the last three marketing years has been about 2.7 million tonnes (Figure 1).

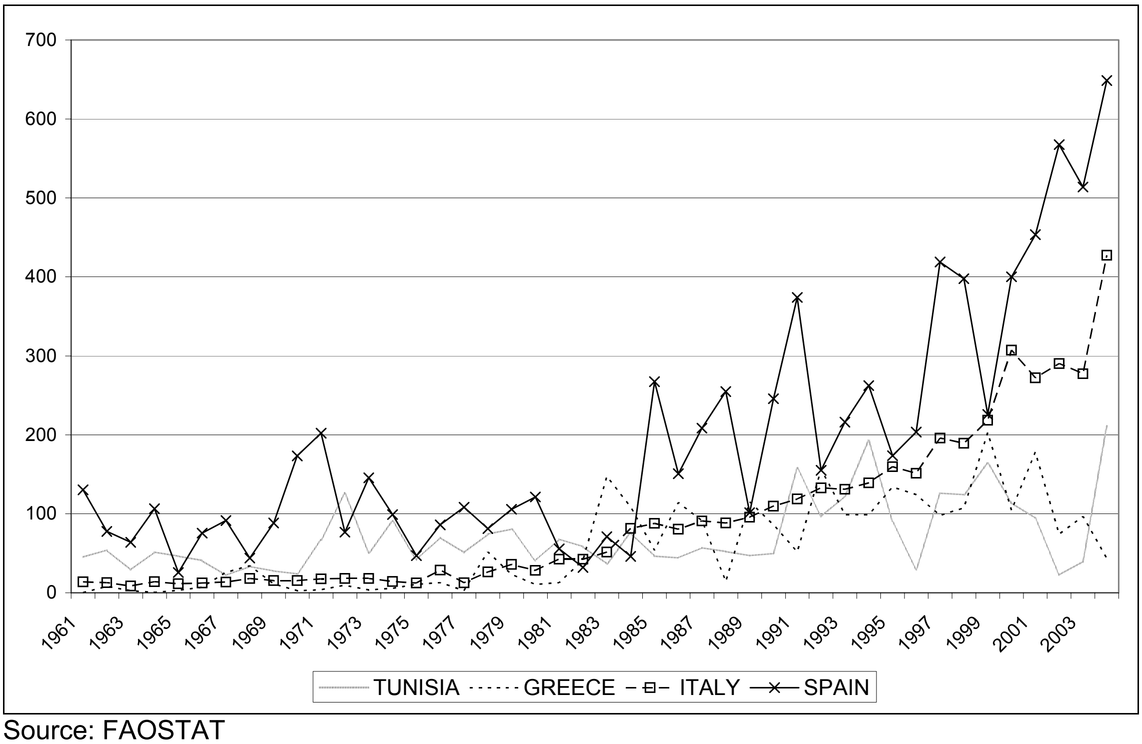
The Community is the dominant player on the olive oil market. Until 1981 its 425 000 tonnes accounted for only one third of world production and it was a net importer. In 1986, after the accession of Greece (1981), Spain and Portugal, the EU became the market reference, averaging 80% of world production. 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 France's production, although very modest in relation to the total for the Community (0.16%), went up slightly. Overall, Community production has gone up 51%. 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%. Among the non-EU Mediterranean countries, Tunisia and Turkey are the main producers, accounting for 6 and 4% of total world production, respectively.

The recent enlargement of the EU has had only a limited impact on Community olive oil production since only three of the new Member States are producers but at a rather small scale. The quotas allocated to them are 6 000 tonnes for Cyprus, 400 tonnes for Slovenia and 150 tonnes for Malta, which together represent 0.4% of the combined national guaranteed quantities of the other Member States.

Since olive oil tends to be consumed in production areas, external trade represents an average of less than 20% of world production. In the beginning of the 1990s the EU accounted for just over half (54.5%) of world exports of olive oil, the corresponding figures for Turkey and Tunisia being 32% and 7.7% respectively. 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 30% (Figures 1).

Figure 1. Evolution of exports from main world suppliers (000 t)

