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AN ECONOMETRIC ANALYSIS OF PRICE DIFFERENTIALS IN THE EEC AUTOMOBILE MARKET

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Abstract

This paper investigates the sources of the observed price disparities in the EEC automobile market.

On the basis of an oligopoly model with product differentiation, this paper tests, and fails to reject, the hypothesis that automobile firms segregate national markets in the EEC.

It is found that value added tax differentials and the existence of different import restraints quotas and VERs are important contributing factors to price disparities. On the contrary, transportation costs differentials are not a significant explanatory variable. Finally, the importance of the preference for domestic products or "national" brands is assessed.

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Introduction

The existence of significant pre-tax price differentials across European automobile markets began to receive widespread attention in the early 80s. These price disparities have been documented by the Bureau of European Consumers Unions for the last few years (see BEUC, 1981, 1986). For some products and pairs of countries before tax transaction prices[1] in common currency, units differ by as much as 50%.

The Commission of the European Communities soon acknowledged that the price differentials constituted the "most complex and wide-ranging" issue facing the automobile industry (Commission of the EC, 1983, p.10). It was perceived that as long as those differentials were not due to "legitimate reactions to the economic environment," they would reflect anti-competitive behavior and jeopardize the EEC market integration policies.

A first attempt at identifying the sources of these price differentials was undertaken by Mertens and Ginsburgh (1985).[2] Using an econometric model with fixed country of sale and country of origin effects, the authors concluded that price variations across European markets were mainly due to price discrimination practices, with a minor contribution of product differentiation.

Mertens and Ginsburgh argue that price discrimination takes place because of differences in demand elasticities across national markets. Yet, this is one among several possible interpretations of the results of their essentially descriptive model.

In contrast, this paper studies the sources of price differentials between markets using an explicit model of oligopolistic interaction. On the basis of related theoretical results (see Gual, 1987), which are briefly reviewed here, we investigate the extent to which firms segment national markets and the empirical impact of value added tax (VAT) differentials, import restraints and transportation cost differentials on observed price disparities. The paper fails to reject the hypothesis that automobile producers segregate national markets, thus possibly undermining

[1] That is, list prices taking into account diverging discount practices in different Community countries (Beuc, 1986, p. 11).

[2] Ashworth et al., 1982; Locksley, 1983, and Bhaskar, 1984; have also looked at the automobile pricing issue. A central concern of these studies was the monetary explanation of price misalignments. Related work of this author (Gual, 1987, chapter 2) shows that these are not relevant contributory factors in a cross-section analysis of the problem.

the Commission's integration policies. Among other results, it is found that VAT differentials and disparities in import restraints are important contributing factors to price differentials.

The paper is organized as follows. The next section presents a short discussion of the theory underlying the econometric model (for a detailed analysis see Gual, 1987). Section III discusses the data set. In section IV we analyze the econometric specification and the estimation procedure employed. Finally, section V presents the empirical findings and their implications for price differentials.

The Theoretical Framework

We will consider a differentiated duopoly model where two firms (1, 2) produce respectively the two differentiated products x and y and are located in countries A and B. The firms serve both markets. The following direct demand system is assumed:

$$x = x\ (p,q) \tag{1}$$

$$y = y\ (p,q)$$

where (p,q) are the prices of the goods produced by the two firms. The demand system is obtained from a symmetric sub-utility function over the set of differentiated products. Apart from the usual properties we will suppose dominance of own-price effects.

We will assume that firms have constant and equal marginal costs of production[3] k. We take the number of firms as given and assume that entry is prevented either by institutional constraints or by the presence of economies of scale due to the existence of large fixed costs F. Specifically, F will be assumed to be at a level that allows both firms (and only them) to make positive profits.

The strategic variables of firms are prices and we will be considering Nash equilibria. The implications of other strategic variables and oligopoly solution concepts have been studied elsewhere (Gual, op. cit.). They do not affect the results significantly. Furthermore, short-term price competition is probably a sensible assumption for the automobile industry.

An important feature of the model is that serving the foreign market involves a constant unit transportation cost g, which is the same for both producers. Additionally we are interested in the impact of VAT rate differentials on equilibrium prices.

Thus we consider two alternative geographical structures of the market. When markets are integrated, firms play a single market game by setting a unique final price for both countries, thus treating the union as a single market.[4] Alternatively, when markets are segmented, both firms set prices for each national market.

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[3] The assumption of constant marginal costs has been central to the literature on oligopolistic rivalry in International markets although it is nothing but an instrumental one (see Brander and Krugman, 1983, footnote 4). For an application to the automobile industry it might well be adequate, especially if firms operate with excessive capacity.

[4] A different integrated markets model arises if firms choose f.o.b. prices. We do not consider here the results of this case because of the empirical goal of this paper. The system of exclusive dealerships and the pricing practices in the EEC automobile market makes the uniform delivered pricing models more adequate. Additionally, the experience of the United States automobile market shows that when firms produce in several locations, a system of uniform delivered prices is more common.

Simple comparative statics analysis has revealed a set of important results for the empirical work which we will review next for the case of linear demands. (For a more general formulation see Gual, op. cit.)

Result 1

Under market segmentation, equilibrium prices abroad differ from domestic prices by less than the transportation cost difference.

This is the "reciprocal dumping" result obtained by Brander and Krugman (1983) for the differentiated products case. This price discrimination has two different sources. First, if demand is linear,[5] we would expect smaller mark-ups in high costs (foreign) markets as in the case of a monopoly. Furthermore, being a high-cost competitor in the foreign market will imply a higher perceived demand elasticity in that market[6] and, thus, further price discrimination against domestic consumers.

