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MACROECONOMIC IMPORT FUNCTIONS WITH IMPERFECT COMPETITION. AN APPLICATION TO THE EC TRADE

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ABSTRACT

Macroeconomic Import Functions with Imperfect Competition. An Application to the EC Trade*

This paper analyses the consequences for the standard import allocation models of assuming monopolistic competition on the supply side. Together with relative prices, this requires additional variables to capture product differentiation effects. To this end, we derive a composite price index from a nested CES-translog demand system. Our empirical work is twofold: first, we try to assess the long-term relationship between market shares and relative prices by using a cointegration technique, and second, we estimate the demand system for domestic, European and foreign products in the main European markets. The results show that a composite price index, integrating product differentiation, tends to perform better than pure price effects alone in a significant number of cases. We use the estimation results to assess the impact of a potential homogenization of tastes over European markets, after the '1992' integration process.

JEL classification: C30, F12, F14, F15, F17
Keywords: import equations, product differentiation, market integration, translog-CES demand system, EC trade

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NON-TECHNICAL SUMMARY

This paper has both empirical and methodological motivations. The first aim is related to the assessment of the consequences for international trade of the '1992' market integration process. To this end we estimate the relevant demand parameters characterizing the competition between European and non-European producers in European markets. We also try to improve the usual methodology on import demand systems in one direction suggested by trade theory by introducing imperfect competition. According to the monopolistic model of international trade, the supply of industries with a great number of firms should be modelled as a bundle of differentiated products rather than as a homogeneous good. As a consequence of assuming product differentiation at the level of the firm, the number of products or equivalently the number of firms at a macro level appears among the determinants of market shares.

We develop a flexible demand system along these lines, in which prices and the number of firms jointly determine market shares. This has the potential advantage of capturing secular shifts in market shares (e.g. an increase in import penetration). In standard import demand equations they remain unexplained and are treated via the introduction of deterministic trends. In our system, however, the number of firms of each country producer can be related to market structure or economic growth. We conduct empirical tests for the four major European countries, France, Germany, Italy and the UK, and for three industries with quite different market structures: textiles and leather, chemical products and electrical machinery. In each market three producers were identified: domestic, other European and non-European. In this way market shares of both foreign and domestic sources of demand are simultaneously determined.

The empirical implementation of this model requires information on the number of representative firms in a given industry. Given the lack of microeconomic data we proxy the number of firms using an activity index. Typically, for a fragmented market structure this index can be a good proxy for capturing the rate of creation of new firms, because the number of firms grows in parallel with output expansion of the industry. We investigate the empirical relevance of our proxy by using available data on the number of domestic firms in France, Germany and the UK for three sectors. The results indicate that in the textiles and leather sector and in the electrical goods sector there is a positive and significant relation between the activity index and the number of firms. For the chemicals sector this relation is not supported by the data, suggesting a segmented market structure

We first try to establish a long-term relationship between market shares and relative prices in a simplified model excluding non-price effects. The results show that in a significant number of cases this relation does not exist. A model based on pure price effects cannot, therefore, account for the behaviour of market shares in the long run. Estimating the demand system including non-price effects, we find that in seven out of twelve cases, this extended model performed better in explaining market shares than relative prices alone. As is common in the literature on trade equations, the estimated price elasticities are low. One interesting empirical finding is that estimated non-price effects display a much higher elasticity than price effects. Under the assumption that our proxy captures product differentiation in an appropriate way, this can be interpreted as a high valuation of product variety in European markets. This has important implications for the debate on the consequences of '1992' market integration.

Finally, we use the estimated parameters to simulate the effects of a counterfactual 10% decrease of domestic producers' price. We show that this price shock has differentiated effects across markets. On average, however, the induced movements of market shares are small. This last result is reinforced in a scenario where the tastes are assumed to be homogenous across domestic markets.

In conclusion, the findings of this paper support the conventional wisdom of the empirical trade literature that price movements induced by market integration would lead to a moderate impact on market shares. None the less, if market integration also leads to large changes in the market structure and in particular in the number of competing firms, our estimates predict that this could have a much greater impact on demand.

1. INTRODUCTION

Macro-economic trade equations - defined at a global or at a sectoral level - usually consider industries as a homogeneous aggregate. Even if products are assumed to be differentiated by place of production (the usual Armington hypothesis), within each particular grouping of goods there is, implicitly, homogeneity among individual components. This may be a serious drawback as estimates of trade equations with usual price variables cannot capture a key element of imperfect competition suggested by trade theory, namely, that industries should be modelled as a group of heterogeneous firms. This paper explores one avenue of introducing into a system of macro-trade equations the impact of the heterogeneity existing at the firm-level.

The first part of the paper analyses the consequences of assuming the supply-side hypothesis of oligopolistic competition with symmetric firms in the usual trade equations framework. Norman (1990) develops this point in a CGE framework, but here, we focus on demand-side aspects. The oligopolistic model suggests that, in a given market, prices and the number of products (or firms) competing are joint determinants of market shares. Traditional equations can be generalised to incorporate both effects in a tractable way. Bismut and Oliveira Martins (1987), and Oliveira Martins (1989) developed this approach with CES trade functions, and this paper extends it to the more flexible Translog system which seems more appropriate for modelling market shares over a long period (1963-1987).

The empirical test of this model is conducted for the four major European countries, France, Germany, the U.K. and Italy, and for three industries with quite different market structures: Textiles and leather, Chemical products and Electrical machinery. Following Winters (1984), we do not impose the stringent separability hypothesis between domestic and foreign sources of domestic demand. Accordingly, in each market, three types of producers are identified: national, other EC and all other non-european producers. The empirical analysis begins with an assessment of the statistical long-term relationship - measured by a co-integration technique - between market shares and relative prices. It turns out that in a significant number of cases there is no evidence of the existence of such a long-term relationship, suggesting that other supply-side variables should be taken into account. The system of market shares is then estimated using as the explanatory variable a "composite price index" embodying prices and an activity variable as a proxy for the number of firms.

The estimates of the demand system enable us to calculate the key substitution parameters characterising the degree of product differentiation among aggregate producers. Finally, we use this set of parameters to simulate the consequences of the 1992'integration process. The integration of European markets is sometimes predicted to lead to the homogenisation of tastes across countries. This implies a hypothetical change in the underlying parameters of the utility function that can be compared with the actual estimates.

2. MACRO-TRADE FUNCTIONS and IMPERFECT COMPETITION

In the spirit of monopolistic competition models, each aggregate regional producer offers a bundle of differentiated products (or varieties) supplied by a given number of heterogeneous but symmetric firms[2]. Each firm produces only one product which makes the number of products equal to the number of firms. In order to focus on the demand equation, we treat as predetermined the price and the number of "representative" firms within an industry.