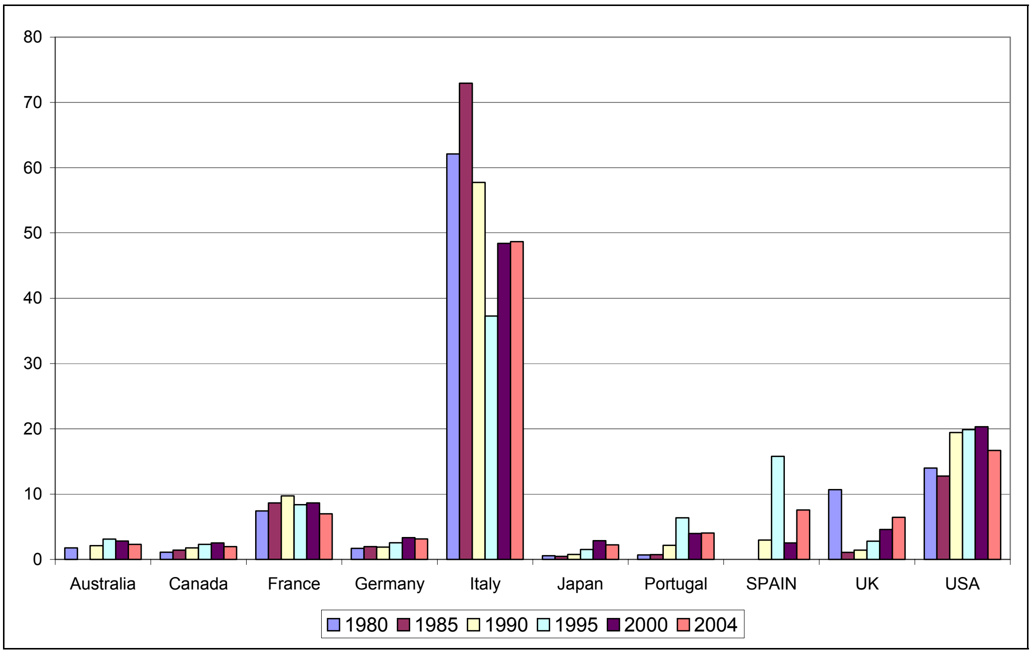


Source: FAOSTAT

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.

Apart from the EU countries (mainly Italy), the United States, Australia, Japan and Canada account for practically all the EU exports (Figure 2) and tend to be in immediate containers of less than 18 kg. The other major exporters to non-producing countries were Turkey (mainly to Canada, the United States, Australia and Japan), Tunisia (mainly to the United States) and Argentina (mainly to Brazil).

Figure 2. Market share of main olive oil importers in the world (%)



Source: FAOSTAT

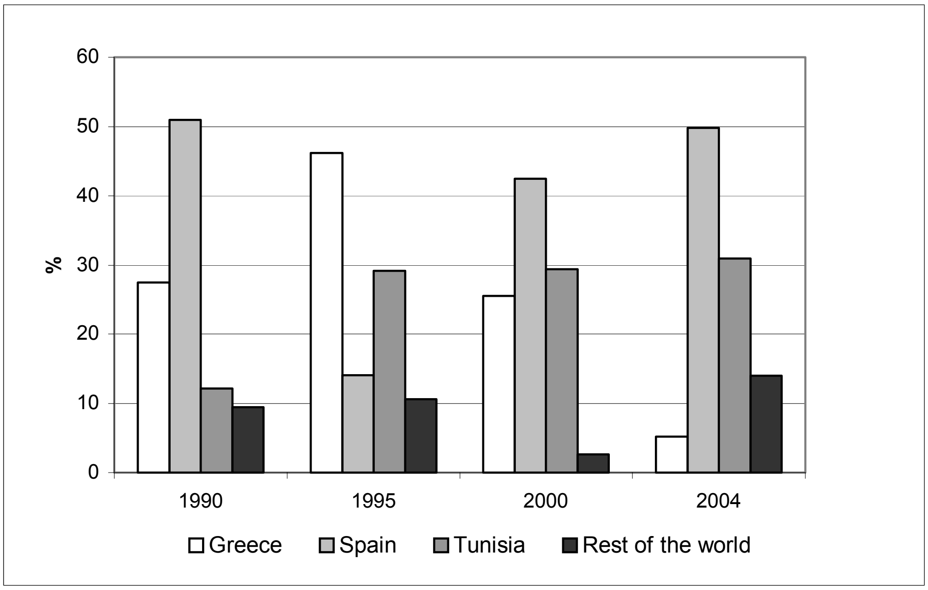
As can be observed in Figure 2, the EU, although a net exporter, is also one of the world's leading importers of olive oil. Unlike its exports, the EU imports are fairly stable, with specific changes brought about by differences in production. Reduced levels of imports correspond to years in which world 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). Italy tends to account for the bulk of the Community's imports: German, Portugal UK and French imports have nearly always been negligible and those of Spain have been only noticeably in 1995 due to the severe drought in that country.

As mentioned above, Intra-Community trade accounts for the bulk of the trade in olive oil. Italy is, by far, the main importer of virgin olive oil, accounting for 80% of total EU imports, followed by France (around 15%), UK, Portugal, Spain and Germany (3%). Generally speaking, Spain and Greece sell oil to the rest of the EU, mainly to Italy. France and Portugal are also major buyers of Spain's oil. Italy buys and sells olive oil within the EU but its purchases tend to exceed its sales. Italy's traditional customers have been Germany (in which it has almost a monopolistic situation) France and the United Kingdom.

Taking into account that more than 80% of total EU imports go to Italy, in this paper we have decided to take Italian imports as representing total EU imports. Moreover, as Italy is an important producer country, this assumption will facilitate us to relate our results to domestic prices in order to specify an appropriate model to calculate import elasticities.

Figure 3 shows the main origins of Italian imports of virgin olive oil. As can be observed, Spain has been traditionally the main supplier, accounting for around 50% of total Italian imports, with the exception of 1995 due to the severe drought that took place in Spain, as mentioned above. Greece is the second main supplier. However, its relative position has varied significantly, depending on the Spanish production. It seems that imports from both countries are highly substitutive. Tunisia is the third main exporter to Italy and, consequently, to the EU. Since 1995, with the EU agreement on zero-rated import quota of 40,000 t (56,000 t, in 2005), Tunisian exports to Italy have been stabilised around 30% of total imports. The rest of the world occupies a marginal position, being Turkey the most outstanding country.

