Identifying changes in behaviour in a multiproduct oligopoly: Incumbents’ reaction to tariffs dismantling

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Abstract

The Spanish automobile market of the nineties simultaneously experienced a perfectly foreseeable tariff dismantling, complicated by a strong demand downturn, with the observed result of an apparently sharpened producer competition, both in products and prices. This paper is aimed at testing whether or not there really was a change in pricing behaviour, using a structural model of competition among oligopolistic multiproduct firms. We understand by behaviour the particular strategies, in the set of well defined, market-specific equilibrium concepts, which are sustained at a given moment. To answer that question, we specify and estimate a pricing equation with panel data for the 164 models belonging to 31 firms which competed in the market during this period. The specification includes several equilibriums as alternative (overlapping) estimating models, considering prominently tacit coalitions by which a group of firms sets prices, taking into account the cross effects on their demands. The statistical test selects as the best model given the data a switch from collusion to competition of domestic and European producers at the beginning of the nineties.

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1. Introduction

At the beginning of the nineties, the Spanish automobile market experienced a perfectly foreseeable tariff dismantling, complicated with a strong demand downturn (see Figure 1), with the observed result of an apparently sharpened producer competition, both in products and prices. Domestic producers (installed multinational firms) and foreign producers (European and non-European) changed car attributes, introduced new models, and engaged in promotional campaigns, at the same time that several signs of price competition appeared. This paper is aimed at testing whether or not there really was a change in pricing behaviour, using a structural model of oligopolistic multiproduct firms which compete in a product differentiated market.

Obviously, it must be understood that firms continuously adjusted model prices to their environment (entry and changes in characteristics). And all producers, multiproduct firms with a rough average of more than three car models on the market at a given moment, must be assumed internalising optimally the cross effects of their model pricing. The central question is whether, in addition, tariff dismantling induced a change in firms’ pricing strategies, modifying their degree of rivalry. We understand by behaviour the particular strategies, in the set of well defined market-specific equilibrium concepts, which are sustained at a given moment.

To answer that question, we specify and estimate a pricing equation with (monthly) panel data for the 164 car models belonging to the 31 firms which competed in the Spanish market during the period 1990-96. By developing the underlying theoretical model, the specification is able to include several equilibriums as alternative (overlapping) estimating models. Among these equilibriums we consider prominently tacit coalitions, by which a group of firms sets prices taking into account the cross effects on their demands. In particular, our statistical test selects as the best model a switch from collusion of domestic and European producers at the beginning of the nineties to a whole competition.

The paper also has, however, a more general interest. We try to uncover whether an economic policy change (tariffs dismantling) triggered a behavioural change. A similar is-
sue becomes relevant when market events may trigger a change in behaviour: e.g. demand downturns or mergers. In fact, quantitative methods of analysis of market competition for backing policy decisions have largely ignored the question of changes in behaviour, with and without government intervention. Some of the pioneering works in the "new empirical industrial organisation" were motivated by and focused on the analysis of behaviour changes (Porter, 1983; Bresnahan, 1987) but, in general, the question of the direct identification of behaviour has shown elusive. More generally, market modelling is often aimed at assessing market power and describe its sources, e.g. product differentiation versus price coordination (for very recent examples see Nevo, 2001, Goldberg and Verboven, 2002, and Pinkse and Slade, 2003.) But all exercises of this type reveal both that integrating behaviour in the analysis matters and that there are few instruments developed to integrate it. There have recently been many advances in the specification and estimation of demand in industries with product differentiation (mainly since Berry, 1994, and Berry, Levinsohn and Pakes, 1995 and 1999.) But the rich modelling of the demand side often comes at the cost of constraining behaviour. Here, using many of the ideas of the recent advances, we try to develop specific and simple ways to address the identification of firms' behaviour.

The methodology consists of specifying and estimating pricing equations which nest the unobservable marginal cost and the margins established by firms. Margins can be shown to be in general a function of demand price effects, firms' market shares and behaviour. By specifying alternative behaviours, one ends with a series of models which predict different margins which depend on different ways on observed shares. Equations may be developed by using methods which range from non-parametric specifications of the price effects to the use of parametric demand models. Then, the alternative models can be estimated and the model that best fits the data selected by means of the suitable test. In this paper we use price effects parametrised according to a non-symmetry differentiation logit model and use a selection test for non-nested models.

The rest of this paper is organised as follows. Section 2 explains in detail the competition changes that took place in the Spanish market and descriptively explores the price data. Section 3 discusses the way to specify and test for behaviour. Section 4 is devoted to detail
the specification and estimation techniques that we apply to the pricing equations, and
Section 5 to explain the empirical results. Section 6 concludes. An Appendix develops the
logit model which the paper employs to specify the price effects.

2. Competition changes

At the beginning of the nineties, the Spanish automobile market was served by three
types of car producers: firstly, domestic producers, or multinationals with plants installed
in Spain during the seventies and the eighties aimed at exporting an important part of
production, manufacturing in them some of the car models they sold; secondly, European
foreign producers, or multinational European producers without manufacturing in Spanish
territory; and thirdly, non-European foreign producers, basically Asian, sometimes possess-
ing an incipient production in European territory.

Domestic producers accounted for seven brands belonging to five groups (Citroen-Peugeot,
Ford, Opel, Renault and Seat-VW), which coincided with the most important non-Japanese
world producers with the absence of Fiat and Chrysler (recall that Opel is a GM subsidiary).
They had dominated the Spanish market during the eighties, and they started the nineties
with a joint market share of 82%. At this time the European foreign producers’ supply
consisted of 14 brands\textsuperscript{1}, with a joint share of only 16%, but they sold more than half of the
cars of the highest segment. And non-European producers accounted for 9 Asian brands\textsuperscript{2}
plus Chrysler, which entered the market in 1992, representing all together a market share
of 2% (see Table 1.)