Note that this price discrimination can take place even if arbitrage is feasible since, in equilibrium, the arbitrage constraints (2) do not bind.

$$|p^B - p^A| \leq g \tag{2}$$

$$|q^A - q^B| \leq g$$

where p^i, q^i (i=A, B) are equilibrium pre-tax prices and for simplicity we set arbitrage costs equal to g.

Result 2

The degree of interpenetration of markets (the volume of intra-industry trade) is lower when firms segment national markets.

With integrated markets, firms discriminate against domestic consumers by charging them phantom freight and favor distant clients who receive the benefits of freight absorption. With segmented markets the possibility of choosing an optimal discriminatory policy implies higher prices in foreign markets, thus reducing the competitiveness of exports.

As for the impact of differential taxation, the following result has been shown to hold with mild restrictions on the demand system.[7]

Result 3

Under market segmentation, tax rate differentials across countries result in less than proportional final price differentials. Therefore, pre-tax prices are lower in high-tax countries.

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[5] Actually, non-extreme convexity is enough.

[6] The perceived demand in the export market will be more elastic if the goods are strategic complements. Then, if marginal costs are constant, a decrease in the price of one good makes the demand for the other good more elastic (see Bulow et al., 1985).

[7] These restrictions relate to the degree of own price curvature of the demand function as well as to the cross price effects.

Market segmentation allows the absorption by the duopolists of part of the tax rate differential and results in discrimination against consumers in low-tax countries. Note that this discrimination will be limited by the arbitrage constraints (2) if arbitrage is feasible.

The following linear version of the model illustrates results 1 and 9 and is the basis of the econometric specification to be used in the empirical work.

The demand system will be now:

$$x = a - bp + cq$$

$$y = a - bq + cp$$

where a, b, c > 0 and dominance of own-price effects implies b > c.

This will be the world demand system when markets are integrated. The population share of the first country is θ ($0 < \theta < 1$). When markets are segmented, a similar system holds for each country.

We will assume a constant[8] unit transportation cost g and value added tax (VAT) rates t^A and t^B (where without loss of generality we assume $t^A > t^B$).

The profit functions corresponding to the integrated markets case under uniform delivered pricing are:

$$\pi_1 = [p - (k + (1-\theta)g)]\ \{\theta[a - bp(1+t^A) + cq(1+t^A)] +$$

$$+ (1-\theta)\ [a - bp(1+t^B) + cq(1+t^B)]\} - F$$

$$\pi_2 = [q - (k + \theta g)]\ \{\theta[a - bq(1+t^A) + cp(1+t^A)] +$$

$$+ (1-\theta)[a - bq(1+t^B) + cp(1+t^B)]\} - F$$

The first order conditions in this case are:

$$R_1 - 2bp + cq(1+t') = 0$$

$$R_2 - 2bq + cp(1+t') = 0$$

where $R_1 = a + bk(1-\theta)g(1+t')$

$$R_2 = a + bk\theta g(1+t')$$

and $t' = \theta t^A + (1-\theta) - t^B$

Equilibrium pre-tax prices will be:

$$p = [1/(4b^2 - c^2)]\ [(2bR_1 + cR_2)/(1+t')]$$

$$q = [1/(4b^2 - c^2)]\ [(2bR_2 + cR_1)/(1+t')]$$

[8] Increasing returns in the transportation technology unnecessarily complicate the algebra without yielding qualitatively different results. The empirical implications of this assumption are discussed in the Appendix.

By construction, we will not expect pre-tax price differences to be related to tax differentials and/or transportation costs.

The profit functions for the segmented markets model are:

$$\pi_1=[p^A-k]\ [a-bp^A(1+t^A)+cq^A(1+t^A)]+$$

$$+[p^B-k-g]\ [a-bp^B(1+t^B)+cq^B(1+t^B)]-F$$

$$\pi_2=[q^A-k-g]\ [a-bq^A(1+t^A)+cp^A(1+t^A)]+$$

$$+[q^B-k]\ [a-bq^B(1+t^B)+cp^B(1+t^B)]-F$$

The first order conditions for market A yield:

$$A_1-2bp^A(1+t^A)+cq^A(1+t^A)=0$$

$$A_2-2bq^A(1+t^A)+cp^A(1+t^A)=0$$

where $A_1=a+bk(1+t^A)$ and $A_2=a+b(k+g)(1+t^A)$ and the mirror image of these conditions hold for market B.

Straightforward computation yields the following pre-tax equilibrium prices:

$$p^A=[1/(4b^2-c^2)]\ [(2bA_1+cA_2)/(1+t^A)]$$

$$q^A=[1/(4b^2-c^2)]\ [(2bA_2+cA_1)/(1+t^A)]$$

By symmetry $p^B=[1/(4b^2-c^2)]\ [(2bB_1+cB_2)/(1+t^B)]$, where

$B_1=a+b(k+g)(1+t^B)$ and $B_2=a+bk(1+t^B)$.

Price differences across countries for the same good will now be related to tax rate differentials and transport costs:

$$p^B-p^A=[a/(2b-c)]\{[1/(1+t^B)]-[1/(1+t^A)]\}+(b/(2b+c)]g. \qquad (3)$$

Result 1 is satisfied since the coefficient on the transportation cost variable is positive and smaller than one under market segmentation. Similarly, Result 3 holds since $a/(2b+c)$ is positive. Then if $t^A > t^B$, one can conclude that $p^B > p^A$, on account of tax differentials.

The simple 2 countries-2 products linear model of the previous section can be extended to accommodate the case of several markets and multiproduct firms. In a multimarket-multiproduct setting with segmented markets, there will also be strategic pricing effects due to the presence of transportation costs and/or tax differentials. Price discrimination against low-tax countries hinges upon certain demand conditions and the assumptions in the interactions between firms, but in no way is dependent upon the number of firms or markets in the model. As for transportation costs, we now have a set of cost variables g_{ri} depending on goods (r,s=1,...,m) and destinations (1,h=1,...,n). Nonetheless, under market segmentation oligopolists will set prices so that marginal costs equal perceived marginal revenues in each market. Price differences across export markets will be lower than transport cost differences with smaller mark-ups in the export markets with larger transport costs.