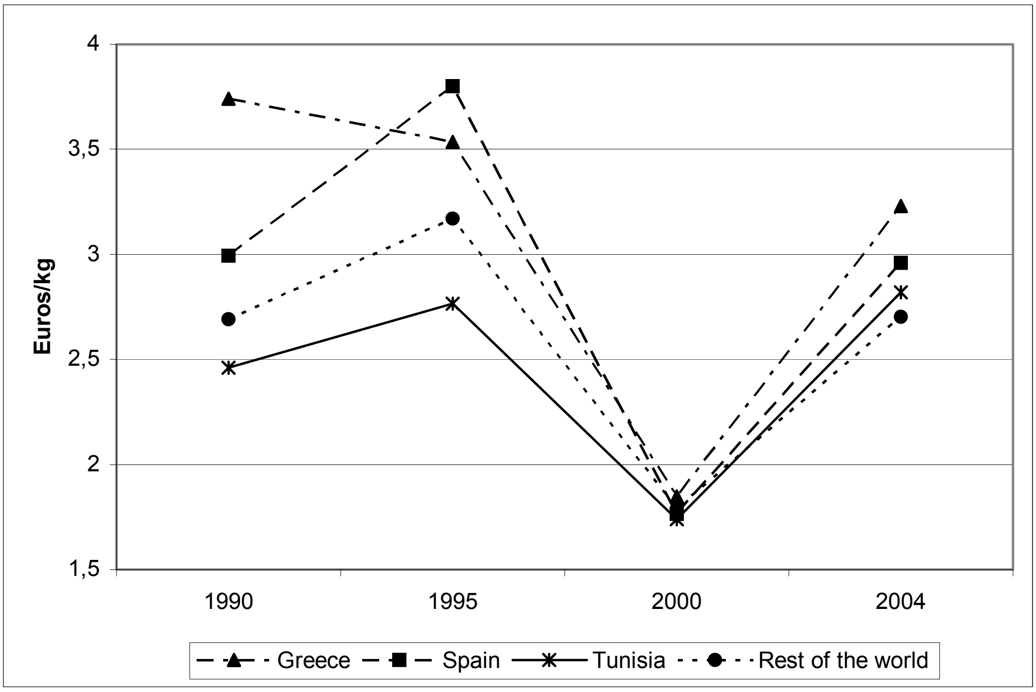
Figure 3. Geographical distribution of Italian imports of olive oil (%)



Source: FAOSTAT

Imports coming from EU countries are more expensive than those coming from non-EU countries (Figure 4). Greek prices are situated slightly above the Spanish prices except in 1995 due to the causes already mentioned. Higher prices are mainly due to the export composition as Greek exports essentially consist of extra virgin olive oil, while in the case of Spain this category only accounts for 44%. In 2000, prices were lower as the EU reached a high production level.

Figure 4. Unit values of Italian olive oil imports by geographical origin (Euro/Kg)



Source: FAOSTAT

3. Theoretical and econometric background

As mentioned in the introduction, in this paper we use a generalisation of the Armington's (1969) model to analyse the olive oil export performance of Mediterranean countries into the EU (Italy). 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.

In relation to the separability assumption, several efforts have been addressed towards allowing the possibility to estimate import demand models that could be at the same time source-differentiated and differentiated by sources of production (Yang and Koo, 1994; Carew et al., 2004). In all these empirical studies, the question of home production and different sources of imports is hardly addressed. This state of affairs is quite understandable in light of the difficulties to obtain comparable price data for domestic production and various sources of import supplies.

This situation is increasingly complicated when more than one product is considered. For instance, for four groups of products and five sources of imports in each group, an unrestricted AIDS model will have 20 equations and 20 x (20+2) = 440 parameters to estimate. Under such circumstances, even the standard assumptions of adding-up, homogeneity and symmetry may not be sufficient to solve the degrees-of-freedom problem.

To reduce the number of parameters to be estimated, Yang and Koo (1994) specify an AIDS model and introduce an assumption of block-substitutability (BLSUB). Contrary to the Armington model, the Yang and Koo's assumption of block-substitutability does not require two-stage budgeting. Expenditures are allocated simultaneously over all products under consideration. This allows for direct cross-price effects among the products belonging to different groups. Their model assumes, however, that while allocating expenditures among different sources of the same good, consumers do not distinguish among sources of other goods.

Thus, the Yang and Koo approach has two major shortcomings. First, because not all symmetry restrictions can be imposed, the gain from block substitutability in terms of degrees of freedom becomes less obvious. Second, the within-group adding-up restriction is not justified by economic theory. In particular, theory does not require that within-group cross-price parameters add to zero.

Soshnin et al. (1999) introduce an improved assumption of block-substitutability (IBLSUB) that makes the source-differentiated AIDS model a better tool for international demand studies as: (a) it is consistent with economic theory; and (b) it further reduces the number of parameters to be estimated. The procedure starts by writing the BLSUB model as nested within the standard LAIDS specification; thus including the same variables on the right-hand side of all equations. Then, they introduced some restrictions based on Hayes et al. (1990).

More recently, Asche et al. (2005) use a different approach, the Generalized Composite Commodity Theorem (GCCT) of Lewbel, to justify aggregation, and therefore the estimation of a demand system of only import demand equations. The main advantage with this approach is that to test whether this theorem holds for a group of goods, one only needs the data that is used when one is estimating a demand system. The GCCT can accordingly be used to easily validate that one can treat the goods in question as a separate group provided that the theorem holds, and without use of additional data that often is not available. Hence, one can use only import data to investigate whether import demand functions can be estimated without taking into account the demand for domestic production of the same good.