Tariff and non-tariff protection made it unprofitable to import cars from abroad during
the beginning of the eighties, dampening even the import of the models from domestic
producers not produced in Spain. All imported cars in 1985 amounted to only 13% of sales.
But this year the Spanish Adhesion Treaty to the EEC, setting the transition framework
to full integration in the single market of 1992, firmly established a different perspective.

\textsuperscript{1} Audi, Alfa-Romeo, BMW, Fiat, Jaguar, Lada, Lancia, Mercedes, Porshe, Rover, Saab, Skoda, Volvo
and Yugo.

\textsuperscript{2} Daewo, Honda, Hyundai, Mazda, Mitsubishi, Nissan, Subaru, Suzuki and Toyota.
Tariffs on cars imported from the EEC had to be decreased as stipulated from the then-current value of 36.7% to zero by the beginning of 1993. And tariffs on cars imported from third countries had to be reduced from the then-current value of 48.9% to the common EEC tariff of 10%, although import quotas applied to Asian cars remained the same at the time.

This perspective immediately started a new competition preparing the coming open market, stimulated by a very dynamic demand. Domestic producers enlarged the range of models distributed in the market with models imported from their production in plants abroad, while foreign producers entered new models. Imports had risen to virtually 40% of sales by 1990 (recall that only 18% are imports by foreign producers). At the same time, domestic producers invested heavily in plant productivity improvements. But, the beginning of the nineties, in which demand transitorily experienced a stagnation and then a sharp downturn, triggered a new competition intensity.

Competition during the nineties adopted several dimensions: producers engaged in a sharp increase in advertising expenditures and an enlargement of sales networks, continuously redesigned existing models introducing quality and more equipment, continued the entry of new models, and started an unprecedented price competition which consumers perceived through promotional advertising. The entry of car models, performed with the double target of replacing old models and introducing in the Spanish market models absent until this time, was particularly important and can be followed in detail by means of our data. The year 1990 began with 79 marketed models, while in the following years 104 models entered the market and 59 exited, which implies 123 marketed models by the end of 1996. Asian cars accounted for a disproportionate share of this entry (28 entering models vs. only 15 in 1990), but entry by the European foreign producers (48 entering models vs. 48 in 1990) and domestic producers (28 entering models vs. 35 in 1990) is also important.

The role of replacement can be seen by noting that 90% of exits are separated from a model entry by the same brand by less than 48 months.

Among all these competition changes, the focus of this paper is on pricing. In particular, did the dismantling of tariffs change firms' price behaviour? Foreign firms found themselves able to sell at significantly lower prices for the same received prices. Domestic producers
experienced the same change for the models which were introduced from abroad and, at the same time, they expected increased competition for all models which included entry and lower rivals’ prices. Obviously, it is understood that firms must continuously adjust each model price to its new environment (entry and changes in characteristics). And all producers are multiproduct firms, with a rough average of more than three car models on the market at a given moment, which implies that they must be assumed to optimally internalise the cross effects of their models pricing. The central question is whether, in addition, tariff dismantling induced any change in firms’ pricing strategies, modifying their degree of rivalry. We understand by behaviour the particular strategies, in the set of well-defined, market-specific equilibrium concepts, which are sustained at a given moment. Testing formally for these changes is the aim of this paper. In what follows, we descriptively explore the likelihood of these changes.

To acquire an impression of possible pricing behaviour changes in our sample period, the cost changes induced by quality changes must be disentangled. With this aim, we will use the hedonic coefficients resulting from regressing prices on car characteristics. Let us define the price corrected by quality changes as

\[ \tilde{p}_{jt} = p_{jt} - (x_{jt} - x_{j0})\hat{\beta} \]

where \( x_{jt} \) is the vector of characteristics of model \( j \) at moment \( t \), \( x_{j0} \) stands for this vector when the model enters the sample, and \( \hat{\beta} \) represents the cost per unit of characteristic estimated in the hedonic regression.\(^3\) Averages of these quality-corrected prices will change with the entry and exit of models, which embody idiosyncratic qualities that shift the mean. To avoid these effects, let us define quality change and entry-corrected prices as

\[ \tilde{p}_t = \frac{1}{N} \sum_j (\tilde{p}_{jt} - (x_{j0} - \bar{x})\hat{\beta}) \]

where \( \bar{x} \) is the sample mean of attributes and \( N \) is the number of models at date \( t \). Entry

\(^3\)We employ the coefficients corresponding to our preferred model (see Section 5), but the exercise produces very similar results using alternative estimates.
and quality change-corrected prices, depicted as indices, give the change in prices which may be attributed to reasons other than quality-induced cost variations. Of course, they can show cost changes attributable to other reasons, but they are likely to clearly reflect changes in pricing.

Figures 2, 3 and 4 summarise the results of descriptively exploring price changes. Figure 2 represents simple average monthly prices for the three producer types, deflated by the consumer price index, and the average received prices; that is, the price received by producers after deducting the relevant tariffs. The figure highlights an apparent parallel evolution of European and domestic received prices during the period, at a different level determined by the diverse sales composition, and a sharp decrease of the Asian received prices. Figure 3 represents the evolution of received prices differing out the quality-induced cost variations (normalised to unity the first year), and Figure 4 represents the evolution differing out the quality composition effects of entry and exit. Both figures stress the same conclusions. Firstly, all prices tend to show a fall during the first three years (1990-92) and some recovery at some point of the following subperiod. This suggests partly procyclical pricing, matching the demand evolution reported above. Secondly, Asian car prices show a sharp new fall by the year 1993. Asian producers price aggressively when the transitory tariff period reaches its end. Thirdly, domestic producers' pricing seems to recover less steadily than the European producers' pricing, denoting its likely engagement in downward price competition at least during 1993. Besides the figures, the hedonic corrections denote that quality increments of marketed cars are introduced at a similar pace for all producers, particularly after 1992, and that Asian entry mainly consists of models of lower quality as time goes by.