However, in the “nxm” model, the equilibrium price of product r in market i depends not only on the transportation cost for this good but also on the costs of bringing all other goods to the market. As a result, the price differential for good r in markets i and h will also be a function of the sum of transport costs differentials for all other goods. Overall prices will tend to be higher in those markets where transport costs for most products are higher. This is shown algebraically in the Appendix for the case of a single-product oligopoly. The case of multi-product firms is also briefly discussed.

The Data Set

The price data used in this paper correspond to (before-tax[9]) transaction prices in ECUs, for June 1986. Transaction prices were computed using the Bureau of European Consumer Unions' estimates (see BEUC, 1986) of the maximum discounts available in different countries, and list prices from the same source.

The sample consists of 28 models and 8 EEC countries.[10] Since we consider price differentials, we have 28 country-pairs or cross-section units. Not all models are available in all countries and, therefore, we have an unbalanced data set. The number of observations per country-pair ranges from 11 to 28, with a total of 568 observations.

The models included in the sample are base models. Thus, we do not have to deal with the pricing of options which might be different across countries and/or models. Similarly, the sample includes only small to large passenger cars, and excludes luxury and extra-small cars which possibly constitute different product groups. The sample does not include either Japanese cars or other low-cost products from Eastern Europe. It must be recalled that the empirical analysis is undertaken under the assumption that there are no significant production cost differentials between the main volume producers in the market; including Japanese products in the sample might not be consistent with this assumption.

Tax information has been obtained from Community sources (see Commission of the EC, 1986b). Similarly, data on market shares, voluntary export restraints and quotes or other restrictions to imports have been supplied by the EEC (Commission of the EC, 1986a).

Transportation costs for the products in the sample have been predicted using an estimated transportation cost function. After consultation with several industry experts it was determined that the main determinants of shipping costs were the distance that the product is to be shipped and its weight. The function was estimated using proprietary data supplied by two of the main manufacturers in the market. Details on the data, the functional form and the estimation procedure are provided in the Appendix.

Econometric Specification

(i) The model: demand asymmetries and stochastic assumptions

The econometric regression model that we specify is derived from the reduced form of a linear model of oligopolistic competition, a la Bertrand, as described above. Since the theoretical

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[9] Only value added taxes are excluded.

[10] Belgium, Denmark, France, Germany, Luxembourg, Italy, Netherlands and the United Kingdom.

framework is a highly stylized symmetric model, some adjustments are made to take into account the impact of possible demand side asymmetries as well as the contribution of other important factors that are not explicitly incorporated in the theoretical model.

The stochastic structure of the econometric specification is derived as follows. It is assumed that the observed prices for product r (r=1,...,m) in market i (i=1,...,n) have a deterministic component and a set of random elements. The non-stochastic component will be the equilibrium price and it is a function of the exogenous variables of the model (marginal production costs k; tax rates ti; transportation costs g_{sh}[11] market sizes θ_i), and a set of parameters Ω.

For the nxm model under market integration, this deterministic component can be written as follows:

$$P_r = P_r\ (k,t_i,\theta_i,g_{sh};\ \Omega);\ r,s=1,\dots,m. \qquad (4)$$

$$i,h = 1,\dots,n.$$

where P_r=P_{ri} for all i.

Similarly, with segmented markets we will have:

$$P_{ri}= P_{ri}\ (k,t_i,g_{sh};\ \Omega);\quad r,s=1,\dots,m.\quad i,h=1,\dots,n. \qquad (5)$$

Equations (4) and (5) are the central regression relations used in the econometric specification. Both are embodied in the following general specification which includes an additive error term:

$$P_{ri}= P_{ri}(k,t_1,\theta_1,g_{sh};\ \Omega)\quad +\tau_{ri};$$

$$i=1,\dots,n.\ r=1,\dots,m.$$

We assume that the error term has three components:

$$\tau_{ri}= er + \in_i + u_{ri}\quad ;\quad i=1 \ldots n.\ r=1 \ldots m.$$

where e and $\in$ correspond to product – and country – specific random terms and u is a purely observational error. It is assumed that the three variables are independent and normally distributed. They are also uncorrelated across observations. These and further assumptions can be summarized as follows:

$$E(e_r)=E(\in_i)=E(u_{ri})=0$$

$$\text{cov}\ (e_r,e_s) = \sigma_e^2 \text{ for } r=s$$

$$= 0 \text{ otherwise}$$

$$\text{cov}\ (\in_i,\in_h) = \sigma_\in^2 \text{ for } i=h$$

$$= 0 \text{ otherwise}$$

$$\text{cov}\ (u_{ri},u_{sh}) = \sigma_u^2 \text{ for } i=h;\ r=s$$

$$= 0 \text{ otherwise}$$

[11] g_{sh} is the unit transportation cost of shipping product s from its country of production to country h.

$cov\ (e_r, \in_i) = 0$ for all r,i; $cov\ (e_r, u_{si}) = 0$ for all r,s,i;

$cov(\in_i, u_{rh}) = 0$ for all i,h,r.

The estimated model considers price differences across countries for the same product. As a consequence, the stochastic structure is simplified and only two error components remain. One is purely observational and the other is associated with the cross-section unit being considered. Additionally, considering price differentials across countries for the same product implies a more direct emphasis on the problem of price differentials within the Community countries and permits the examination of the sources of those differentials.[12]

The estimated regression model is the following:

$$PD_{rj}=\beta_1 TD_{rj}+\beta_2 TCD_{rj}+\beta_3 0TCD_{rj}+\beta_4 DOM_{rj}+\beta_5 VER_{rj}+\beta_6 SCR2_{rj}+w_{rj} \tag{7}$$

where r,s=1,...,28 indexes products and j,k=1,...,28 indexes country-pairs (i,h).