Taking this result into account, in this paper import demand analyses is carried out by only estimating the second-step import demand system (following the Armington terminology) and testing for separability using the GCC Theorem. The AIDS flexible functional form (Deaton and Muellbauer, 1980) is chosen for that purpose. This system was first applied to the analysis of the demand for imports by Winters (1984) and has the following expression:

$$w_{it} = \alpha_i + \beta_i \ln\left(\frac{M}{P}\right)_t + \sum_j \gamma_{ij} \ln P_{jt} + \varepsilon_t \quad (1)$$

where: w_{it} represents, in this case, the market share of the i-th country on total imports of a specific product in period t (t=1...T)); p_{jt} is the unit value of imports coming from country j in period t (t=1...T); M_t represents the total value of imports in period t (t=1...T); and log P_t is a price index in period t (t=1...T) defined as:

$$\ln P_t = \sum_{i=1}^{n} \overline{w}_i \ln p_{it} \qquad (2)$$

where: $\overline{w}_i$ is the average market share of imports coming from country i.

The theoretical restrictions of adding-up, homogeneity and symmetry hold if the parameters satisfy the corresponding expressions:

- Adding-up: $\sum_{i=1}^{n} \alpha_i = 1$, $\sum_{i=1}^{n} \gamma_{ij} = 0$ and $\sum_{i=1}^{n} \beta_i = 0$;
- Homogeneity: $\sum_{j=1}^{n} \gamma_{ij} = 0$;
- Symmetry: $\gamma_{ij} = \gamma_{ji}$.

The negativity condition will hold if matrix C with elements c_{ij}:

$$c_{ij} = \gamma_{ij} - \lambda_{ij}\overline{w_i} + \overline{w_i}\ \overline{w_j} \qquad (3)$$

is semi-definite negative, where: λ_{ij} = 1 when i=j, and zero, otherwise; and $\overline{w_i}$ and $\overline{w_j}$ are average market shares of countries i and j, respectively, of total imports. The sufficient condition for matrix C to be semi-definite negative is that its eigenvalues are negative. Otherwise, this condition has to be imposed either following a Bayesian approach (Chalfant et al, 1991; Hasegawa et al, 1999) or using the Cholesky decomposition (Barten and Geyskens, 1975; Moschini, 1998; and Ryan and Wales, 1998; among others).

However, results from the specification and estimation of AIDS systems when dealing with time series data have been criticised from the theoretical point of view due to the frequent rejection of homogeneity, symmetry and negativity restrictions (although we must note that no attention has been paid up to now to the last restriction in import demand models). The rejection of theoretical restrictions has been related to misspecification problems. Deaton and Muellbauer (1980) found that for every commodity group for with the homogeneity restriction was rejected, the imposition of homogeneity generated positive serial correlation in the residuals of the equation fitted. This strongly suggests dynamic mis-specification. Moreover, the disappointing rejection of the homogeneity and symmetry restrictions was shown to be compatible with data when applied to the long run solution of a completely specified dynamic model by Anderson and Blundell (1983, 1984).

Recent developments in non-stationary time series and cointegration techniques have opened an alternative approach for introducing dynamics in demand systems. When non-stationary variables are used, homogeneity and symmetry tests, carried out on the estimated system in level form using least squares, are no longer valid. On the other hand, if the variables are cointegrated, the specification of a demand system in first differences is biased due to the misspecification of the long-run relationships. While the stochastic properties of series have been considered in some studies dealing with traditional demand analyses, this approach has yet to be applied to import demand systems.

Ng (1995) and Attfield (1997) specified a system of variables in triangular form, estimating the system using Dynamic Ordinary Least Squares (Phillips, 1991; and Stock and Watson, 1993) and testing homogeneity with a Wald statistic. Balcombe and Davis (1996) used the canonical cointegrating regression (CCR) (Park, 1992) to estimate an AIDS for food consumption in Bulgaria. In all these papers no attempt was made to identify the cointegrating relationships. It was assumed *a priori* that among the (2n + 1) variables (n budget shares, n prices and real expenditure) there are n-1 cointegrating vectors each of which corresponds exactly to an AIDS equation[1].

Recently, Pesaran and Shin (1999) have used the Johansen's (1988) approach to specify, estimate and test theoretical restrictions on cointegrated demand systems. This approach overcomes most of the above-mentioned criticism. However, it relies on the idea that all variables in the system are non-stationary, which is not always the case. It is not unrealistic to assume that, for specific products, market shares could be quite stable over time[2]. Thus, before econometrically specifying the demand system, it is relevant to analyse the stochastic properties of all series involved in the model. This is precisely the aim of next section.

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[1] One equation is arbitrarily deleted due to the adding-up restriction

[2] On the other hand, from a statistical point of view, market shares are bounded between 0 and 1, so it is expected to be stationary in the long run. However, in some cases they show the typical characteristics of I(1) processes. Thus, we have followed Ng (1995), Attfield (1997) and Pesaran and Shin (1999) and have tested to see if market shares are or are not stationary.

4. Data sources and preliminary analyses

Data used in this study consists of monthly importation values and quantities to Italy 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 2005:12. As mentioned in Figure 3, in the case of Italy the main sources of imports are Spain, Greece 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.

The stationarity of the time-series variables has been analysed using the recently proposed Ng and Perron (2001) unit root tests, namely the Modified Generalized-Least-Squares (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 Phillips and Perron (1988) 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. Results are presented in Table 1. Results, clearly indicate that all price series are I(1). In the case of the four market shares and the total imports, results indicate that these series are stationary around a deterministic trend.