As a starting point, we assume that a given national market is supplied by two aggregate producers (for example, domestic and foreign producers), referred to as N and E . For convenience, we define the ratio between the market share of each producer, V_{NE} :

(1) $$V_{NE} = \frac{\sum_{i=1}^{n_N} p_{N_i} \cdot q_{N_i}}{\sum_{i=1}^{n_E} p_{E_i} \cdot q_{E_i}}$$

where n_N and n_E are, respectively, the number of firms of the aggregate producers N and E. With symmetry, the ratio V will be equal to $(n_N \cdot p_N \cdot q_N / n_E \cdot p_E \cdot q_E)$, where p_N and q_N are the price and output of the representative firm in N, and similarly for E.

Models of international trade with monopolistic competition often assume Dixit-Stiglitz CES preferences (see the survey by Helpman, 1990). In that

[2]Following the mainstream of this literature, only a symmetric market structure will be considered in this paper. A recent paper by ABD-EL-RAHMAN (1991) incorporates empirical information on intra-firm heterogeneity together with trade data in order to explain the overall composition of trade.

case, with rational behaviour, the ratio between the market shares of the two aggregate producers is given by:

$$(2) \qquad V_{NE} = \frac{n_N \cdot p_N \cdot q_N}{n_E \cdot p_E \cdot q_E} = c \cdot \left[\frac{n_N}{n_E}\right] \cdot \left[\frac{p_N}{p_E}\right]^{(1-\sigma)}$$

where $\sigma > 1$ is the elasticity of substitution between each pair of products and c is a constant depending on the parameters of the CES utility function. This equation says that the ratio V is determined by relative prices between N and E and by the relative number of firms. As Norman (1990) observes, this result is quite different from the traditional Armingtonian equation. Indeed, with perfect competition [3], the term n_N/n_E - which can be viewed as an effect related with imperfect competition on the supply-side - will not appear in the equation. If the "true" model is the imperfect competition one, an econometric estimate of V using only relative prices as the explanatory variable will suffer a specification bias. Indeed, available price data are generally based on weighted averages of individual prices, and therefore they cannot capture the valuation of product diversity as it is in the case of a CES aggregate[4].

It is very likely, then, that supply-side effects have been underestimated by empirical work on trade equations, as noted by Goldstein and Khan (1985). There have been attempts at introducing non-price competitiveness effects in trade equations (e.g. Barker (1977), Geracci and Prewo (1982)) but generally without an explicit link to trade theory, and indeed the practice of introducing time trends to trade equations can be interpreted as an attempt to proxy missing supply-side effects. All these heuristic approaches can be improved upon, taking account of our recent understanding of how imperfect competition interacts with trade flows. Along these lines, the imperfect competition model suggests that one should construct a "composite price index", encompassing the ratio (n_N/n_E) that combines both price and non-price effects.

The market share equation (2) can serve as basis for estimation if consumers perceive difference among varieties but do not perceive a global difference between the bundles offered by the aggregate producers N and E. However, as trade equations are usually estimated at a broad level of aggregation, it seems more convenient to allow for different degrees of differentiation within and outside each

[3] Actually, perfect competition in the supply-side was not an explicit assumption in the original paper by Armington (1969). He derived only a demand system with imperfect substitutes. Afterwards, the so-called Keynesian approach of trade flows made this hypothesis explicit by assuming an infinite elastic supply.

[4] This is a well-known problem in price index theory. See, for example Lloyd(1975) for a discussion of the bias arising from a Laspeyres approximation of a two-level CES price index.

group. To this end, Bismut and Oliveira Martins (1987) used a two-level CES function with an intra-firm layer and an aggregate producer layer, with three elasticity parameters characterizing the substitution possibilities. Their market-share ratio V is then given by:

$$(3) \quad V_{NE} = c \cdot \frac{n_N^{\beta_N}}{n_E^{\beta_E}} \cdot \left[\frac{p_N}{p_E}\right]^{(1-\sigma)}$$

where $\beta_N = (1-\sigma) / (1-\sigma_N)$ and $\beta_E = (1-\sigma) / (1-\sigma_E)$

σ_N and σ_E, (both >1) are the elasticities of substitution inside each group. The parameter σ which characterises the substitutability between the two bundles of products is required only to be positive. Compared with equation (2) the two-level system has the same price elasticities but the "product elasticities" β_N and β_E can now be different from one. Whereas, in the previous model increasing product variety by one producer always increased market shares, in equation (3) its effect depends on the elasticity of substitution between the two bundles of differentiated products. For a very low degree of substitution ($\sigma < 1$) the entry of new firms can have an adverse effect on the ratio V. Also, one should expect (but not necessarily) that the upper-level elasticity σ is lower than the intra-varieties elasticities of substitution[5]. As the equation (3) is embodied in equation (2), it would be possible to choose between the two models by testing the hypothesis $\beta_N=\beta_E=1$.

In order to carry out an econometric estimation with more than two bundles of products, the generalisation of equation (3) would entail severe restrictions on the parameters. Two options are suggested by the literature (see, Deaton and Muellbauer (1980), chapter 2):

i) Assume that the overall substitution between each pair of differentiated product bundles is the same. This is equivalent to assuming strong separability among all aggregate sources of domestic demand.
ii) Or, assume weak separability among groups of products. This hypothesis will lead to a nested framework in which groups of products are in their turn gathered together in broader groups over several layers. A structure of this type would be very similar to the nesting used in applied general equilibrium models.

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[5]As shown by Sato (1967), inside a group, say N, the Allen partial elasticity of substitution equals:

$A_{ij} = \sigma + \frac{1}{w_N} \cdot (\sigma_N - \sigma)$ for $i,j \in N$, $i \neq j$.

where w_N is the market share of aggregate producer N. When σ is higher than σ_N, the varieties are complementary and a monopolistic competition equilibrium will be unstable.

The first option seems very restrictive on cross-price (and "product") effects. On the other hand, following the second option needs a careful choice of the separability hypothesis embodied in the nesting. Indeed, the price effects can be radically different according to the separability assumptions. Empirical work has consistently reported that, as far as price effects are concerned, separability between foreign and domestic source is rejected by the data (see Winters, 1984, 1985).

Moreover, over a long period it also seems quite restrictive to assume that the overall substitution effects remain constant. Accordingly, it seems best to abandon the search for constant elasticities of substitution and adopt instead a more flexible demand system. Among the many candidates, we chose the Translog functional form. The main advantage of the Translog system is that it allows for variable own- and cross-effects within a tractable form. The AIDS model (see, Brenton and Winters, 1991), could also be a possible alternative, but it would be somewhat more complex to handle, namely when dealing with the composite prices defined below.