The variables are defined as follows:

PD_{rj} is the price differential for product r between countries i and h: $P_{ri}–P_{rh}$;

TD_{rj} is the difference of the reciprocals of 1 plus the tax rate, multiplied by 100.

That is: $100\{[1/(1+t_i)\ ] – (1/(1+t_h)]\}$

TCD_{rj} is the difference in transportation costs of product r to both countries i and h. That is, $g_{ri}–g_{rh}$.

$0TCD_{rj}$ is the summation of TCD_{rj} over all products but r within each cross section unit.

DOM_{rj} is a dummy variable that takes value 1 if the product is produced in country i and –1 if it is produced in country h. It is zero whenever the product is not produced in any of the countries or in both of them.

VER_{rj} is the difference between the degree of penetration of Japanese imports in markets h and i.

$SCR2_{rJ}$ is the difference in the two-seller concentration ratio in countries I and h, multiplied by 100.

The error term $w_{rj} = \in_i - \in_h + u_{ri} - u_{rh}$ implies that there is a constant correlation across observations corresponding to the same cross-section unit. That is, we will have:

$$\begin{aligned} cov\ (w_{rj}, w_{sk}) &= 2\sigma_\in^2 + 2\sigma_u^2 \text{ if } r=s,\ j=k \\ &= 2\sigma_\in^2 \quad \text{if } r=s,\ j\neq k \\ &= 0 \text{ otherwise} \end{aligned}$$

The resulting variance covariance matrix is block diagonal with a representative matrix of the form:

$$\Phi_j = \sigma_u^2\ I + \sigma_\in^2\ J.$$

[12] Furthermore, considering price differentials implies that, given our assumptions on costs, we do not need production cost data, which in any case are difficult to obtain.

where I is the Identity matrix and J is a square matrix of ones with nj rows, as many as the number of observations in cross-section unit j.

The variables DOM, VER and SCR2 attempt to capture the potential impact on price differentials of some aspects of demand asymmetries and diverging competitive conditions, which can be important under a regime of market segmentation. In particular, a more inelastic demand for domestic products can result in comparatively higher prices for domestic goods being compatible with a leading domestic market share.

Similarly, the impact of import restraints on equilibrium prices can be important. Other things being equal, we would expect higher equilibrium prices in those markets where import competition is restricted either by means of a quota or a voluntary export restraint (see Harris, 1985, and Krishna, 1985).

Finally, the inclusion of the seller concentration ratio attempts to capture the possible impact of varying degrees of collusion at the member state level. This was ruled out by the assumption of Bertrand competition and, although no attempt is made to specify alternative conjectural variations, this variable should capture the effect of varying degrees of competition on equilibrium prices.

It must be finally repeated that all these variables should not contribute significantly to price variation under a regime of market integration but they could be important if markets are segmented.

(ii) Estimation procedure

Efficient parameter estimates for model (7) can be obtained by Estimated Generalized Least Squares (EGLS). Since our sample size is large, we can use asymptotic results and, under these conditions, the properties of the EGLS estimator are those of the Aitken or true GLS estimator.

A variance components model of this sort can be estimated either by direct maximization of the likelihood function or by a two-step procedure whereby initial unbiased estimates of the variance components are first obtained. In a second stage, the slope parameters are derived using the estimated variance-covariance matrix. Again, for large samples this two-step procedure is equivalent to the maximum likelihood estimator (Harvey, 1981, pp. 138-141). We have chosen to work with a two-step method due to its computational simplicity. Furthermore, it does not require the assumption that the errors are normally distributed.

The use of a random effects specification merits some remarks. Modeling country-pair effects as random effects as opposed to fixed effects is unavoidable given our interest in explaining both the between- and the within-country-pair variation in the data. The fixed effects or Least Squares Dummy Variables (LSDV) model removes all variation between cross-section units. Since some of the right-hand side (RHS) variables in model (7) do not vary much or at all within the cross-section unit, their effect is absorbed by the dummy coefficients, and the parameters of interest cannot be estimated by this method.[13]

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[13] Additionally, the fixed effects method accounts for cross-section specific variation but does not explain the sources of this variability. The dummy coefficients do not have a clear interpretation. Similarly, another drawback of the fixed effects specification is that it results in an important loss of degrees of freedom.

On the other hand, OLS represents a possibly inefficient weighting of the between and the within variation in the data. In fact, the variance components model that results from the specification of random effects leads to an estimator which can be shown to be an efficient combination of the two types of variation in the data (Maddala, 1971).[14]

An important restriction that has to be satisfied by the random effects model is that the country-pair specific error should be uncorrelated with the RHS regressors. Inconsistent parameter estimates would obtain if the error component included left-out regressors correlated with included variables. In the present case we do not expect the orthogonality assumption to be violated since the random specification arises not because of a missing regressor problem but rather for efficiency reasons.

Several methods of estimation of the variance components have been suggested in the literature (for a comparative analysis of their relative performance, see Maddala and Mount, 1973).