Table 1. Results from Ng and Perron (2001) unit root tests

Variables (lags)	$\overline{MZ}_{\alpha}^{GLS}$	$\overline{MSB}_{\alpha}^{GLS}$	ADF^{GLS}	Variables	$\overline{MZ}_{\alpha}^{GLS}$	$\overline{MSB}_{\alpha}^{GLS}$	ADF^{GLS}
	With constant and trend				With constant		
LIM (2)	-18.38*	0.12*	-3.31*	LIM (2)	-8.88*	0.24*	-1.92
LP_{Spain} (2)	-3.95	0.328	-1.38	LP_{Spain} (2)	-1.29	0.57	-0.80
$LP_{Tunisia}$ (2)	-7.08	0.26	-1.69	$LP_{Tunisia}$ (2)	-4.83	0.32	-1.66
LP_{Greece} (0)	-3.76	0.34	-1.35	LP_{Greece} (0)	-1.01	0.60	-0.67
LP_{RW} (4)	-6.71	0.26	-2.45	LP_{RW} (4)	-3.28	0.37	-1.82
ω_{Spain} (2)	-18.32*	0.11*	-3.31*	ω_{Spain} (2)	-7.00	0.26	-1.85
ω_{Greece} (1)	-26.44*	0.14*	-3.92*	ω_{Greece} (1)	-17.24*	0.16*	-2.94*
$\omega_{Tunisia}$ (1)	-18.74*	0.16*	-3.25*	$\omega_{Tunisia}$ (1)	-4.89	0.31	-1.36
ω_{RW} (0)	-25.33*	0.14*	-4.12*	ω_{RW} (0)	-9.99*	0.21*	-2.28*
CV (5%)	-17.3	0.168	-2.91	CV (5%)	-8.10	0.233	-1.98

Notes: An L indicates that the variable is in logs. IM represents Total Italian Imports of virgin olive oil; P indicates the corresponding price and ω the corresponding market share. Finally, RW indicates imports from the Rest of the World. See Ng and Perron (2001) for a description of the different test statistics.

Before specifying an import demand system, it is important to determine if (as Italy is both a producer and an importing of virgin olive oil) we can estimate a demand system of only import demand equations; in other words, if Italian domestic production is separable from imports. To tackle with this issue we have used the Generalised Composite Commodity Theorem (GCCT) proposed by Asche et al. (2005) (see Section 2). As shown previously, all price series are nonstationary in levels, then, the GCCT requires imported and domestic prices to be cointegrated (i.e. price differences to be stationary). Thus Ng and Perron (2001) unit root test have been applied to differences between domestic and main virgin olive oil exporters' prices. Results are shown in Table 2. As can be observed, for the four prices, it is not possible to reject the null hypothesis of one unit root. Hence, we must conclude that the GCCT does not hold for this group of goods; thus justifying the estimation of an import demand system without taking into account the demand for domestic the virgin olive oil (i.e. the second step of the Armington's approach).

Table 2. Results from Ng and Perron (2001) unit root tests to price differences

Variables	$\overline{MZ}_{\alpha}^{GLS}$	$\overline{MSB}_{\alpha}^{GLS}$	ADF^{GLS}	Conclusion
	With trend			
LP_{Italy} vs. LP_{Spain}	-2.05 (1)	0.213	-2.02 (1)	I(1)
LP_{Italy} vs. LP_{Greece}	-1.92 (1)	0.341	-1.95 (1)	I(1)
LP_{Italy} vs. $LP_{Tunisia}$	-2.09 (3)	0.511	-2.03 (3)	I(1)
LP_{Italy} vs. LP_{RW}	-4.08 (2)	0.485	-2.23 (2)	I(1)
CV (5%)	-17.3	0.168	-2.91	
	With constant			
LP_{Italy} vs. LP_{Spain}	-4.45	0.53	-1.21	I(1)
LP_{Italy} vs. LP_{Greece}	-5.34	0.42	-1.42	I(1)
LP_{Italy} vs. $LP_{Tunisia}$	-3.28	0.64	-0.37	I(1)
LP_{Italy} vs. LP_{RW}	-6.31	0.38	-1.61	I(1)
CV (5%)	-8.10	0.23	-1.98	

As mentioned above, all expenditure shares and total imports are stationary around a time trend in their levels, but all prices are I(1). It is well known that an equation must be balanced, in the sense that all variables or linear combinations of variables must be integrated of the same order, to represent long-run relationships (Hendry, 1995). Moreover, the estimation of the static model (1) with a combination of stationary and non-stationary variables could lead to serious inference problems. Under such circumstances, two alternatives could be considered. The first one consists of calculating the first differences for the prices and then estimating a system similar to (1). We think that in this case the economic interpretation of the parameters is not straightforward.

The second alternative, as prices are nonstationary, is to test for cointegration among prices. If non-stationary prices are cointegrated, there exists at least a linear combination among them that is stationary, thus making each AIDS equation balanced and allowing us to estimate the system. However, the procedure is not straightforward. In fact, with n non-stationary prices, one can find at most (n-1) cointegrating vectors. In this case, all prices would follow the same stochastic trend. Finding (n-1) cointegration vectors implies that each pair of prices is cointegrated with a β_i parameter ($[1, -\beta_i]$) In addition, if the cointegrating parameter (β_i) can be restricted to the unity, then the relative prices are stationary, indicating that nominal prices are proportional and the Law of One Price Holds. This special case will only occur when the markets for the good in question are fully integrated. If this is the case, the AIDS can be specified in terms of relative prices (i.e. imposing homogeneity) and then, the system will contain only stationary variables.