Consider the Translog indirect utility function (introduced by Christensen, Jorgenson and Lau, 1975):

(4) $$- \mathrm{Log}\, U = \alpha_0 + \sum_i \alpha_i \,\mathrm{Log}(\Pi_i/Y) + \frac{1}{2}\sum_i \sum_j \beta_{ij}\,\mathrm{Log}(\Pi_i/Y)\,\mathrm{Log}(\Pi_j/Y)$$

where $[\Pi_i]$ is the vector of prices and Y total income. In our experiment, we assume rational behaviour implying constraints on the parameters:

$$\sum_j \beta_{ij} = 0\,, \quad \beta_{ij}=\beta_{ji} \quad \text{and} \sum_i \alpha_i = 1\,.$$

These constraints implying homogeneity of degree one with respect to income. By using the logarithmic form of Roy's identity, one gets the the market-share w_i of producer *i* :

(5) $$w_i = \alpha_i + \sum_j \beta_{ij}\,\mathrm{Log}(\Pi_j)$$

According to the argument above, the price vector $[\Pi_j]$ should be defined over a bundle of differentiated products, incorporating both a pure price effect and a "variety" effect. As in the Dixit-Stiglitz approach, we assume a CES functional form for the composite price index[6]:

$$\Pi_j = \left[\sum_{k=1}^{n_j} P_k^{(1-\sigma_j)} \right]^{\frac{1}{(1-\sigma_j)}} \tag{6}$$

where σ_i is the intra-group elasticity of substitution. Moreover, with intra-group symmetry (hence each product has the same price P_j) the composite price index for source j will be:

$$\Pi_j = n_j^{1/(1-\sigma_j)} . P_j \tag{7}$$

It can be noted that, if the first-level Translog system is homothetic, by introducing the CES second-level, we possibly are also introducing other income-type effects via the number of firms (or the number of products). This is an important point related with the effects of product differentiation embodied in this system [7].

We recall the usual direct (ε_{ii}) and cross-price (ε_{ji}) elasticities between groups of products derived from the translog parameters:

$$\varepsilon_{ii} = (\beta_{ii} + w_i^2 - w_i) / w_i \tag{8}$$

$$\varepsilon_{ji} = (\beta_{ji} + w_j . w_i) / w_j \tag{9}$$

This demand system can be used to characterise the degree of substitution among different bundles of products. The direct (Hicksian) or the Allen elasticities of substitution are not very appealing in the n-commodity case, but there is an alternative measure that can be interpreted in terms of the curvature of indifference surfaces - thus measuring the ease of substitution. This is the so-called Morishima

[6]Note that this Translog-CES system is a nested structure which is different from the more general CES-Translog system proposed by Pollack, Sickles and Wales (1984); the latter combines both a CES and a Translog function at the same level.

[7] See Krugman (1989) for a discussion on the relation between income elasticities and growth in the context of a monopolistic competition model of international trade.

partial elasticity of substitution M_{ij}. It can be defined as (see, Blackorby and Russel, 1981,1989):

(10a) $$M_{ij} = \varepsilon_{ji} - \varepsilon_{ii}$$

By using (8) and (9) one gets:

(10b) $$M_{ij} = \beta_{ji}/w_j - \beta_{ii}/w_i + 1$$

M_{ij} is asymmetric since it refers only to the situation in which the composite price of group i varies. Hence, the possibilities of substitutability between groups i and j will be different if only the price of group j varies[8]. This measure, however, has the appealing property of being a straightforward generalisation of the 2-group case by relating clearly the impact of relative prices over the market share ratio V_{ij}. Indeed, by taking the logarithmic derivatives of market shares w_i and w_j with respect to the relative composite price and comparing with (10b), one gets:

(11) $$\frac{\partial \mathrm{Log}(w_i/w_j)}{\partial \mathrm{Log}(\Pi_i/\Pi_j)} = 1 - M_{ij}$$

We have now set up the basic framework of a macro-economic import system with imperfect competition. We now turn to the data sources and the empirical estimates.

3. EMPIRICAL IMPLEMENTATION

3.1 THE DATA

For France, Germany, the U.K. and Italy, data were collected for trade and output variables for domestic, other European and non-European producers, on an annual basis covering the period 1963-88 (except for production data in Italy which cover 1967-1988). In order to cover various market structures, three quite different

[8] In this case all relative prices Π_j/Π_k, $k \neq i$ would vary, whereas they remain constant in the previous case.

industrial sectors were chosen: Textile and leather, Chemical products and Electrical goods.

Trade data (values and unit values) were derived from the EC Volimex data base[9]. Unfortunately, the production data were not available from an unique source. The primary source was the EC sectoral B.D.S. data base[10]. As, there are many missing data concerning production, this data base has to be completed with other sources - in order of preference, the OECD IAI data base, the Statistical Office of the European Communities (SOEC) Industrial Statistics and the UNIDO's Industrial Statistics. Data for France before 1970 were derived from a particular source, the PROPAGE data base[11]. Annex 1 summarizes the data collecting process. Domestic demand is derived by using the usual identity:

$$(12) \qquad P^d D + P^x X = P^q Q + P^m M$$

where D is demand, X exports, Q domestic production, M imports, with their respective prices (based in 1980).

3.2 A PROXY FOR THE PRODUCT DIFFERENTIATION EFFECT

As data for the number of firms do not exist for a sufficiently long time period, the variables corresponding to the number of firms must be replaced by proxies. Nonetheless, some information is available for France, Germany and the U.K. is available in the SOEC-Industrial Statistics, that can be used to qualify the proxies. We constructed our proxies from the industrial activity indexes of the UNIDO Industrial Statistics data base. These indexes were constructed in two steps. First, an aggregate index for each of the major trading partners of the Volimex classification was calculated: individual EC countries, USA, Japan, Australia+New-Zealand+South Africa, rest of OECD and the dynamic Asian economies (Singapore, South Korea, Hong Kong, Taiwan and Malaysia). Second, the group was weighted together into our aggregates by their import market share in the base year 1980 for each market/sector we consider (4 markets x 3 products).

The rationale behind the proxy relationship is grounded on the monopolistic competition model. Therefore, the proxy will capture better the

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[9]This data base contains bilateral trade flows for each OECD country and a world breakdown of 30 groups of countries at the Nace-Clio disaggregation level (25 products), for the period 1963-88.
[10]This data base contains value added (value and volume), production (value and volume), investment and employment for the twelve EC countries at an aggregation of Nace-Clio level.
[11]This data base can be provided by the I.N.S.E.E., Paris, upon request.

product differentiation effects when the market structure is fragmented, i.e. when the number of firms grows in parallel with output expansion of the industry. We used the available data from the SOEC Industrial Statistics for the period 1975-87 to test the relation between the activity index (I_N) and the number of domestic firms (n_N) in each of the three sectors. The results of the pooled regressions[12] are shown in table 1:

Table 1. Results of the Pooled regressions

-Textiles and leather:

$$\text{Log } n_N = \underset{(7.84)}{1.283} \text{ Log } I_N \qquad R^2 = 0.65 \quad \text{See} = 0.088 \quad \text{ndf} = 36$$

-Chemical products:

$$\text{Log } n_N = \underset{(-2.87)}{-0.302} \text{ Log } I_N \qquad R^2 = 0.22 \quad \text{See} = 0.062 \quad \text{ndf} = 36$$

-Electrical goods:

$$\text{Log } n_N = \underset{(8.50)}{0.672} \text{ Log } I_N \qquad R^2 = 0.69 \quad \text{See} = 0.053 \quad \text{ndf} = 36$$

Note: Period 1975-87. All variables are defined as deviations from their country sample means. Student t-ratios are in parenthesis. See=standard error of the regression; ndf= number of degrees of freedom.