3. Identifying behaviour.

This section presents the framework to specify and test for behaviour by means of price equations. Firms are assumed to be multiproduct, competing in a product-differentiated industry given products and their characteristics. We understand by behaviour the partic-
ular strategies, in the set of well-defined equilibrium concepts, sustained by firms. A wide range of different behaviours may be covered, although we confine ourselves to price games equilibriums. One of the lessons of this section is that the testing for price behaviour can be based on price equations alone. Efficiency can, of course, be improved with demand information, but the relevant information is implicitly estimated in these equations. This is because firms’ imperfectly competitive pricing crucially employs this information, while the converse is not true: consumer behaviour is generally price taking. Our equations are structural, stemming from the equilibrium relationships which emerge between prices and output, represented by shares, in a broad class of price games. Firms’ shares are hence endogenous variables and must be treated as such.

We firstly set out the basic framework and show that non-parametrical specifications of behaviour amount to include the own and the behaviour-specific relevant rivals’ shares as right-hand variables of the pricing equations. In this context, testing behaviour can be done by means of testing exclusion restrictions. But we also show that non-parametric specifications impose strong data requirements when the number of products is not very small. Then we show that the use of demand models for the price effects, particularly logit models, dramatically increases feasibility and efficiency. But it is also shown that models then become non-nested and testing behaviour needs the use of non-nested selection tests.

3.1 Basic framework.

Let us assume a product-differentiated industry consisting of $F$ multiproduct firms, indexed $f = 1, ..., F$. Each firm produces $J_f$ products and there are in total $J = \sum_f J_f$ products. We will often generically refer to a product $j$, which is implicitly assumed to be one of the products of the set $j = 1, ..., J_f$ produced by firm $f$. Demand for each product $j$ is a function of the $J \times 1$ vector of prices $p^t$. Demand for each product $j$ is a function of the $J \times 1$ vector of prices $p^t$. Write demand, for convenience, in the shares form $q_j = s_j(p)M$, where $M$ is market size\(^5\) (usually the number of potential consumers.)

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\(^4\)In addition, demands will have, in general, other arguments (e.g. product characteristics) which we drop here for simplicity.

\(^5\)The model can be equally applied to demands in the form $q_j = q_j(p)$. 

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Share \( s_0(p) \) stands for the fraction of consumers buying nothing. Let \( s \) be the \( J \times 1 \) vector of shares and define the \( J \times J \) price effects matrix \( D = \left( \frac{\partial s_k}{\partial p_j} \right) \), where row \( j \) collects the own and cross-demand effects of price \( j \). Product \( j \) constant marginal cost\(^6\) is \( c_j \).

With price-strategic complements (reaction curves are upward-sloping), a price increase on a product generates a positive externality on the other product profits, including rival firms’ product profits. Assume that firms care about internalising the cross-price effects of their own profits. That is, firm \( f \) sets prices, even in the most competitive setting, by maximising \( \sum_{k \in J_f} (p_k - c_k)q_k \). But firms can also set prices which internalise the positive cross-price effects among a group of rivals. That is, they can form price coalitions (Deneckere and Davidson, 1985), by maximising \( \sum_{k \in J_h} (p_k - c_k)q_k \), where summation is extended to the \( J_h = \sum_{f \in h} J_f \) products of the firms which take part in the coalition.\(^8\) Let \( H \) be the number of coalitions. Belonging to a coalition implies setting prices to maximise over a set of products which contain the own products as a subset. Hence, from now on, we will speak exclusively about the grouping of products at the level of coalition without loss of generality (firms can simply be thought of as one-member coalitions).

Prices result from maximising behaviour given the other prices, that is \( p_j = \arg \max \{ \sum_k \delta_{jk}(p_k - c_k)q_k \delta_j \} \), where \( \delta_j \) is a \( 1 \times J \) vector of ones and zeroes, with element \( \delta_{jk} \) being the indicator of inclusion of product \( k \) in the relevant profits sum (i.e., \( \delta_{jk} = 1 \) if \( k \in J_h \) and \( \delta_{jk} = 0 \) otherwise.) Let us stack all vectors in a \( J \times J \) matrix \( \delta \). The set of FOC conditions which define a price equilibrium can be written in matrix form as

\[
 s + (\delta \circ D)(p - c) = 0
\]

where \( \circ \) represents the Hadamard product (element by element product). Equilibrium prices are easily obtained as

\(^6\)Marginal cost is assumed here to be known to simplify notation.

\(^7\)Non-constant marginal costs may be considered by specifying relevant marginal cost as \( c_j(1 + k) \), where \( k \) is the elasticity of cost with respect to output.

\(^8\)We assume that firms enter price coalitions with all their products. Other equilibriums can be considered, but notably complicate the analysis and notation.
This system, consisting of equations in the form $p_j = c_j + m_j(D, s, \delta)$, shows that margins corresponding to a particular equilibrium are a function of demand price effects, firm shares, and behaviour as represented by $\delta$. These equations are structural equilibrium relationships, relating the endogenous variables $p$ and $s$, with $s$ determined additionally through the system of demands.

**Proposition.** With firms setting prices, product $j$ margin can be written as a linear combination of the shares of the products included in the coalition, with weights which are a function of the coalition submatrix of demand price effects.