Taking this in to consideration, our next step has been to check if the four prices in our demand system are cointegrated. The cointegration rank is determined using the likelihood ratio test introduced by Johansen (1988). A two-lag VAR system with restricted constant and trend was specified as the underlying model for carrying out such tests. The optimum lag was selected on the basis of the Akaike Information Criterion (AIC) and the Likelihood Ratio test proposed by Tiao and Box (1981). Misspecification tests for autocorrelation (Doornik and Hendry, 1997) indicated that the specified model was quite satisfactory. Table 3 shows the results from cointegration tests. As can be observed, at the 5% level of significance, the null hypothesis of three cointegrating vectors can not be rejected.

Table 3. Results from the cointegration trace statistic for the cointegration analysis between prices in the Italian demand system

r	LR	p-value	CV 95%
0	66.79	0.002	53.94
1	39.57	0.014	35.07
2	21.67	0.041	20.16
3	6.02	0.195	7.60

Given that the cointegrating rank is (n-1), we tested whether the Law of One Price holds. As mentioned this Law implies that that each cointegrating vector should satisfy the long-run condition (1,-1). Restriction tests on the cointegrating vector are asymptotically $\chi^2(v)$ distributed, where v is the number of imposed restrictions[3]. In our

[3] For further details, see Johansen (1995).

case, the Likelihood Ratio (LR) statistic was 5.21, which is well under the critical value at the 5% level of significance ($\chi^2(3)$=7.81). Therefore, it can be concluded that price homogeneity holds in the long run. In other words, model (1) can be expressed using relative prices.

5. The Threshold Almost Ideal Demand System (TAIDS)

Theoretical models assume linear adjustments of import quantities to changes in main economic determinants (prices and total imports). However, as Italy is also an important producer country of olive oil, it is not unrealistic to assume that import behaviour highly depends on the competitiveness of Italian production (excess supply or demand conditions). In this context, we could think that Italian imports of virgin olive oil imports depended on the existing relationship between domestic and import prices. To tackle with this issue, in this study, we have finally specified a dynamic two-regime Threshold Almost Ideal Demand System ($TAIDS_2$), which adopts the following expression:

$$\begin{aligned} w_{it} &= \alpha_i^1 + \delta_i^1 t + \lambda_i^1 w_{it-1} + \beta_i^1 \ln\left(\frac{y}{P}\right)_t + \sum_j \gamma_{ij}^1 \ln P_{jt} + \varepsilon_{it}^1 && \text{if } RIP_t \le \lambda \\ w_{it} &= \alpha_i^2 + \delta_i^2 t + \lambda_i^2 w_{it-1} + \beta_i^2 \ln\left(\frac{y}{P}\right)_t + \sum_j \gamma_{ij}^2 \ln P_{jt} + \varepsilon_{it}^2 && \text{if } RIP_t > \lambda \end{aligned} \quad (4)$$

where: $RIP_t = \ln(P_{Italy}/P_{importer})$ is the threshold variable; P_{Italy} represents the domestic price for virgin olive oil; $P_{importer}$ is a weighted average importer price; λ is the threshold parameter that delineates the different regimes; and t is a trend variable which accounts for changing tastes in Italian consumers.

As can be observed, the $TAIDS_2$ in (4) specifies that the imported demand system is regime specific. This model says that virgin olive oil import demand elasticities depend on the magnitude of the Italian price relative to the weighted average importer price. The two-regime TAIDS given in (4) can be compactly expressed as the following multivariate regression model:

$$w_{it} = X_t' A_i^1 I_t^1(\lambda) + X_t' A_i^2 I_t^2(\lambda) + \varepsilon_t \quad (5)$$

where $I_t^r(\lambda) = I(\omega(\beta) < \lambda)$ is a heavyside indicator function such that I(A)=1 if A is true and 0, otherwise; and

$$X'_t = \begin{pmatrix} 1 & t & w_{t-1} & \ln(\frac{y}{P})_t & \ln P_{1t} & \ldots & \ln P_{nt} \end{pmatrix} \quad A_i^r = \begin{pmatrix} \alpha_i^r \\ \delta_i^r \\ \lambda_i^r \\ \beta_i^r \\ \gamma_{i1}^r \\ \vdots \\ \gamma_{in}^r \end{pmatrix}$$

Note that when the threshold parameter (λ) is fixed (known a priori), the model is linear in the remaining parameters. In such circumstances, and under the assumption that errors ε_t are iid gaussian, parameters in model (5) can be estimated using Zellner's (1962) seemingly unrelated regressions (SUR) method.

However, in general, the value of the threshold parameter (λ) is unknown and needs to be estimated along with the remaining parameters of the model. Hansen and Seo (2001) provide a search procedure to estimate the values of λ:

$$(\hat{\lambda}) = \underset{\lambda \in [\lambda_L, \lambda_U]}{\arg\min} \left(\log \left| \hat{\Sigma}(\lambda) \right| \right)$$

where $\hat{\Sigma}(\lambda)$ is the estimated covariance matrix of model (5) conditional on (λ).