The results indicate, as far as the main European producers are concerned, that the proxy should perform much better for Textiles and leathers and Electrical goods industries than for the Chemical sector. Indeed, in the latter sector the negative correlation between the activity index and the number of firms suggests that the market is closer to a segmented rather from a fragmented structure. In this case, the proxy may not be consistent with the underlying hypothesis of oligopolistic competition with symmetric firms and free entry. The estimates presented in section 3.4 below confirm at some extent these presumptions.

[12]In order to take into account the cross-country differences, we estimated a fixed effects model (or the so-called "within" estimator) over the pooled data of France, Germany and the U.K. for the period 1975-87. A break in the statistical coverage of the number of firms for Italy, this country was removed from the sample.

3.3 MARKET SHARES AND RELATIVE PRICES: Is there a long-term relation ?

This section conducts a preliminary test. Before estimating our extended model, we decided to test for the existence of a long-term relation between market shares and relative prices. If such a relation do not hold, this would suggest that a more general model may be required by the data. The stationarity and co-integration tests described in this section can be viewed as an analysis of this question.

Intuitively, cointegration among a set of variables implies that there exist fundamental economic forces which make the variables move stochastically together over time. Johansen (1988) and Johansen and Juselius (1990) provided a unified approach based on a maximum likelihood procedure for estimation and testing in the context of a multivariate system[13].

We applied this procedure to test whether our sample was compatible with the following simple (often used) long-run empirical relation between market shares and relative prices:

$$\text{(13)} \qquad \mathrm{Log}(q_i / D) = a \,.\, \log(p_i / P) + b$$

where q_i is the demand for product i in a given market, p_i is its respective price, D is total demand, P is the price index of total demand. Variable q_i is equal to (Q-X) for domestic producers and equal to imports for foreign producers. Table 2 applies the Johansen's procedure to the two-dimensional vector composed of market shares and relative prices taken in logarithms. We assume an autoregressive process of

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13 Johansen's approach can be summarised as follows. Consider a p-dimensional gaussian autoregressive vector:

$X_t = \sum_{t=1}^{k+1} \pi_t X_{t-i} + e_t$, with non-singular covariance matrix. By reparameterising the process in first differences, we get: $\Delta X_t = \sum_{i=1}^{k} \Gamma_i \Delta X_{t-i} + \Gamma_{k+1} X_{t-k-1} + e_t$. The rank of Γ_{k+1} gives the dimension of the cointegration space. Under the null hypothesis that this dimension is equal to r, this matrix can be decomposed into $\Gamma_{k+1} = \alpha\beta'$ where α and β' are p.r full rank matrices. The dimension of the cointegration space can then be determined sequentially, by analysing the canonical correlations between levels and first differences corrected for lagged differences. The determination of this dimension is based on a likelihood ratio statistic with a known asymptotic distribution. Given this dimension, Johansen's procedure allows for testing an hypothesis on the structure of the co-integration space, e.g. that this space contains or is contained in another space. The test statistic is distributed as chi-squared. In particular, it can be used to test whether the series are stationary or not.

order 2 and that series can be integrated up to order 1. Moreover, we also assume that there are no deterministic trends in the series. When the dimension of the cointegration space is different from zero, we test whether the vector(s) which spans this space embodies one of the two variables. In this way the stationarity of the series can be tested.

Over the 36 instances (4 markets x 3 sectors x 3 producers), the dimension of the cointegration space was found to be zero in 22 cases (61%). In this group, market share and relative prices are integrated but not cointegrated, and the existence of a long-run relation between the two series is rejected by the data. For the remaining 14 cases, the dimension of the cointegration space is one. The tests on the structure of the cointegration space show that in 8 of these, the stationarity of only one of the two variables is accepted. In only 6 cases (14%) was a cointegration relation found. This happens in France for domestic producers in Chemical and Electrical products, and for the non-European producers for Electrical products; in Italy, for European producers in Textiles and Chemical products, and in the U.K., for European producers in Textiles.

These results require two caveats however. First, the small size of our sample makes the tests very weak. Second, there is a strict inconsistency in a bounded variable like a market share generating a non-stationary time series[14]. This probably can only reflect our finite sample size. In spite of these limitations, we conclude that equation (13) is rejected by the sample in the majority of cases we studied.

It would be difficult to continue this road and extend the dimensionality of the co-integration test without imposing non-linear constraints or relations among the variables, hence the need for a structural modelling approach of demand systems.

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[14]Note that as the estimates were made in Logs, the market share variable is only bounded upwards.

Table 2. Tests for the dimension and the structure of the cointegration space between the logarithms of market share and price competitiveness.

(Sample: 1963-88, except for Italy :1967-88)

Market	Producer Sector	Domestic	European	Non-European
France	Textiles	0	0	0
	Chemicals	1 (*)	1 (a)	1 (a)
	Electrical	1 (*)	0	1 (*)
Germany	Textiles	1 (a)	1 (b)	0
	Chemicals	0	1 (a)	0
	Electrical	0	1 (b)	0
Italy	Textiles	0	1 (*)	0
	Chemicals	1 (a)	1 (*)	0
	Electrical	0	0	0
U.K	Textiles	0	1(*)	1 (a)
	Chemicals	0	0	0
	Electrical	0	0	0

Note: The numbers indicate the dimension of the cointegration space.

(a) : The tests on the structure of the cointegration space indicate that relative prices are stationary while market share is not.

(b) : The tests on the structure of the cointegration space indicate that market share is stationary while relative prices are not.

(*) : The hypothesis of cointegration between market shares and relative prices is not rejected.