Most methods involve the use of mean square errors from the OLS, LSDV and Least Squares Between Groups (LSBG) models to generate estimates of the variance components. In particular, the analysis-of-variance technique equates mean square errors for the LSBG and LSDV regressions to their expected values. However, when the data set is unbalanced, the resulting estimators are biased (Searle, 1971). For this reason we have used the fitting-of-constants method developed by Henderson (Searle, op. cit., chapter 10). With this methodology, reductions in sum of squares due to fitting different parts of the model are set equal to their expected values, which include constants and are linear in the variance components. Solving the resulting system of equations yields the variance parameter estimates. Fuller and Battese (1974) provide the solution of that system for a model like (7) where there are two variance components and some of the independent variables have to be dropped in the LSDV specification because they do not vary within cross-section unit or they are linearly related to the "fixed effects." The estimators derived by Fuller and Battese are:

$$s_u^2 = u'u/N - d - z_1 + z_2$$

$$s_\in^2 = \{\in'\in - (N-z)\, s_u^2\}/(N - \text{trace}\ [\ (X'X)^{-1}\Sigma_{jnjxj'xj})\}$$

where u is the vector of residuals from the LSDV regression, ε is the residual vector for OLS and the other variables and parameters are defined as follows:

N: total number of observations,

z: total number of regressors,

z_1: number of regressors that are not linear functions of the fixed effects in the LSDV specification,

z_2: number of regressors which are linear combinations of the fixed effects in the LSDV specification,

d: number of cross-section units,

[14] The EGLS estimator is also a matrix weighted average of the LSDV and a related estimator, the Least Squares Between Groups estimator, which regresses cross-section means of the dependent variable on cross-section means of the independent variable, thereby accounting only for between-cross-section unit variation in the data.

X: is the (N x z) matrix of regressors used in the OLS procedure, and

x_j: is the (1 x z) vector of means for cross-section unit J.

Once the variance components have been consistently estimated, one can proceed to estimate the GLS model using the estimated variance-covariance matrix. However, since this involves the inversion of a (568 x 568) matrix, estimation has been carried out using a transformation of the data which results in a scalar variance covariance matrix.[15] This transformation requires taking weighted deviations from cross-section means where the weights are obtained as follows (Fuller and Battese, op cit., p. 628):

$$\alpha_j = 1 - \{s_u^2/(s_u^2+njs_\epsilon^2)\}^{1/2}$$

V. Estimation Results and Implications for Price Differentials

Before proceeding to the estimation of the model, a preliminary test was performed to check the stochastic specification. The null hypothesis that no random effects are present can be tested using a conventional F test like the one that would be used if the effects were not random (Judge et al., 1980, p. 336). For the country-pair effects, the value of the F statistic was 2.055 with (25,537) degrees of freedom. This statistic is significant at a 99% confidence level, so that the null hypothesis of no country pair effects is rejected.

Given the two-way structure of the data, a preliminary test to ascertain the presence of product effects was also run. These product effects are assumed to be fixed as opposed to the random country-pair effects. This mixed model specification is consistent with our theoretical framework. According to our specification there are no random product effects since these cancel out when we take price differences across countries for the same product. Modeling product effects as fixed effects is equivalent to disregarding the variation of the data between products and concentrating on the within-product variation. Specifically, the dummies could capture left-out regressors that might account for model-specific variations in the data that are left unexplained by our theoretical model. The result of the F test for the products effects model were somewhat inconclusive. The F statistic was 1.591 with 28 and 535 degrees of freedom. This value is not significant at a 99% confidence level but it is at a 95% level. As we will see, these product effects do not seem to be quantitatively very important but they do provide a way to analyze the sensitivity of our results to the problem of left-out regressors.

The preliminary regressions also indicated the presence of high collinearity among the RHS variables. Even though partial correlation coefficients were low, the determinant of the correlation coefficient matrix was close to zero (0.026). Further examination of the data revealed that the presence of the SCR2 variable was causing the correlation problem. The determinant of the coefficient correlation matrix improved to (0.21), when this variable was deleted.

Additionally, the coefficient of the variable had a negative sign and was not significantly different from zero in the OLS regression (under the Ho that $\sigma_\epsilon^2=0$). For these reasons it was decided that the variable would be dropped in further analysis.[16]

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[15] Direct EGLS procedure was used for the small sample where Denmark, Italy and the United Kingdom are excluded and the sample size is reduced to 253 observations.

[16] The F tests that have been reported correspond to the regressions where SCR2 was already not included.

The estimation of the variance components for the full sample yielded the results indicated in Table I.

Table I

Variance Components Estimates (Full Sample)

	Country-pair error	Pure error	Ratio $\sigma_\epsilon^2/\sigma_u^2$
Model without product effects	109772.6	570234.7	.1925
Model with product effects	85530.9	503822.9	.1698

(Reduced Sample)

	Country-pair error	Pure error	Ratio $\sigma_\epsilon^2/\sigma_u^2$
Model without product effects	133569.7	193047.8	.6919

The values of the weighting parameters αj ranged between .437 and .604 when the product effects were not present, and between .410 and .583 when those effects were included. We expect, therefore, that the EGLS model will give results different from both the LSDV model (weight of one), and the OLS model (weight of zero). This is confirmed by the point estimates presented in tables II and III, respectively for models without and with product effects. The results for the OLS, LSDV, LSBG and the EGLS models are reported to facilitate the analysis of the estimates.

Table II

Parameter Estimates. Models without product effects

	Estimation Procedure			
Variables	**OLS**	**LSDV**	**LSBG**	**EGLS**
TD	18.50	-64.89	13.47	17.89
	-2.844	(58.82)	(.7507)	-5.754
TCD	.4001	-1.087	-36.27	-.3550
	(.6559)	(.6626)	-2.016	(.6462)
OTCD	.6178		2.694	.5935
	(.0989)	(.1130)	(.1916)	
DOM	501.52	165.2	1305.7	247.62
	(70.71)	(80.73)	(31.01)	(77.25)
VER	25.33		19.92	27.24
	-3.885		(.9835)	-8.158
Adjusted R^2	.4470	.5285	.9368	.5242

Note: The figures in parenthesis are standard errors. For each model these standard errors are valid only under the hypothesis that the corresponding specification is correct.