Once the parameters of model (4) have been estimated, the next step is to test if the AIDS model is linear or exhibits threshold non-linearity. This hypothesis can be formulated as:

H_0: $A_1 = A_2$ (symmetric adjustment) against the alternative

H_a: $A_1 \neq A_2$ (asymmetric adjustment)

The statistic to test such a hypothesis suffers from the problem of the so-called unidentified nuisance parameters under the null hypothesis. Given that the test statistic has a non-standard distribution, the critical values have to be determined by simulation methods such as the bootstrapping technique (for more details, see Hansen, 1997). As a solution to the above-mentioned problem, Hansen and Seo (2001) propose the following Sup-LM statistic based on the Lagrange Multiplier (LM) Principle:

$$\text{SupLM} = \sup_{\lambda_L \leq \lambda \leq \lambda_U} \text{LM}(\lambda) \tag{6}$$

where LM(λ) is the heteroskedasticity-robust Lagrange Multiplier (LM) statistic, which tests the restriction as given by the null hypothesis.

Taking equation (4) as the underlying model, Table 4 shows the results from the linearity tests. As can be observed, linearity is rejected at the 5% level in favour of the threshold model. The estimated threshold value is $\hat{\lambda}_1$= 0.059. Thus, the Italian imported demand system for virgin olive oil can be characterised by the following two-regime threshold process.

$$w_{it} = \sum_{l=1}^{11} \varphi D_{lt} + \begin{cases} \alpha_i^1 + \rho_i^1 w_{it-1} + \beta_i^1 \ln(\frac{y}{P})_t + \sum_j \gamma_{ij}^1 \ln P_{jt} + u_{it}^1 & \text{if } RIP_t \leq 0.059 \\ \alpha_i^2 + \rho_i^2 w_{it-1} + \beta_i^2 \ln(\frac{y}{P})_t + \sum_j \gamma_{ij}^2 \ln P_{jt} + u_{it}^2 & \text{if } RIP_t > 0.059 \end{cases} \quad (7)$$

Table 4. Results from non-linearity test for the Italian import demand system for virgin olive oil

	LR test for linearity
SupLm statistic	146.21
p-value	0.002
Threshold parameter	0.059

Once the TAIDS model has been identified and estimated, the first result from the model estimation concerns the test for theoretical restrictions (homogeneity and symmetry)[4]. Since homogeneity has already been tested and imposed, we will concentrate on the symmetry restrictions. The symmetry restriction has been tested using the Likelihood Ratio (LR) test. Results indicate that the symmetry restriction cannot be rejected by the data since the LR statistic is 9.29, which is well under the critical value at the 5% level of significance (χ_6^2 =12.59). Consequently, considering the whole set of results mentioned above, we can conclude that the estimated system to analyse the Italian import demand of virgin olive oil satisfies all theoretical restrictions and, therefore, the calculated elasticities will be consistent with theory.

Using the estimated parameters[5], Tables 5 and 6 show the calculated import demand elasticities[6] for both regimes. All estimated expenditure elasticities are statistically significant al the 5% level, except for the Rest of the World. Additionally,

[4] The eigenvalues of matrix C (see expression (3)) were all negative.

[5] The estimated parameters are not included but they are available from the authors upon request.

[6] Given the elasticities' definition in Chapter 3, their variances, based on which the t-values are calculated, are computed as follows:

$$\sigma^2_{\eta_i} = \frac{\sigma^2_{\beta_i^j}}{(\overline{w}_i^j)^2} \; ; \; \sigma^2_{\varepsilon_{ii}} = \frac{\sigma^2_{\gamma_{ii}^j}}{(\overline{w}_i^j)^2} + \sigma^2_{\beta_i^j} - \frac{2}{(\overline{w}_i^j)^2} \operatorname{cov}(\gamma_{ii}^j, \beta_i^j) \; ; \; \sigma^2_{e_{ij}} = \frac{\sigma^2_{\gamma_{ij}^j}}{(\overline{w}_i^j)^2}$$

statistical tests indicate that in the case of Spain, the expenditure elasticity is not significantly different from the unity. Results indicate that only in the case of imports coming from Tunisia the expenditure elasticity is higher than one. Consequently, Tunisia is likely to benefit/lose more from increases (decreases) of total Italian imports of virgin olive oil. In the case of Spain, exports to Italy seem to be very stable, varying according to total Italian imports.

Table 5. Uncompensated Italian import demand price and expenditure elasticities for the each regime.

	Regime 1 Italian relative price below the threshold level		Regime 2 Italian relative price above the threshold level	
	Expenditure	Price	Expenditure	Price
Spain	1.028*	-2.981*	0.944*	-1.887*
Greece	0.702*	-0.278*	0.900*	-2.261*
Tunisia	1.496*	-0.132	1.541*	-1.017*
R.O.W	1.311	-2.32	0.405	-1.826

* Denotes significance at 5% level.

If we compare the situation under the two regimes, differences are not very significant in terms of expenditure elasticities. As mentioned in the section above, the first regime is associated with lower domestic prices relative to imported prices while the opposite takes place in the second regime. Higher domestic prices benefit marginally to Greece and Tunisia. This situation implies that Spanish exports seem to be more price-oriented while exports from Grrece seem to be more quality-oriented. Tunisia is a different case, as imports from this country are subject to contingents and are mainly bulk to be mixed with domestic olive oils.