3.4 ESTIMATION AND RESULTS WITH THE NESTED TRANSLOG-CES DEMAND SYSTEM

The estimation of the Translog-CES market share equations defined above was carried out for each of the 12 markets (4 countries x 3 products). In each market three producers are identified: domestic, other European and non-European. The system was estimated simultaneously by a maximum likelihood technique[15]. Given our limited data set, the restrictions of homogeneity and symmetry were imposed *a priori*. The adding-up constraint is fulfilled by dropping the equation for the foreign producers, but as is well known the maximum likelihood estimates are independent from the choice of the equation which is dropped from the system[16]. Based on Anderson and Blundell (1982), the only dynamic form which seemed tractable in our framework was a very simple partial adjustment process with a common adjustment speed (λ) across all suppliers. Preliminary estimates, not reported here, showed that the dynamic model tends to perform much better than the static model with respect to residual autocorrelation. By combining equations (5) and (7) with a proxy for the number of firms and a partial adjustment process, one finally gets the system of equations to be estimated:

$$(14)\quad \Delta w_i = \lambda . \left(\alpha_i + \sum_j \beta_{ij} . \left(\mathrm{Log}(P_j) + 1/(1-\sigma_j).\mathrm{Log}(I_j) \right) - Lw_i \right) + u_i$$

where I_j is a proxy for the number of firms (in our case, an activity index), (i,j) stands for national (N), other European (E) and non-European (F) producers, L for the lag operator and u_i is a normally distributed random term. The first aim was to estimate the full system in a systematic way, but it turned out that because of convergence problems, it was necessary to adopt a specific estimation strategy[17] and to add more structure to the system. In general, the near collinearity between activity indexes made it difficult to estimate all the second-level CES elasticities of substitution freely. In that case it was necessary to add more structure to the system by imposing equality on some of the σ_j parameters across the three producers. For

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[15] We used the non-linear least squares procedure of TSP 4.1a.

[16] See ITALIANER (1986) for a exhaustive review of simultaneous systems of equations techniques applied for import allocation models.

[17] As very often when estimating non-linear systems, the initialisation point is crucial to get convergence of the maximisation procedure. To overcome this problem, we adopted a linear iterative procedure over two subsets of parameters; this procedure performed quite well to supply initial estimates. More information on these technical aspects can be supplied by the authors upon request.

Germany, the U.K. and Italy, we had to impose equality of the three elasticities. For France, it was possible to obtain a more general form by imposing only the equality of two σ_j [18]. In addition, in very few cases, the adjustment speed λ was also constrained in order to ease the convergence process[19].

For each market the nested Translog-CES model can be compared with the one-level system where varieties are implicitly supposed to be homogeneous ($\sigma=\infty$). Table 3 gives the results of LR test between the two models[20]. Tables 4-7 show the parameters estimates.

Table 3. Comparison between the homogeneous and the differentiated product model.

Sector/ Country	TEXTILES	CHEMICALS	ELECTRICALS
France	12.4**	3.4	9.6**
Germany	3.4	6.2*	0.4
U.K.	10.2**	10.4**	5.2*
Italy	1.6	0.2	13.2**

Note: Log-Likelihood ratio (LR) test. The test statistic follows a $\chi^2(2)$ for France and a $\chi^2(1)$ for the other countries.
(*) the homogeneous product model is rejected at the 5% level.
(**) the homogeneous product model is rejected at the 1% level.

Several inferences can be drawn from the results. In seven out of twelve cases, the model embodying product differentiation effects increases significantly the likelihood of the sample (see table 3). In the Electrical goods sector, the composite price effect also appears to be more significant than in the other sectors. On the other hand, the extended model works better in France and in the U.K than in the other two countries. Parameter σ_j is significatively different from zero and tends to be greater than one in the majority of the cases, a result compatible with the monopolistic competition hypothesis. However, one could expect to find greater values for these parameters[21]. It is possible that our high level of aggregation and weak proxy have downward biased our estimates of the intra-variety elasticity of

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[18]The choice of the constrained elasticity was based on the likelihood of the estimates.
[19]The value was chosen according to a grid search over the range]0,1].
[20]The null hypothesis corresponds to the homogeneous product model where $1/(1-\sigma)$ is constrained to be zero.
[21]Namely, when they are compared with the equivalent parameters calibrated in AGE imperfect competition models (e.g., Smith, Venables and Gasiorek, 1992).

substitution. On the other hand, it should be noted that the effect of expanding the variety of products over market shares reduces very quickly with increases in this elasticity[22]. Thus, the estimated intra-variety elasticities imply quite a high impact of the non-price effects. Under the assumption that our proxy reflects the number of products, this can also be interpreted as a high valuation of the variety embodied in the preferences. The value of this intra-variety elasticity of substitution can be crucial for the assessment of welfare effects of market integration (see Burniaux and Waelbroeck, 1992). In applied GE models with high values of σ_j the variation of the number of products has a small impact on welfare; typically, the benefits from market integration coming from firm specialisation will not be outweighed by the decrease in the number of products. For small values of this elasticity, the variety effect may dominate the sign of welfare gains.

In brief, in a significant number of cases, the broad picture seems to be consistent with the extended model incorporating non-price effects. But one must be bear in mind that the very simple symmetric market structure which underlies the proxy for product differentiation represents only one possible source of non-price effects.

[22]The elasticity of the market share of producer i with respect to the number of products of producer j is given by $\dfrac{\beta_{ij}}{w_i \cdot (1-\sigma_j)}$.

Table 4. ESTIMATES FOR THE FRENCH MARKET, 1965-1987.

Sector/parameters	TEXTILES	CHEMICALS	ELECTRICAL
Estimates with homogeneous products ($\sigma_j = \infty$):			
β_{NN}	-0.177	-0.15	-0.213
	(-14.3)	(-3.9)	(-9.9)
β_{NE}	0.091	0.11	0.08
	(10.6)	(3.7)	(5.9)
β_{NF}	0.085	0.040	0.133
	(8.3)	(2.8)	(12.3)
β_{EE}	-0.117	-0.037	0.007
	(-3.7)	(-0.9)	(0.5)
β_{EF}	0.026	-0.073	-0.087
	(0.7)	(-1.4)	(-9.4)
β_{FF}	-0.111	0.033	-0.046
	(-2.6)	(0.6)	(-5.0)
λ	0.334	0.157	0.5
	(3.0)	(2.2)	(a)
LL	177.7	177.2	156.1
Estimates with differentiated products (σ_j estimated):			
β_{NN}	-0.031	-0.008	-0.147
	(-2.0)	(-0.2)	(-6.6)
β_{NE}	-0.017	0.005	0.082
	(-1.1)	(0.2)	(8.9)
β_{NF}	0.048	0.003	0.066
	(4.3)	(0.2)	(2.9)
β_{EE}	0.063	-0.004	0.011
	(2.5)	(-0.02)	(0.7)
β_{EF}	-0.046	-0.001	-0.092
	(-3.1)	(-0.2)	(-6.5)
β_{FF}	-0.002	-0.002	0.026
	(-0.4)	(-0.2)	(1.2)
λ	0.659	0.431	0.5
	(4.6)	(3.4)	(a)
σ_N	1.371	1.016	3.79
	(14.5)	(12.0)	(1.5)
σ_E	1.181	1.016	1.604
	(22.2)	(a)	(8.6)
σ_F	1.181	1.044	3.79
	(a)	(4.4)	(a)
LL	183.9	178.9	160.9

Note: Student-t are in parenthesis. LL: Log of likelihood function.