LSDV: Least Squares Dummy Variables, or the "within" estimator. LSBG: Least Squares Between Groups, or the "between" estimator. EGLS: Estimated Generalized Least Squares.

Table III

Parameter Estimates. Models with product effects

	Estimation Procedure			
Variables	**OLS**	**LSDV**	**LSBG**	**EGLS**
TD	10.42	-193.4	12.09	9.045
	-2.947	(65.56)	(.8215)	-5.577
TCD	.1695	-.6182	-33.72	-.3795
	(.6890)	(.6947)	-2.104	(.6820)
OTCD	.3358	–	2.485	.2782
	(.1035)		(.1221)	(.1900)
DOM	426.16	217.0	1212.1	262.72
	(73.86)	(81.59)	(39.04)	(78.95)
VER	18.79	–	19.02	20.28
	-3.827		-1.027	-7.694
Adjusted R^2	.5196	.5841	.9364	.5759

Note: The figures in parenthesis are standard errors. For each model these standard errors are valid only under the hypothesis that the corresponding specification is correct.

LSDV: Least Squares Dummy Variables, or the "within" estimator. LSBG: Least Squares Between Groups, or the "between" estimator. EGLS: Estimated Generalized Least Squares.

Several comments on the relative performance of the estimators are necessary before interpreting the specific parameter estimates.

First of all, there is not much difference between the estimates when the product effects are included or removed. The main differences affect the TD and OTCD parameters, which are substantially reduced for both the OLS and the EGLS model. Excluding the product effects from the regression might create bias in the parameter estimates of the included regressors if those effects are correlated with included explanatory variables. It is not clear that this is a problem in our model. And, if the effects are in fact unnecessary regressors, the estimates reported in Table III are inefficient and we are likely to reject the null hypothesis more often than we should. Overall, the magnitude of the product effects does not seem to be very relevant due the relatively small changes in adjusted R^2. We will thus concentrate on the point estimates reported in Table II.

As we pointed out when discussing the LSDV model, due to the characteristics of the data this procedure cannot provide estimates of the OTCD and VER coefficients. When we take deviations from country-pair means, TCD and OTCD are perfectly correlated and, therefore, the parameter estimate obtained for TCD in the LSDV model cannot be compared with those obtained with the other techniques. Similarly, the variation in the tax variable is mostly between cross-section units and is almost completely removed by the within estimator, which results in unreliable estimates of the tax parameter.[17] For our sample, almost 65% of the total variation in the dependent variable is within cross-section unit variation and 35% is between variation.[18]

[17] In the LSDV model, the country-pair intercepts capture unexplained variation between cross-section units; that is, the influence of variables like TD and VER.

[18] The within variation is computed as follows: $Wyy=\Sigma_r\Sigma_j(y_{rj}-y_j)^2$, where y is the dependent variable and y_j are cross-section means. Byy is the between variation: $BYY= \Sigma_r\Sigma_j\, y_j^2$.

The between-group regression gives results similar to the OLS regression for the variables that do not vary much or at all within cross-section unit (TD and VER). For the rest of the parameters, the within variation is important and, by removing it, this method gives widely different results with respect to both the OLS and the EGLS methods. Finally, the LSBG method results in artificially low standard errors because, for each cross-section unit, as many observations as data in the original sample are included.

The EGLS method provides results which are intermediate between both the OLS and LSDV model and the LSDV and LSBG model. This is consistent with the theoretical proposition which asserts that the EGLS estimator can be viewed as an efficient use of both the within and the between variation in the data.

The F test on the significance of the regression as an explanation of the behavior of the LHS variable is performed first. This corresponds to the test of the null hypothesis of market integration. For all models and under the alternative stochastic specifications, the null hypothesis is rejected at a 99% confidence level.

As for the different parameters in the model we have the following results.

The tax variable is a significant explanatory variable for most models (with the exception, of course, of the LSDV model). The sign of the effect is consistent with the theoretical proposition that higher tax rate differentials result in increasing discrimination against consumers in low tax countries (recall that $TD_{ij} = 100 \{[1/(1+t_i)] - [1/(1+t_h)]\}$). The evaluation of the magnitude of the effect deserves some comments. A 1% increase in the tax variable results in an increase in the price differential of 17.89 ECUs. However, TD does not correspond strictly to tax differentials of the form $(t_h - t_i)$.[19] The impact of a 1% increase in $(t_h - t_i)$ will depend on the current levels of taxation, with a larger effect when taxation rates are low. For example, the impact of a 1% increase in the tax differential when the differential is of 20 points will be of 14.91 ECUs if the tax rates are 20% and 0 but only of 10.65 ECUs when the tax rates are 40 and 20%.

The magnitude of the tax effect is also consistent with the theoretical prediction that a 1% increase in taxes leads to a decrease in pre-tax prices which is proportionally smaller (Gual, op. cit., p. 45). Our specification does not allow a direct test of this proposition. However, if the impact of a 1% increase in (t_h-t_i) is around 10 to 15 ECUs, this figure is clearly smaller than 1% of the average car price in the sample (7323 ECUs). Note also that the average value of TD in the sample is 10.78, which will result in a price differential of 192.85 ECUs (the average price differential in the sample is 677.91 ECUs and the average predicted price differential is 527.88 ECUs).

The coefficient for the VER variable indicates that the influence of competitive conditions as related to the presence of Japanese imports is important. A unit increase in the penetration of low-cost imports is predicted by the model to result in a 27.24 ECU increase in the price differential. For our sample, the average value of the VER variable is 9.104, which would correspond to a price differential of 247.99 ECUs. As we mentioned earlier, the ad hoc characteristics of the inclusion of this variable in the specification should caution the interpretation of the magnitude of the coefficient. As the discussion of the tax coefficient indicates (footnote 19), alternative specifications could substantially alter the results.