A similar conclusion can be obtained when analysing the own-price elasticities. In both regimes, own price elasticities for the, traditionally, two main exporters are significant. However, significant differences are found when comparing both regimes. When Italian prices are more competitive (Regime 1) only the Spanish price elasticity is higher than unity, indicating than under such a situation, Spanish exporters could gain market share in the Italian market through competitive prices. During the second regime, exports from the three main suppliers become price elastic, especially in the case of Tunisia and Greece. In other words, when Italian prices are less competitive (i.e. higher that import prices), Greece and Tunisia have more market opportunities and are more price-sensitive. This result is quite consistent with Figure 4, in which we showed that Greek prices used to be the highest among the main virgin olive oil exporters. Again, Spain seems to be very competitive in an environment of Italian decreasing prices while competition becomes harder when Italian prices are increasing.

Compensated cross-price elasticities are shown in Table 6. Substitutability and complementary among the different countries are indicated by positive and negative cross-price elasticities, respectively. In both regimes cross-price elasticities, in all cases except one, have the same sign. In general terms, between Spain and Greece, there exists a high degree of substitution, which has to do with production conditions in both countries (see Figure 3). Greece increases exports to Italy when Spanish production is relatively low.

Table 11. Hicksian cross-price elasticities from the Italian import demand system.

	Regime 1 Italian relative price below the threshold level			
	Spain	Greece	Tunisia	R.O.W
Spain	-	1.328*	0.475	0.821*
Greece	1.121*	-	-0.599*	-0.530*
Tunisia	0.868	-1.297*	-	0.279
R.O.W	4.893	-3.742	0.912	-
	Regime 2 Italian relative price above the threshold level			
	Spain	Greece	Tunisia	R.O.W
Spain	-	1.055*	0.297	0.012
Greece	2.149*	-	-0.255*	0.122
Tunisia	1.158	-0.488*	-	0.128
R.O.W	0.211	0.567	1.034	-

* Denotes significance at 5% level.

Spanish exports show also a certain degree of substitutability with imports coming from the Rest of the World, mainly when Italian prices are competitive in relation to the import price. On the other hand, the cross price elasticity between Greece and Tunisia is negative, indicating complementary among them. This result is related to the fact that imports from Tunisia are subject to quotas, reducing price competitiveness. Once the quota is surpassed, imports from Greece play an important role. As indicated in Figure 3, Imports from Tunisia have been stabilised, with those from Greece have been decreasing during the last decade Finally, we can conclude that, in general terms and as expected, the magnitude of the elasticities are lower during the second regime, that is, when Italian prices are less competitive.

6. Concluding remarks

This paper has analysed the import demand for virgin olive oil in the EU. However, given that Italy concentrates more than 80% of EU imports, we have concentrated in Italian imports, which, on the other hand, is the most interesting EU country due to its two-fold condition of virgin olive oil exporter and importer. The ultimate objective has been to determine the relative position of Mediterranean EU and non-EU countries exports and their degree of substitutability or complementarity. The methodology used is based on the specification of a Threshold Almost Ideal Demand System in which special attention has been paid to the stochastic properties of the series involved, that is, if they are or not stationary and, in the latter case, if they are cointegrated. In an empirical context, the paper has aimed to provide a set of import demand elasticities that can be useful in trade models.

Results from unit root tests indicate that prices are non-stationary, while the market shares and total imports are stationary. This is an interesting case as most of the literature to date has considered all variables either stationary or non stationary. A two-step modelling approach has been followed here. In the first step, relationships among international prices are considered to check if price homogeneity holds. Since the null of price homogeneity has not been rejected, we next have estimated the imports demand system using relative prices. In the estimated model other theoretical restrictions (symmetry and negativity) were empirically tested and imposed. Moreover, separability with the domestic production has been tested. As domestic production could be considered separable from imports, we finally have estimated a strictly imports demand system.

However, in spite of domestic production being separable from imports, it is plausible that imports could be affected by domestic prices, thus generating non linear demand systems. This fact has been explicitly considered in this paper, which is one of the main methodological contributions of this study to the existing literature. In fact, we have differentiated imports behaviour depending whether the domestic price relative to an average import price is above or below a threshold.

To analyse the relative situation of the different exporting countries into the three importing countries, we have computed the expenditure and price elasticities.

These measure the response of imports from a specific country to changes in its own price, to changes in prices from other origins and to changes in the total imported volume. Here, different conclusions could be obtained, which are quite consistent with the evolution of market shares. Results point to Spain as the leader in the Italian virgin olive oil market. What's more, it is expected that this position will be maintained in the future. Greece has improved its relative position after its accession into the EU. However, imports coming from Greece are highly dependent on the situation in Spain. Traditionally, Greek olive oil has been the main substitute for Spanish oil when there are shortages in the Spanish production. In the case of Tunisia, future prospect in a new context of trade liberalisation are positive as expenditure elasticity is higher than unity, in spite that its exports are currently constrained due to existing quotas. Moreover, results related to the Italian market, seems to indicate that while Spanish exports are more price-oriented, Greek exports are more quality oriented.

Our results indicate that Tunisia has good potential for future exports development as a consequence of new perspectives of trade liberalisation taking into account its relative position in the Italian market. However, little can be said for other non EU Mediterranean countries (Morocco and Turkey) as still they represent an insignificant market share, waiting for market opportunities related to bad production conditions in traditional olive oil producers. However, the crucial question for further research is to what extent these new perspectives for the Tunisian exports are going to increase the welfare of Tunisian producers.

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