N: National, E: other European, F: non-European; (a) The parameter was constrained.

Table 5. ESTIMATES FOR THE GERMAN MARKET, 1965-1987.

Sector/parameters	TEXTILES	CHEMICALS	ELECTRICAL
Estimates with homogeneous products ($\sigma_j = \infty$):			
β_{NN}	-0.422 (-5.8)	-0.125 (-2.2)	-0.203 (-0.9)
β_{NE}	0.160 (3.1)	0.048 (0.8)	-0.154 (-0.6)
β_{NF}	0.263 (3.0)	0.077 (1.3)	0.357 (1.1)
β_{EE}	0.048 (0.5)	-0.084 (-0.5)	0.122 (0.6)
β_{EF}	-0.208 (-1.7)	0.035 (0.2)	0.031 (0.2)
β_{FF}	-0.055 (-0.3)	-0.112 (-0.7)	-0.389 (-0.9)
λ	0.098 (1.9)	0.066 (1.0)	0.030 (0.9)
LL	170.5	181.7	182.6
Estimates with differentiated products (σ_j estimated):			
β_{NN}	-0.125 (-1.9)	-0.003 (-0.1)	-0.067 (-1.7)
β_{NE}	0.049 (1.8)	0.000 (0.1)	0.013 (0.9)
β_{NF}	0.077 (1.5)	0.003 (0.1)	0.054 (1.6)
β_{EE}	-0.033 (-0.7)	0.002 (0.1)	0.000 (0.0)
β_{EF}	-0.016 (-0.5)	-0.002 (-0.1)	-0.013 (-1.5)
β_{FF}	-0.061 (-1.6)	-0.001 (-0.1)	-0.041 (-1.6)
λ	0.264 (2.8)	0.145 (2.3)	0.213 (2.7)
$\sigma_N=\sigma_E=\sigma_F$	1.466 (3.8)	1.005 (18.5)	1.275 (6.3)
LL	172.2	184.8	182.8

Note: Student-t are in parenthesis. LL: Log of likelihood function.

N: National, E: other European, F: non-European.

Table 6. ESTIMATES FOR THE U.K MARKET, 1965-1987.

Sector/parameters	TEXTILES	CHEMICALS	ELECTRICAL
Estimates with homogeneous products ($\sigma_j = \infty$):			
β_{NN}	-0.160	-0.045	-0.203
	(-3.5)	(-3.1)	(-5.3)
β_{NE}	0.070	0.048	0.079
	(3.5)	(4.8)	(5.4)
β_{NF}	0.091	-0.003	0.124
	(2.1)	(-0.6)	(4.8)
β_{EE}	-0.314	-0.019	-0.002
	(-1.5)	(-0.4)	(-0.1)
β_{EF}	0.245	-0.029	-0.077
	(1.1)	(-0.7)	(-3.3)
β_{FF}	-0.335	0.032	-0.047
	(-1.3)	(0.8)	(-1.8)
λ	0.136	0.535	0.205
	(1.9)	(3.9)	(2.4)
LL	153.7	133.1	152.7
Estimates with differentiated products (σ_j estimated):			
β_{NN}	0.100	0.016	-0.121
	(1.6)	(0.6)	(-4.4)
β_{NE}	-0.055	-0.010	0.037
	(-1.9)	(0.6)	(2.7)
β_{NF}	-0.044	-0.006	0.084
	(-1.3)	(-0.6)	(5.3)
β_{EE}	0.027	0.001	0.029
	(1.8)	(0.3)	(1.4)
β_{EF}	0.028	0.009	-0.066
	(1.3)	(0.7)	(-3.6)
β_{FF}	0.016	-0.003	-0.018
	(1.1)	(-0.9)	(-1.3)
λ	0.354	0.608	0.300
	(4.7)	(4.8)	(3.2)
$\sigma_N=\sigma_E=\sigma_F$	0.824	0.976	1.788
	(10.1)	(3.9)	(4.9)
LL	158.8	138.3	155.3

Note: Student-t are in parenthesis. LL: Log of likelihood function.

N: National, E: other European, F: non-European.

Table 7. ESTIMATES FOR THE ITALIAN MARKET, 1969-1987.

Sector/parameters	TEXTILES	CHEMICALS	ELECTRICAL
Estimates with homogeneous products ($\sigma_j = \infty$):			
β_{NN}	-0.034 (-6.6)	-0.067 (-2.1)	-0.050 (-2.9)
β_{NE}	0.014 (4.7)	0.035 (1.4)	0.023 (2.0)
β_{NF}	0.020 (6.8)	0.031 (3.0)	0.027 (3.9)
β_{EE}	0.017 (1.1)	0.030 (0.4)	0.030 (1.2)
β_{EF}	-0.031 (-1.8)	-0.065 (-1.0)	-0.053 (-2.4)
β_{FF}	0.011 (0.6)	0.033 (0.6)	0.026 (1.3)
λ	0.420 (3.3)	0.2 (a)	0.462 (3.5)
LL	158.2	128.6	120.2
Estimates with differentiated products (σ_j estimated):			
β_{NN}	-0.035 (-5.4)	-0.046 (-1.1)	0.060 (3.8)
β_{NE}	0.009 (1.5)	0.014 (0.4)	-0.038 (-3.7)
β_{NF}	0.026 (3.8)	0.032 (1.9)	-0.023 (-3.6)
β_{EE}	0.020 (1.2)	0.043 (0.6)	0.024 (3.4)
β_{EF}	-0.030 (-1.7)	-0.057 (-0.9)	0.014 (3.6)
β_{FF}	0.003 (0.2)	0.025 (0.5)	0.009 (3.4)
λ	0.349 (2.6)	0.2 (a)	0.725 (5.5)
$\sigma_N=\sigma_E=\sigma_F$	1.848 (1.9)	1.451 (1.6)	0.938 (87.0)
LL	159.0	128.7	126.8

Note: Student-t are in parenthesis. LL: Log of likelihood function.

N: National, E: other European, F: non-European; (a) The parameter was constrained.

4. SIMULATION OF THE EFFECTS OF '1992' MARKET INTEGRATION

4.1 SUBSTITUTABILITY PARAMETERS CHARACTERIZING THE COMPETITION IN EACH MARKET.

Using the previous estimates it is possible to calculate all the substitution parameters characterizing the nature of competition in each market. As discussed above, we report Morishima partial elasticities of substitution M_{ij} derived from equation (10b), given the estimates of coefficients β_{ij}. The M_{ij} are not constant over the period and are asymmetric; each measures the impact of a change of producer *i* price over the market share ratio between *i* and *j* , all other prices being held constant but all quantities adjusting to their optimal levels. The point estimates of this parameter for the year 1987 are shown in Table 8. From equation (11) above, one minus the Morishima elasticity can be interpreted as the impact of relative prices over the corresponding market share ratio; hence, a value of M_{ij} greater than one indicates that a decrease in relative prices induces a market share gain. Given this appealing interpretation, it is easier to design an alternative hypothesis on the value of these parameters rather than on the values of the coefficients β_{ij}.