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[19] Using (t_h-t_i) as a regressor would imply an erroneous specification of the model, since we would be disregarding the nonlinear relationship between (t_h-t_i) and price differentials. Under the linear specification of model (7), regressing PD_{rJ} on (t_h-t_i) would result in biased and inconsistent parameter estimates.

Overall, the transportation cost variables do not seem to be of great importance. For the TCD variable, the null hypothesis that the coefficient is not significantly different from zero cannot be rejected for both EGLS models with and without product effects. The results for the OLS model are similar. Although Table II shows the standard errors for the OLS specification, the true standard errors are likely to be higher (see Table IV below), and there is no doubt that the coefficient is not significant. The OTCD parameter is significant in the model without product effects but its magnitude is very small. Transportation costs do not appear to play a significant role in segmenting national markets. This is probably due to the fact that, for this industry, almost all firms operate several plants which are geographically dispersed and sometimes located in more than one member state.

Finally, the DOM variable contributes significantly to the explanation of observed price differentials. This coefficient shows that the same product will be sold with a mark-up of 247.62 ECUs in the country where it is manufactured. The result clearly indicates that domestic firms are able to profit from market segmentation by discriminating against domestic consumers who are willing to pay more for domestically produced products.

To further check the robustness of our results, the same estimation techniques were applied to a reduced sample, where the data for Denmark, Italy and the United Kingdom were deleted. As a result, the sample was reduced to 253 observations with 10 cross-section units. Denmark was excluded on account of its extremely high tax rates, while. Italy and the United Kingdom were dropped because for both countries it is commonly perceived that the domestic product factor plays an important role. The objective of considering this smaller market within a market was to check to what extent the core of Community countries form an integrated market. The variance components results for the small sample are presented in Table I and the parameter estimates in Table IV. For brevity, only the OLS and EGLS models without product effects are reported and, for the OLS regression, the standard errors under alternative hypothesis about the variance covariance matrix are included.

Reducing the sample does not significantly change the point estimates obtained with the full sample. The standard errors increase because of the smaller sample size. As a result, only the VER coefficient is significant in the EGLS regression. For the OLS model, both DOM and VER are significantly different from zero when the true standard errors are considered.[20]

The point estimates for the DOM parameter is significantly reduced for the EGLS model, both as compared to the full sample result and to the OLS estimate with a reduced sample. A decreased importance of the DOM parameter is consistent with the hypothesis that domestic product effects are important factors in Italy and the United Kingdom, countries excluded from the reduced sample. Moreover, the large divergence between the OLS and the EGLS estimates for this parameter is due to the fact that, for the email sample, the EGLS estimator weights the within variation more heavily.

Overall, the analysis of the smaller sample confirms that two of the main factors in price disparities across countries in the European Community car market are: a) the lack of uniform taxation on value added at an EEC level, and b) the different national policies versus extra-EEC low-cost imports undertaken by EEC member states. Domestic product preferences are found to be an important phenomenon possibly associated with the Italian and British markets. The cost

[20] As expected, the correct standard errors are larger than those of the EGLS model. Overall, OLS standard errors underestimate the true errors by a factor of two.

of shipping finished products is not found to be a contributory factor to the fragmentation of the EEC market. This segmentation of the market at a member state level is ascertained, in particular for the model with all eight EEC countries considered.

Table IV

Parameter Estimates. Model without product effects (reduced sample)

	Estimation Procedure	
Variables	**OLS**	**EGLS**
TD	22.6091	19.8228
	(17.726)	(17.095)
	(5.1319)	
TCD	.9292	-1.6629
	(1.3434)	(1.0971)
	(1.0052)	
OTCD	1.2262	1.1088
	(.7231)	(.7088)
	(.2029)	
DOM	460.503	25.455
	(155.153)	(70.003)
	(67.079)	
VER	31.0104	35.4315
	(11.326)	(1ls231)
	(3.2597)	
Adjusted R^2	.4675	.3867

Note: The figures in parentheses are standard errors under the hypothesis that the country-pair variance component is significantly different from zero. For the OLS procedure, the figures in brackets are standard errors under the hypothesis that there is no country-pair variance component.

Appendix 1

The n x n linear model

Consider a model with n single-product firms (r,s=1,...,n), located in countries i,h=1,...,n. Prices and quantities are denoted by pr and x_r and the following linear demand system is assumed:

$$x_r = a-bp_r + C\Sigma s\neq rp_s \quad r,s = 1,...,n.$$

The unit transport coat of bringing product r to market i is assumed constant and equal to gri. The marginal cost of production is constant and the name for all firms (k), and 8_i denotes relative country size.

For the integrated markets case we have the following maximization problem for firm r:

$$\max\ (p_r-k-\Sigma_i\theta_i g_{ri})\ (a-bp_r+c\Sigma_s\neq rp_s)$$

which results in the following first order conditions:

$$-2bp_r+c\Sigma_s\neq rp_s = -(a+bk) - b\ \Sigma_i(\theta_i g_{ri}) \quad r=1,..., n$$

or in matrix notation:

$$Ap = -(a+bk)\ e - b\ \Sigma\ \theta_i g_i \qquad \text{(A.1.)}$$

where p and g_i are n-dimensional vectors of prices and transport costs and e is a vector of ones; A=–(2b+c)I+c J; J is a n-dimensional matrix of ones and I is the identity matrix.