Proof. Let $P_H$ be the permutation matrix which induces a re-ordering of firms (products) according to the coalition they belong to. $P_H s + P_H(\delta \circ D)P_H^tP_H(p - c) = 0$ is a system equivalent to (3), and hence $P_Hp = P_Hc - (P_H(\delta \circ D)P_H^t)^{-1}P_Hs$ gives the same prices. But, by definition of $\delta$, $(P_H(\delta \circ D)P_H^t)^{-1}$ is a block diagonal matrix of elements $D_h^{-1}$, $h = 1, \ldots, H$. □

3.2 The specification of behaviour.

An important implication of the above proposition is that, if price effects can be considered given, and hence estimable, all non-parametric econometric specifications of behaviour are in principle a question of adding enough rivals shares among the right-hand side variables. Consistent estimation will imply the use of IV, because shares are endogenous, but tests between equilibriums may be easily carried out as tests of exclusion restrictions. Efficiency could in principle be improved by simultaneously estimating the price effects with demand data, but the highly non-linear form in which these effects enter the price equations makes this approach extremely impractical. Given the stated price equation properties, a completely non-parametric specification is a very attractive alternative. However, this approach to estimation imposes strong data requirements and the use of parametric demand models improves the situation. In what follows, we explain the data requirements and
Non-parametric specification. Table 2 summarises the price effects and price equations corresponding to several key specification choices. The estimating equation corresponding to a fully non-parametric alternative (see the left-top panel of Table 2) can be written

\[ p_j = c_j + \beta_{jj} s_j + \sum_{k \neq j, k \in J_h} \beta_{jk} s_k + v_j, \quad j \in J_h \]  

where \( v \) represents a disturbance term with \( E(v|z) = 0 \) for a suitable set of instruments \( z \).

That is, the \( J_h \) product shares in the coalition enter the \( J_h \) equations for \( j \) \((j = 1, \ldots, J_h)\) with specific coefficients. But equilibriums described by equation (5) imply the estimation of \( \sum_h J_h^2 \leq J^2 \) parameters, a number which increases with the number of products involved in big coalitions. Hence, equations like (5) can only be estimated for a very small number of products and enough repeated observations for each product \( j \), either over time or across markets.

An obvious alternative is the assumption of similar price effects for groups of analogous products or “nests” (see the bottom-left panel of Table 2). Let us suppose \( G \) nests, indexed \( g = 1, \ldots, G \), with \( N_g \) products each. The estimating equation becomes

\[ p_j = c_j + \beta_{gh} s_{gh} + \sum_{m \neq g} \beta_{mh} s_{mh} + v_j, \quad j \in J_h, j \in N_g \]  

where \( s_{mh} = \sum_{k \in J_h, k \in N_m} s_k \) stands for the share of products of coalition \( h \) in nest \( m \). The number of parameters to be estimated is now \( \sum_h G_h^2 \leq G^2 H \), which will generally imply a dramatic fall. In addition, a further reduction of the number of parameters to a maximum of \( \frac{G(G+1)}{2} H \) can be obtained by imposing (when reasonable) symmetry of the price effects. However, even this number may be high for many sample sizes (with 5 nests and 2 coalitions, the number of parameters to be estimated amounts to 30.) If coalitions are not very disimilar in the number of products in each nest, the idiosyncratic parameters may be approximated by common estimates, i.e., \( \beta_{gh} \approx \beta_{gm} \), with the result of only \( G^2 \) or \( \frac{G(G+1)}{2} \) parameters to be estimated. This is better, but can still be too large a number.
for the size of most samples.

**Demand model for the price effects.** Demand models are, in this context, a potentially powerful tool for dealing with the dimensionality of the price effects. Theoretically-founded demand models, by specifying price effects as a function of a few parameters and observable variables, can drastically reduce the number of parameters to be estimated, increasing feasibility and efficiency. Logit demand models, for example, have recently been profusely used in estimating elasticities and other characteristics of demand. And a number of researchers have even used them to specify price equations, although mostly under tight constraints on behaviour (references...). In what follows, we show that logit-discrete choice models of market demand are particularly useful here.

One potential problem is that the simplest logit model embodies unrealistic constraints on the price effects, up to the point that a complete line of research has been devoted to develop richer specifications of consumer heterogeneity as the random coefficients model (Berry, Levinshon and Pakes, 1995). However, it can be shown that, at least for the purpose of modelling firm behaviour, a simple generalisation of the basic model helps to attain reasonable flexibility. Here we base the specifications on a varying marginal utility version of the basic logit model, which is developed in the Appendix. Its main characteristics are the following: a) differentiation is not symmetric for all products, combining both horizontal (symmetric) and vertical (ordered) dimensions; b) marginal utility of income is allowed to vary, and different average income utilities for buyers of different products are a key characteristic for producing sensible elasticities; c) the model is easily invertible, facilitating the specification of behaviour.

The estimating price equation corresponding to the price effects of the non-nested version of the logit model (see the top-right panel of Table 2) has the form

\[
p_j = c_j + \beta_j \theta_j(s_h) + \sum_{k \neq j, k \in J_h} \beta_k \theta_k(s_h) + v_j, \quad j \in J_h
\]

where \(s_h\) stands for the vector of product shares of the coalition to which product \(j\) belongs. Regressors are now product-specific known functions of this vector and only enter the equa-
tion if the product also belongs to the coalition. Coefficients are only regressor-specific (they do not depend on \( j \)) and hence the number of parameters to be estimated is \( J \). Estimation and testing of this equation provides, in addition to inference on behaviour, estimates of key demand parameters of the logit model (the product-specific \( \alpha' s \); see Table 2). This is, hence, an outstanding alternative when the number of products, \( J \), is not too high.

The problems associated with the presence of two (or more) levels of (horizontal) substitutability of the products can be remedied by the nested version of the logit\(^9\). The criticism usually made against this model is the exogenous choice of the grouping of products, but in many cases the own market data clarify the relevant clusters of products, which in any case can be tested if necessary. With nested logit price effects, the estimating equation becomes (see the bottom-right panel of Table 2)

\[
p_j = c_j + \beta_g \theta_g(\sigma, s_h) + \sum_{m \neq g} \beta_m \theta_m(\sigma, s_h) + v_j, \quad j \in J_h, j \in N_g
\]

where regressors are now nest-specific known functions of the vector of shares of the coalition to which product \( j \) belongs (and of parameter \( \sigma \)), and only enter the equation when the coalition has a non-zero share in the nest. Coefficients are only nest-specific (they do not depend on \( j \)) and hence the number of parameters to be estimated is simply the number of nests \( G \). Estimation and testing of this equation also provide estimates of key demand parameters of the nested logit model (the nest-specific \( \alpha' s \); see Table 2). Parameter \( \sigma \), which expresses the degree of substitutability inside the nests, must, however, be given a previous value.