Table 8: MATRIX OF PARTIAL ELASTICITIES OF SUBSTITUTION in 1987.
Model with differentiated products

		FRANCE			GERMANY			U.K.			ITALY		
		N	*E*	*F*	*N*	*E*	*F*	*N*	*E*	*F*	*N*	*E*	*F*
Textiles	*N*	-	0.971	1.403	-	1.459	1.549	-	0.560	0.597	-	1.150	1.332
	E	0.687	-	0.372	1.233	-	1.068	0.774	-	1.017	0.771	-	0.426
	F	1.089	0.805	-	1.387	1.158	-	0.841	1.054	-	0.998	0.607	-
Chemicals	*N*	-	1.030	1.042	-	1.004	1.034	-	0.935	0.921	-	1.124	1.389
	E	1.022	-	1.004	0.990	-	0.971	0.980	-	1.077	0.875	-	0.295
	F	1.024	1.016	-	1.014	1.000	-	1.018	1.063	-	0.808	0.560	-
Electrical	*N*	-	1.631	1.632	-	1.201	1.398	-	1.422	1.539	-	0.739	0.718
	E	1.077	-	0.392	1.019	-	0.927	0.905	-	0.585	0.835	-	1.009
	F	0.948	0.398	-	1.307	1.125	-	1.219	0.702	-	0.890	0.988	-

Note: For each market (country x product), producers (ij) are ranked in the following order: National (N), other European (E) and non-European (F). For example, the first line corresponds to the Morishima partial substitution elasticities M_{NE} and M_{NF} .

4.2 PRICE AND NON-PRICE EFFECTS

The composite price index (7) enables us to simulate shocks either to relative prices or to aggregate producer output - the latter supposed to proxy the creation of new products. Since both operate via the composite price the two shocks are qualitatively equivalent, the numerical equivalence depending on the value of the estimated intra-variety elasticity of substitution σ_j. The lower this parameter, the higher the relative impact of non-price effects. As noted above, our system embodies quite a strong impact from the differential in output growth on market shares. For example, with an intra-variety elasticity of substitution equal to 2, a 11% growth in the number of products (or firms) would lead to a decrease of the composite price index of 10%; with an elasticity of 1.5, only a 5.4% shock would be required to achieve the same shock over the composite price.

4.3 THE EFFECTS OF MARKET INTEGRATION

The purpose of this exercise is two-fold. First, it illustrates the impact of a price shock on market shares when preferences remain unchanged. Secondly, it aims to explore the impact of a particular homogenisation of tastes over European markets after the 1992 integration process. As an illustrative case, we assumed a counterfactual shock of a 10% decrease in the composite price of the National producers in each market. Two scenarios were considered:

-Differentiated tastes: this is the base case using the estimated parameters, differentiated by producer and market.

-Homogenisation of tastes: The design of this scenario relies upon the value of the Morishima elasticities of substitution. Here, the values of the Morishima elasticity are assumed to be a cross-country average of the estimated elasticities used in the first scenario. This can be viewed as a possible homogenisation of tastes across European countries whereby their behaviour with regard to the substitution between home, European and foreign goods become more similar. The values of the average Morishima elasticities are given in table 9. Given these values and the observed market shares, it is possible to derive the parameters β_{ij} of the demand system which correspond to this change in preferences (see Annex 2). The impact on market shares can then be calculated in a straightforwardly.

Table 9: MATRIX OF AVERAGE PARTIAL ELASTICITIES OF SUBSTITUTION in 1987.
Model with differentiated products

		N	*E*	*F*
Textiles	*N*	-	1.035	1.220
	E	0.866	-	0.721
	F	1.079	0.906	-
Chemicals	*N*	-	1.023	1.097
	E	0.967	-	0.837
	F	0.966	0.910	-
Electrical	*N*	-	1.248	1.322
	E	0.959	-	0.728
	F	1.091	0.803	-

Note: (see note table 8).

The results of these simulations, sector by sector, are shown in tables 10-12. The responsiveness of markets shares is higher for the Textile and Electrical goods than for the Chemicals products. Depending on the value of the partial elasticity of substitution, the price decrease can have a positive or a negative effect on National producers' market share and on the competitive position of the other-European and foreign producers. Accordingly, the effects are differentiated by market. However, as is common in econometric work on trade equations, the impacts on market shares are rather low.

As a result of the fall on the domestic producers' price, in the first scenario for the textiles industry in France, the foreign producers lose market share whereas the share of the other European producers increase slightly (for the latter producer the Morishima elasticity M_{NE} is lower than one). In Germany and Italy, National producers record market share gains over the two latter producers. For the U.K. there is an adverse effect for the National producers (the Morishima elasticities are both lower than one).

In the chemicals sector, the effects are typically very low. The Italian market is an exception, as there is a sizeable market share loss for the foreign producers.

The highest impacts are in the Electrical goods industry. In France, national producers have a 2.5 per cent increase of their market share in comparison with a 4.2 per cent market share loss for other two producers. The same pattern applies for Germany and the U.K. In Italy, the price decrease has a negative impact on the market share of national producers.

The effect of a homogenisation of tastes across European markets is shown in the third column of tables 14-16. As all markets have the same elasticities, the results are naturally much less contrasted than in the previous scenario. The magnitude of the market share deviations also tends be lower than in the preceding case, because there is a compensation between negative and positive effects across countries. Except for the textile industry, the decrease of national producers' price induces a market share loss for the other producers. In all sectors, the market share of foreign producers falls by around 2 per cent relative to the base shares. The losses for the other European producers are relatively higher for the electrical goods than for the other sectors. For chemicals, they are very small. For the textiles sector, the European producers benefit from a low substitutability with national products and gain market shares.

Even by using an arbitrary assumption on the hypothetical effect of the '1992' market integration, these experiments show that homogenization of tastes does not necessarily lead to an increase of substitution elasticities. In this case, a more homogeneous market can embody more uniform but also more rigid, responses to price changes.