Solving system (A.1.) we obtain the equilibrium price for a representative product:

$$p_r=[1/D]\{(2b+c)(a+bk)-b(2b-(n-2)c)\Sigma_i\theta_i g_{ri}-cb\Sigma_s\neq r\Sigma_h\theta_h g_{rh}\}$$

where D=(2b+c) (2b–c(n–1))

In the segmented markets model, the maximization problem for firm r in market i is the following:

$$\max p\ (pr^i-k-g_{ri})\ (a-bp_r^i+c\Sigma s\neq rp_s)$$

The first order conditions for market i in matrix notation will be:

$$Ap^i = -(a+bk)\ e - b\ g_i \qquad \text{(A.2.)}$$

where p^i and g_i are vectors of prices and transportation costs for the ith market. A representative equilibrium price for market i will be:

$$pr^i=[1/D]\{(2b+c)(a+bk)-b(2b-(n-2)c)g_{ri}-cb\Sigma_s\neq rg_{si}\}$$

Systems A.1. and A.2. reveal that equilibrium prices in the integrated markets case are linear combinations of equilibrium prices for the segmented markets model. If we look at cross-country price differences for the same product we obtain:

$$pr^i-pr^h=[1/D]\{b(2b-(n-2)c)(g_{ri}-g_{rh}) + cb(\Sigma_s\neq r\ g_{si}-\Sigma_s\neq g_{sh})\} \qquad \text{(A.3.)}$$

Whenever $g_{si}=g_{ri}$ and $g_{sh}=g_{rh}$ for all s and r, we obtain the 2x2 model result (3). In general, however, we should take into account the impact of the second term in A.3. on equilibrium prices.

Appendix 1 (continued)

Finally, the fact that firms generally produce several differentiated products merits some consideration. The multiproduct firm takes into account the impact of its pricing behavior on the profits derived from selling other differentiated products. The transport costs differentials coefficients in equation A.3. will be different. In particular, the price differential will be a function of a weighted sum of transportation cost differentials of other products since products manufactured by the same firm will play a distinctive role. The use of (g_{ri}-g_{rh}) and ($\Sigma_s \neq r\ g_{si}$– $\Sigma_s \neq r g_{sh}$) as regressors in the econometric specification is therefore an approximation which should be taken into account when we evaluate the parameter estimates.

Estimation of the transportation cost function

The transportation cost data that we require for the empirical analysis is the marginal cost of shipping a finished product between the countries of production and destination. The theoretical model assumes that this marginal cost is constant, and equal to average cost, as a function of the number of cars shipped. This is an empirically plausible assumption if a significant number of cars are sold in foreign markets. Increasing returns to scale arise mainly from the fixed costs involved in the shipping activity.

To estimate the average cost per car, we determined first the factors that contribute to the variation of these costs across countries and products. Of course, one key determinant is the distance that the car has to be shipped. In the EEC, the transportation of finished vehicles is usually performed by rail and truck. Truck transportation is mostly used for short distances and rail is the main means of transport from factories to distribution centers. According to industry experts, railway tariffs do not change much with distance and it is not clear that the cost function shows "increasing returns to scale" with respect to this variable. As we will see, this presumption is confirmed by our estimates.

A second determinant of costs is the weight of the product. Railway tariffs change by railcar weight brackets. Industry experts indicate a clear pattern of declining average cost per unit weight which again was confirmed by the data.

Two more determinants of cost were considered. The length of the product was highlighted as a relevant factor. After all, the cost per car will depend on how many cars are loaded in a railcar. However, mixed loads with small and large cars are common in the industry and the relationship between the length of the product and its shipping costs is unclear.

Finally, not only the distance between production and distribution centers but also the route that had to be used was deemed to be an important factor. Since the European railway sector is heavily regulated at a member state level, it was suspected that some runs might be more costly because of higher national railway fares.

To take into account possible variable returns to scale to each factor, the following functional form was specified:

$$g_{ri} = A\ (d_{ri})^{\Gamma}\ (wr)^{\beta} \qquad \text{(A.4.)}$$

Appendix 1 (continued)

where g_{ri} is the transportation costs (in ECUs) of product r from its home market to market i; Γ, β, and A are constants; dri is the distance (in km) from the home market to i; and w_r is the weight (in kg), of product r.

This function was estimated using two data sets provided by two firms in the industry. The first data set comprised only 12 observations for a representative car of 1000 kg. Thus, the weight parameter cannot be estimated with this information. The data correspond to three ports of origin and five alternative destinations. Model A.4. was estimated in loglinear form, alternatively with origin and destination effects. The presence of both effects was rejected at a 99% confidence level with F values of 1.979 for destination effects and 1.016 for origin effects.[21] The estimate of the distance parameter for the restricted model was .925 with a standard error of .098. The R^2 of the regression was .90.

The second data set contains 72 observations corresponding to four origins and nine alternative destinations. For each origin-destination pair, data on cars of two different weights is provided. The tests for origin and destination effects for this sample give opposite results. The F statistic for destination effects is 3.43 and that of the origin effects is 7.58. Both are significant statistics at a 99% confidence level and therefore we cannot reject the null hypothesis of no effects for both cases. However, for both sets of effects the adjusted R2 is only slightly reduced when the effects are removed (from .902 to .879). This suggests that, although significant, the two effects are of little importance.

The estimates for the restricted model with no effects are the following:

intercept	–.454475
distances	.901460
	(.03987)
weight	.664495
	(.2475)

with standard errors in parentheses.

These are the parameters that were used for prediction of transportation costs for the products in the price sample. Although route effects could be significant, the transport costs data did not allow the estimation of alternative intercepts for all products in the sample and, therefore, the pooled regression was used.

Computation of predicted transport costs was performed using model A.4. and the estimated parameters. When a product is manufactured in more than one plant, the locations with minimum distance to each market were considered. Plant locations were obtained from the firms involved. These data, together with the rest of information (for example, technical characteristics of the products in the price sample) will be supplied by the author on request. Distances were calculated from the plants to major distribution centers in each country. The distribution centers chosen were: Frankfurt, Paris, London, Verona, Brussels, Amsterdam, Luxembourg and Copenhagen.

[21] Of course, the reduced number of observations imply that very high F statistics are required to be able to reject the null.

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