The logit models used here are suitable for situations in which product differentiation is not symmetric, i.e., there is some vertical differentiation dimension which induces an ordering of the products. With fully-symmetric products, the formulae of Table 2 simplify dramatically, and specifications include a unique parameter to be estimated. However,

\(^9\)The nested logit mitigates the effects of the property of independence of the irrelevant alternatives (IIA) by which the ratio of any two logit probabilities of choice remains invariant to the introduction of new alternatives. This is not a desirable property if the new alternative is a close substitute for one type of the existing alternatives.
this is hardly a realistic situation for many markets. Other possible ways to specify price effects is to make them a deterministic function of market structure or a function of product characteristics. Both approaches have been successfully applied to model demand and price equations (see Jaumandreu and Lorences, respectively, 2002, for the first approach, and the “distance-metric” approach of Pinkse, Slade and Brett, 2003).

3.3 Testing behaviour.

Logit-based price equations provide a nice alternative to reducing the number of parameters to be estimated on theoretical grounds. But they imply an important difference in the logic of the specification and testing with respect to the non-parametric specifications. In the non-parametric setting, coefficients are non-linear functions of the underlying price effects, but shares enter linearly. Behaviours can be tested among them by means of exclusion restrictions. In the logit-based equations, coefficients directly estimate parameters of theoretical interest, but regressors are now behaviour-specific, non-linear functions of the observed shares. That is, alternative behaviours imply non-nested, partially-overlapped models.

We are going to use the nested logit-based specification, and hence we need a test for selection among non-nested models. To statistically compare the results, we then use a Vuong-type test of selection among non-nested or overlapping models (see, for similar applications, Gasmi, Laffont and Vuong, 1992, and Jaumandreu and Lorences, 2002.) We compute it with the GMM analogous to the likelihood ratio. That is, for every two models, we compute the value

$$ V = \frac{N(J_2 - J_1)}{\sum (J_{j1} - J_{j2})^2 - N(J_1 - J_2)^2}^{1/2} $$

where $J_1$ and $J_2$ are the corresponding minimised values of the objective function, $J_{j1}$ and $J_{j2}$ are the individual observation values of the objective function evaluated at the minimum, and $N$ the total number of observations. We expect this statistic to be distributed as a $N(0, 1)$.
4. Econometric specification.

In this section we discuss the specification and estimation techniques to be employed in the empirical exercise. We are going to estimate price equations using data on the car models sold on the Spanish market from 1990 to 1996 by the 31 firms with a presence in the marketplace. The data consist of unbalanced panel observations for a rather standard number of individuals (182 models) but with the more unusual characteristic of monthly data frequency (a maximum of 84 observations per individual). Using the price equilibrium relationships established in Section 3, the final objective of this empirical exercise is to obtain estimates under the assumption of different behaviours and to test their relative likelihood given the market data.

We must carry out the simultaneous estimation of a nested marginal cost function and the firms’ markups. Section 3 has shown us that equilibrium prices are additively separable in two components, marginal costs and unit markups. To estimate marginal costs, we adopt the “hedonic” approach to the estimation of the cost function\(^\text{10}\): we will take cost as a function of a set of product attributes.

Using the nested logit specification of the price effects, our estimating equation then has the form

\[
p_{jt}/(1 + tariff_{jt}) = \alpha + \bar{x}_{jt}\beta + \sum_g \beta_g \theta(\sigma, s_{nt}) + \eta_j + u_{jt}
\]  

where the dependent variable is received prices; that is, model monthly prices deflated, when relevant, by the tariff. Variables \(\bar{x}\) represent the attributes (in the form of deviation with respect to the sample mean and squares of these deviations, to approximate cost around its mean \(\alpha\)); the sum of terms in \(\theta\) stands for the specification of behaviour according to Section 3; the term \(\eta\) is an unobservable individual error, representing cost unobservable

\(^{10}\text{Cost estimations starting from regressions on product characteristics can be called “hedonic” because they use the methodology of the traditional hedonic price regressions (see Griliches (1961), and also Pakes (1997)). Hedonic price equations relate price to product characteristics. They can be theoretically based on equilibria under the approach of consumer preferences for bundles of attributes and the willingness to pay for these characteristics (Rosen (1974)).}
advantages or disadvantages, and \( u \) is a zero mean disturbance given the set of instruments \( z \). In practice, to control for seasonality, we also include a set of monthly dummies.

The sum of regressors \( \theta \) is close to unity by construction, given their theoretical specification. This implies a serious difficulty in separately identifying the average cost and the average level of margins, at least without more cost data and estimating the pricing equation in isolation. Some cost data allowing for identification of mean cost, and/or joint estimation with a demand equation, imposing cross constraints on the coefficients, would probably improve the identification conditions. But, for the moment, we will limit ourselves to identifying relative margin differences by constraining the behavioural coefficients to add up to zero. This does not, in any case, affect the model capacity of discrimination among conducts.

The optimal pricing expressions developed in Section 3 starting from first order conditions are equilibrium relationships; that is, they relate the endogenous variables prices and shares (or functions of shares) in equations that also include products’ marginal cost and parameters of demand. They define only implicitly the “reduced-form” equations for equilibrium prices, in which prices would depend only on marginal costs, demand parameters, and product attributes appearing in the shares (demand) equations in addition to prices. Given the highly non-linear form in which shares depend on prices, explicit reduced-form equations cannot be obtained.