Table 10. RESULTS OF THE SIMULATIONS, TEXTILES AND LEATHER.
Simulation of a -10% shock on the National producers' composite price index

Market/Producer	(shares of total demand in 1987)	(per cent deviations relative to base shares)	
	Base shares	differentiated tastes	homogenization of tastes
FRANCE			
National	64.5%	0.51	0.39
other-European	21.9%	0.82	0.03
non-European	13.5%	-3.74	-1.93
GERMANY			
National	45.9%	2.87	0.74
other-European	26.2%	-1.97	0.37
non-European	27.9%	-2.91	-1.58
U.K.			
National	62.1%	-1.70	0.49
other-European	19.7%	2.94	0.13
non-European	18.2%	2.55	-1.83
ITALY			
National	82.7%	0.45	0.24
other-European	8.3%	-1.14	-0.13
non-European	9.0%	-3.05	-2.08

Table 11. RESULTS OF THE SIMULATIONS, CHEMICALS PRODUCTS.
Simulation of a -10% shock on the National producers' composite price index

Market/Producer	(shares of total demand in 1987)	(per cent deviations relative to base shares)	
	Base shares	differentiated tastes	homogenization of tastes
FRANCE			
National	59.9%	0.14	0.18
other-European	29.6%	-0.18	-0.07
non-European	10.5%	-0.30	-0.84
GERMANY			
National	69.1%	0.05	0.15
other-European	20.8%	0.00	-0.09
non-European	10.1%	-0.31	-0.86
U.K.			
National	63.8%	-0.26	0.17
other-European	25.1%	0.42	-0.07
non-European	11.1%	0.57	-0.84
ITALY			
National	60.7%	0.80	0.18
other-European	29.1%	-0.51	-0.07
non-European	10.2%	-3.30	-0.84

Table 12. RESULTS OF THE SIMULATIONS, ELECTRICAL GOODS.
Simulation of a -10% shock on the National producers' composite price index

Market/Producer	(shares of total demand in 1987)	(per cent deviations relative to base shares)	
	Base shares	differentiated tastes	homogenization of tastes
FRANCE			
National	62.8%	2.47	1.10
other-European	20.7%	-4.18	-1.51
non-European	16.6%	-4.20	-2.29
GERMANY			
National	69.8%	1.01	0.93
other-European	12.4%	-1.11	-1.69
non-European	17.9%	-3.19	-2.46
U.K.			
National	56.0%	2.28	1.35
other-European	18.0%	-2.17	-1.26
non-European	26.0%	-3.40	-2.04
ITALY			
National	65.5%	-0.96	0.99
other-European	22.4%	1.79	-1.62
non-European	12.1%	2.01	-2.40

4. SUMMARY AND CONCLUSIONS

In this paper we have estimated a demand system for domestic, European and foreign products in the four main European markets, allowing for both price and non-price determinants of market shares. Non-price effects are related to a supply-side hypothesis of oligopolistic competition with a large number of firms, in which price and product differentiation effects are channelled through a composite price index. Such a demand system has the potential advantage, compared with usual import demand equations estimated in the literature, that the composite price index can capture secular shifts in market shares, (e.g. increase in import penetration), that usually are treated in a *ad hoc* fashion via the introduction of deterministic trends. Moreover, it is a flexible demand system which does not impose the habitual separability over national and foreign sources of domestic demand.

The empirical implementation of this model requires the use of a proxy for the number of representative firms in a given industry. We used a weighted index of industrial activity of each aggregate producer present in each market. A better approximation of this variable could be constructed by using microeconomic data at the firm level, but we were enable to do this because of lack of available data. Before the estimation of the demand system a test of the long-term relationship between relative prices and market shares was performed by means of a cointegration technique. This confirmed that in a significant number of cases such a relation does not exist, which suggests that the usual model based exclusively on pure price effects should indeed be ruled out in favour of a more general one.

The results of estimating the demand system suggest that the composite price indexes - incorporating product differentiation - may perform better in explaining market shares than relative prices alone. Secondly, we obtained plausible and significant estimates of the intra-variety elasticity of substitution. As is common in the literature on trade equations, the estimated price effects over market shares are in average rather moderate. However, one must note that the non-price effects intervene with a much higher elasticity. Finally, we attempted two simulations related to the 1992'market integration effects. We showed that a counterfactual decrease of 10% on domestic producers' price has a differentiated, but rather moderate, impact across markets. In the hypothetical case that a homogenisation of tastes would lead to the same average elasticities in all countries these effects could even be lower.

ANNEX 1: SOURCES OF THE DATA

The production data used in this paper were derived from the reconciliation of several sources, summarised in Table A.

Table A : PRIMARY SOURCES FOR PRODUCTION DATA.

Country	Data type	BDS [a]	OECD [b]	UNIDO [c]	SOEC [d]	INSEE [e]
France	value	1970-87				1963-69
	volume	1970-87				1963-69
Germany	value	1960-88				
	volume			1963-80	1980-87	
Italy	value				1980	
	volume		1980-88	1963-69		
	price		1970-88	1967-69		
U.K.	value	1970-88		1963-69		
	volume			1963-69		
	price		1970-88			

Note: The figures indicate the period for which the data source was used.

(a) Banque de Données Sectorielles, EC-DG II.

(b) OECD, IAI data base.

(c) UN, Industrial Statistics.

(d) Statistical Office of European Communities, Industrial Statistics.

(e) Institut de la Statistique et des Etudes Economiques (INSEE), Propage Model data base.

For the base year 1980, the production value was derived from the input-output table of the OSCE. Trade and production data were reconciled by using the correspondences between the Nace-Clio classification and the CITI industrial products list described in the table B.

Table B : Correspondences between the CITI, Nace-Clio and BDS Classifications.

Sectors	CITI	Nace-Clio	BDS
Textile and leather	321+322+323+324	43+44+45	14
Chemical products	351+352	25+26	17
Electrical goods	383	34	11

ANNEX 2: DERIVATION OF THE PARAMETERS OF THE DEMAND SYSTEM FROM THE MORISHIMA ELASTICITIES

Given the values of the Morishima elasticities (M_{ij}) and the constraints on the demand system, it is possible to derive the parameters of the Translog function. Indeed, from equation (10b) in the text he have:

(a1) $\quad M_{ij} = \beta_{ji}/w_j \; - \beta_{ii}/w_i \; + 1$

(a2) $\quad M_{ik} = \beta_{ki}/w_k \; - \beta_{ii}/w_i \; + 1$

The homogeneity and symmetry constraints imply that, $\beta_{ii} + \beta_{ji} + \beta_{ki} = 0$. By using this relation and rearranging (a1) and (a2) we get:

(a1)' $\quad \beta_{ji} \cdot (w_i + w_j) + \beta_{ik} \cdot w_j = w_i \cdot w_j \cdot (M_{ij} - 1)$

(a2)' $\quad \beta_{ji} \cdot w_k + \beta_{ik} \cdot (w_i + w_k) = w_i \cdot w_k \cdot (M_{ik} - 1)$

Finally, by using the adding-up condition $w_i + w_j + w_k = 1$, and solving the above system we find that:

$$\beta_{ji} = - \; w_j \cdot w_k \cdot (1 - M_{ik}) - w_j \cdot (1 - w_j) \cdot (1 - M_{ij}) \quad , \text{for } i \neq j .$$

and by symmetry $\beta_{ji} = \beta_{ij}$.

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