In this context, estimation of markups raises an endogeneity problem. Shares and functions of shares are likely to be correlated with \( \eta \) and \( u \). It is natural to assume that prices are explained by the observed attributes that determine costs, some additional unobserved cost components summarised in the individual \( \eta \)'s and the shocks \( u \), and markups which depend on shares. But market shares are a function of prices. Hence, shares will be correlated with the cost side unobservable terms \( \eta \) and \( u \). On the demand side, shares can also be influenced by unobserved components of utility that are likely to be correlated with the cost unobserved components.

We have an estimating expression that is linear in parameters but includes a set of nonlinear endogenous variables. This implies the use of instrumental variables (IV) methods.
of estimation. We are going to employ GMM methods suitable to unbalanced panel data. The GMM estimator exploits the moments $E(Z_j' \xi_j) = 0$, where $Z_j$ stands for the matrix of instruments and $\xi_j$ represents the vector of elements $\xi_j = \eta_j + u_{jt}$. We employ the estimator $\hat{\beta} = (M_{xx}'AM_{xx})^{-1}M_{xx}'AM_{xx}$, where $M_{xx} = \sum_j Z_j'X_j$, with the consistent “one-step” choice $A = (\sum_j Z_j'Z_j)^{-1}$.

The most standard way to treat the setting explained above is to estimate the equation taking differences in order to difference out the individual component of disturbances, and to use lags of the endogenous variable in order to set valid moment restrictions (see, for example, Arellano and Honoré (2001)). In our case, this is an undesirable alternative because the time dimension of the data is short in relation to the pace of variation of attributes. Monthly data are likely to contain useful information about prices and shares, which change frequently, but much less about reactions to attributes, which change less frequently and whose change has mainly longer-term effects. The differentiation of the attributes would eliminate crucial information contained in the levels equation and would exacerbate the variance of the disturbances. Instead, we will use as instruments the differences of the shares with respect to their individual time means, $\tilde{s}_{jt} = s_{jt} - (1/T_j) \sum_l s_{jl}$, lagged a number of periods. These instruments share with their past time variations, avoiding the use of their level variations across models. Instruments of this type were first proposed by Bhargava and Sargan (1983), and moment restrictions of this type have been studied in Arellano and Bover (1995). A recent discussion may be found in Blundell and Bond (1998). As additional instruments, we will employ the set of 31 brand dummies. To test the validity of the employed instruments, we employ the Sargan test statistic of the overidentifying restrictions.

To obtain inferences robust to serial correlation, we will use a robust estimate of the variance-covariance matrix (see Newey and West, 1987). All the statistics are then computed using this robust-to-heteroskedasticity and serial autocorrelation “two-step” weighting matrix. To estimate a robust inverse of $E(Z_j'\xi_j\xi'_jZ_j)$, we assume that $\xi_j\xi'_j = J_j$ are matrices corresponding to conditional homoskedastic errors, and we obtain $\tilde{J}_j$ values using the Newey-West Bartlett-kernel computations for the autocovariances of individual $j$. Then
we employ the usual “two-step” estimate $A_R = (\sum_j Z_j^T \tilde{Z}_j Z_j)^{-1}$. We use 72 time observations as the maximum lag that we take into account in the Bartlett kernel specification.

5. Empirical results.

Estimation of the pricing equation is carried out employing as attributes the power measure ratio cubic centimetres to weight (CC/Weight), the fuel efficiency ratio km to litre (Km/l), used in the particular form of the relative efficiency in city driving with respect to 90 Km/h driving, the measure of size and safety length times width (Size), the maximum speed in km/h (Maxspeed) and air conditioning (ac) as “luxury” indicators, and the materials use indicator weight (Weight). We try to be deliberately close to Berry, Levinsohn and Pakes’ (1995) specification for the sake of comparisons. The use of other characteristics or a more complete list does not change the main results.

We group models into 5 categories that closely follow common industry and marketing classifications. The nests of cars considered are: small, compact, intermediate, and luxury. We separately group minivans, which were at that moment a product beginning their market penetration. The number of models in each segment are, respectively, 33, 37, 56, 47 and 9. Variables of behaviour are computed using the nested logit formulas of Section 3 (see Table 2) using a $\sigma$ value of 0.8. This is a standard value and also the value which is obtained in an independent demand estimation with the same data.

Table 3 presents the results of estimating some key models. All attribute coefficients show the sign expected in a cost function and sensible values (cost increases in weight in all the sample values). Moreover, they do not change dramatically from estimate to estimate, although some changes are significant.

Estimates of Table 3 implement the three more straightforward specifications of behaviour. The first equation assumes that behaviour was Bertrand-Nash all the time and for all players, a common assumption in many models and estimates of this type. The second equation makes the unrealistic assumption that behaviour was collusion of all players all the time. The third estimate makes the sensible assumption that domestic and European
producers set prices at the beginning of the period, internalising the cross effects of their prices; i.e., they constituted a price coalition, but this coalition broke up at the end of 1991. Domestic and European producers are then supposed to switch to play Bertrand, while Asian producers are assumed to play Bertrand the entire period. The estimate of a model in which the breaking up of the coalition is assumed at the end of 1992 gives very similar results.

This third estimate is the best in economic and statistical terms. On the one hand, the coefficients of the variables’ modelling behaviour exactly follow the pattern expected according the theoretical specification, and they are roughly consistent with the coefficients and price elasticities of demand which have been obtained in a demand equation estimated separately. On the other, the Vuong-type test of selection among non-nested or overlapping models (see Section 3) selects it as the best among the three models.

Given the central role of the year 1992, and to check that the conclusion is also robust to closer, less differentiated alternatives of behaviour, two additional (a-priori), more or less likely sequences of behaviour are estimated. One specifies that domestic producers left the domestic-European coalition at the end of 1991, i.e., began to play Bertrand, but that European producers continued to coordinate pricing. The other is that European producers were never involved in price coordination: there was a coalition exclusively formed by domestic producers which broke up at the end of 1991.

Table 4 identifies all the estimated behaviour sequences by numbers and reports the results of the Vuong-type test. From inspection of the matrix of values, it becomes clear that model 3, the coalition of domestic and European producers which broke up at the end of 1991, is a model which clearly statistically dominates the others, i.e., best fits the data.

6. Conclusion.

This paper has addressed the question of whether the Spanish car market underwent a change in pricing behaviour that coincided with the tariffs dismantling attained by 1992. The answer is yes, that it was tacit coordination in pricing maintained up to this moment
by domestic and European producers, and that this coordination seems to have broken up by this time. The result is, of course, conditional on the modelling. This is simply the model which best explains the data among the behaviour sequences proposed to match the data. And the model also does not inform us about the relative role played in this change of behaviour by the tariffs dismantling and, for example, the demand downturn. But the answer is by no means useless because it provides, at least, a first step for focussing on narrower hypotheses and more complex structural models. In addition, the estimates carried out permit us to assess the weight of price coordination in the pre-change prices and the welfare effects of the change.

More generally, the exercise carried out shows that the estimation and test of suitably-specified price equations can be the basis for identifying behaviour and behaviour changes. In addition, these estimates provide a method for assessing the sources and the effects of market power, and even estimates of some demand parameters. The methods proposed here can be applied at least to any situation in which firms of a differentiated product market are likely to price (or are going to price) coordinately, and hence they are useful and reasonably simple tools to be added to the set of quantitative methods for backing decisions of competition policy.
References


University of Chicago, mimeo.

Pinkse and Slade (2003)

Pinkse, Slade and Brett (2003)


Table 1
The Spanish car market in the 90s

<table>
<thead>
<tr>
<th>Year</th>
<th>Sales index</th>
<th>Sales entry</th>
<th>Models entry</th>
<th>Models exit</th>
<th>No. of models</th>
<th>Share of domestic producers</th>
<th>Share of European producers</th>
<th>Share of Asian producers</th>
</tr>
</thead>
<tbody>
<tr>
<td>1990</td>
<td>971,466</td>
<td>100.0</td>
<td>19</td>
<td>2</td>
<td>98</td>
<td>82.0</td>
<td>16.0</td>
<td>2.0</td>
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<tr>
<td>1991</td>
<td>878,594</td>
<td>90.4</td>
<td>10</td>
<td>5</td>
<td>106</td>
<td>80.0</td>
<td>16.9</td>
<td>3.1</td>
</tr>
<tr>
<td>1992</td>
<td>973,414</td>
<td>100.2</td>
<td>16</td>
<td>10</td>
<td>117</td>
<td>81.3</td>
<td>14.6</td>
<td>3.9</td>
</tr>
<tr>
<td>1993</td>
<td>735,993</td>
<td>75.8</td>
<td>13</td>
<td>9</td>
<td>120</td>
<td>80.7</td>
<td>13.9</td>
<td>5.2</td>
</tr>
<tr>
<td>1994</td>
<td>897,492</td>
<td>92.4</td>
<td>13</td>
<td>14</td>
<td>124</td>
<td>78.6</td>
<td>15.8</td>
<td>5.3</td>
</tr>
<tr>
<td>1995</td>
<td>822,593</td>
<td>84.7</td>
<td>17</td>
<td>10</td>
<td>127</td>
<td>77.0</td>
<td>15.7</td>
<td>6.8</td>
</tr>
<tr>
<td>1996</td>
<td>897,906</td>
<td>92.4</td>
<td>16</td>
<td>9</td>
<td>133</td>
<td>75.0</td>
<td>15.8</td>
<td>8.4</td>
</tr>
</tbody>
</table>
| Table 1  
Price effects and equations for some specifying choices |
<table>
<thead>
<tr>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Non-parametric specifications</strong></td>
<td><strong>Theoretical parameters:</strong></td>
</tr>
<tr>
<td>Logit models with non-symmetric differentiation</td>
<td></td>
</tr>
<tr>
<td>$D = -A\Pi + A\pi'$, $A = \begin{bmatrix} \alpha_1 &amp; \cdots &amp; \alpha_J \end{bmatrix}$, $P = \begin{bmatrix} P_1 &amp; \cdots &amp; P_J \end{bmatrix}$ and $\pi = \begin{bmatrix} \pi_1 &amp; \cdots &amp; \pi_J \end{bmatrix}$</td>
<td></td>
</tr>
</tbody>
</table>

**Non-nested**

$$p_j = c_j + b_{j\pi} s_j + \sum_{k \neq j, k \in J_h} b_{jk}s_k, \quad j \in J_h$$

where $\{b_{jk}\} = D^{-1}$

**Nested**

$$D = -D_G + E D_{GG} E', \quad D = -\sigma A\Pi + A\Pi \Sigma E' \Pi, \quad \Sigma = ee^T$$

where $D_G = \{d_g + d_{gg}\}$, $D_{GG} = \{d_{gm}\}$ is $G \times G$, $E = \begin{bmatrix} e_1 & \cdots & e_G \end{bmatrix}$ is $G \times G$ with $\Pi_G = \begin{bmatrix} P_1 & \cdots & P_G \end{bmatrix}$

$$p_j = c_j + \frac{1}{\alpha_g} w_{gh} (1 + \frac{w_{gh}}{1 - s_{gh}})s_{gh} + \sum_{m \neq g} \frac{w_{gh} w_{mh}}{\alpha_m} s_{mh}, \quad j \in J_h, j \in N_g$$

where $\{b_{gmk}\} = (D_{GG}^{-1} - E' D_G E)^{-1}$

$$p_j = c_j + \frac{1}{\alpha_g} w_{gh} (1 + \frac{w_{gh}}{1 - s_{gh}})s_{gh} + \sum_{m \neq g} \frac{w_{gh} w_{mh}}{\alpha_m} s_{mh}, \quad j \in J_h, j \in N_g$$

where $w_{gh} = (1 + \frac{\sigma}{1 - \sigma} (1 - \frac{s_{gh}}{s_g}))^{-1}$ and $w_h = \sum_g w_{gh} s_{gh}$

---

1 Logit matrices $D$ written in terms of theoretical probabilities, price equations in terms of observed shares.

2 $J$ products, $J_h$ in each coalition.

3 $G$ nests, $N_g$ products in each nest. Products are sorted by nests within each coalition. Matrices $D$ are coalition-specific because of different number of products by nest, but coalition subindices are dropped for simplicity.
Table 3
Results from model estimation and testing

Dependent variable: \( p_{jt}/(1 + tariffs_{jt}) \)
Sample period\(^1\): I – 1991 to XII – 1996; Observations\(^1\): 7,122; N\(^o\) of models\(^1\): 164
Estimation method: GMM\(^2\)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>t-ratio(^3)</th>
<th>Coefficient</th>
<th>t-ratio(^3)</th>
<th>Coefficient</th>
<th>t-ratio(^3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>2.151</td>
<td>13.74</td>
<td>2.036</td>
<td>11.69</td>
<td>1.985</td>
<td>11.33</td>
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<tr>
<td>Attributes:</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>CC/Weight</td>
<td>2.034</td>
<td>7.85</td>
<td>2.068</td>
<td>8.06</td>
<td>1.859</td>
<td>6.89</td>
</tr>
<tr>
<td>Maxspeed</td>
<td>1.144</td>
<td>2.26</td>
<td>1.043</td>
<td>2.05</td>
<td>0.832</td>
<td>1.75</td>
</tr>
<tr>
<td>Km/l</td>
<td>0.360</td>
<td>0.76</td>
<td>0.364</td>
<td>0.76</td>
<td>1.031</td>
<td>2.17</td>
</tr>
<tr>
<td>Size</td>
<td>-0.146</td>
<td>-0.75</td>
<td>-0.162</td>
<td>-0.84</td>
<td>-0.381</td>
<td>-2.05</td>
</tr>
<tr>
<td>Weight</td>
<td>3.610</td>
<td>5.41</td>
<td>3.814</td>
<td>5.54</td>
<td>3.908</td>
<td>6.04</td>
</tr>
<tr>
<td>Air</td>
<td>-0.085</td>
<td>-0.77</td>
<td>-0.091</td>
<td>-0.80</td>
<td>0.037</td>
<td>0.31</td>
</tr>
<tr>
<td>(CC/Weight)(^2)</td>
<td>1.570</td>
<td>3.90</td>
<td>1.560</td>
<td>3.90</td>
<td>1.440</td>
<td>2.70</td>
</tr>
<tr>
<td>(Maxspeed)(^2)</td>
<td>3.220</td>
<td>4.21</td>
<td>3.193</td>
<td>4.25</td>
<td>3.589</td>
<td>4.11</td>
</tr>
<tr>
<td>(Size)(^2)</td>
<td>0.100</td>
<td>1.82</td>
<td>0.099</td>
<td>1.87</td>
<td>0.088</td>
<td>1.69</td>
</tr>
</tbody>
</table>

Behaviour\(^4\)

Always Bertrand
- Small: -0.16, -0.09
- Compact: -1.67, -1.55
- Intermediate: 3.45, 2.80
- Luxury: 2.64, 1.70
- Minivan: -4.26, -1.69

Always Collusion
- Small: 0.084, 0.23
- Compact: -0.219, -0.91
- Intermediate: 0.827, 3.10
- Luxury: 0.667, 1.93
- Minivan: -1.358, -2.12

D+E switch from collusion to Bertrand
- Small: -1.102, -2.22
- Compact: -0.952, -1.05
- Intermediate: 3.013, 3.08
- Luxury: 3.487, 3.35
- Minivan: -4.445, -2.46

Sargan test: 49.46, 50.19, 36.22
(28 degrees of freedom)

Voung-type test: -8.51, -9.21

Notes:
1. Instruments lagged 12 months imply that models with 12 and fewer observations must be removed.
2. Instruments: differences of segment-shares with respect to their time mean lagged 6 and 12 months, 31 brand dummies.
3. Standard errors are robust to serial correlation and heteroskedasticity across individuals.
4. Coefficients of behavioural variables are constrained to add up to zero.
### Table 4
Testing behaviour

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Collusion</td>
<td>Collusion</td>
<td>Coal. European</td>
<td>Bertrand</td>
</tr>
<tr>
<td>1</td>
<td>2</td>
<td>3</td>
<td>4</td>
</tr>
<tr>
<td>Coalition Domestic+European</td>
<td>2</td>
<td>3</td>
<td>4</td>
</tr>
<tr>
<td>Coalition Domestic</td>
<td>4</td>
<td>5</td>
<td>5</td>
</tr>
<tr>
<td>Bertrand</td>
<td>5</td>
<td>5</td>
<td>5</td>
</tr>
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</table>

<table>
<thead>
<tr>
<th>Test results (Vuong type test)</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>5.54</td>
<td>-9.21</td>
<td>-1.64</td>
<td>-1.20</td>
</tr>
<tr>
<td>2</td>
<td>-34.27</td>
<td>-6.70</td>
<td>-6.40</td>
<td></td>
</tr>
<tr>
<td>3</td>
<td>9.12</td>
<td>8.51</td>
<td></td>
<td></td>
</tr>
<tr>
<td>4</td>
<td></td>
<td></td>
<td>1.28</td>
<td></td>
</tr>
</tbody>
</table>

Model # versus model #. A row value above (below) the critical value of 1.96 (-1.96) means that the row model is better (worse) than the column model.
Figure 1
Sales evolution in the Spanish car market

Informe de la industria española, MCYT  Sample sales
Figure 2: Prices and received prices
Figure 3: Quality adjusted price index

- Domestic
- European
- Asian
Figure 4: Entry and quality change adjusted price index