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Exchange Rate Pass-Through and Dynamic Oligopoly: An Empirical Investigation[1]

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Abstract

This paper explicitly takes into account the dynamic oligopolistic rivalry among source producers to evaluate the degree of exchange rate pass-through. Using recent time-series techniques for the case of imported automobiles in Switzerland, the results show that prices are strategic complements and that the degree of pass-through is lower in the long run than in the short run. We attribute this to the fact that, although some rivals match long-term price changes, others do not, inducing the producer who faces a change in exchange rate to absorb a greater proportion of the variation.

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I. INTRODUCTION

The purpose of this paper is to analyze the dynamic properties of the exchange-rate pass-through in the presence of oligopolistic competition. The analysis takes advantage of recent developments in time-series econometrics to fully endogenize the pricing by producers from individual source countries and thereby investigate the effect of rivalry on the pass-through. To our knowledge, there is no empirical study in the pass-through literature which has analyzed how prices in a market interact following an exchange-rate shock. For instance, in his study of Japanese automobile imports to the US, Feenstra (1989) uses a competing price of imports to control for domestic competition but there is no feedback effect from the change in foreign price on the domestic price. Feenstra, Gagnon and Knetter (1996), and Gross and Schmitt (1996) construct aggregate prices for competitors to control for substitute products. These studies use single-equation estimation methods and do not explicitly develop the pricing interactions among rivals. The related literature on pricing-to-market is primarily interested in investigating how export prices can differ across destination markets (for instance, see Gagnon and Knetter, 1995, and Knetter, 1993), not how competing prices in a given destination market react.

Recognizing links among individual prices is an important undertaking since the theoretical literature on exchange rate pass-through shows that imperfect competition may be one of the main causes for the existence of incomplete pass-through (Dornbusch, 1987, Krugman, 1987). This suggests that when competition is oligopolistic, the interdependency between prices (both within and across periods) must be taken into account when assessing the pass-through relationship. The aim of this paper is a first attempt at filling this gap.

The empirical analysis deals with the Swiss automobile market where we treat oligopolistic rivalry among producers at the country level. The choice of the Swiss market for the study of exchange rate pass-through and oligopolistic interactions is interesting for at least two reasons. First, it is one of the few markets where there is no quantitative restriction particularly against Japanese imports. The results about the behavior of Japanese automobile producers can therefore be considered free of distortions.[3] As we shall see, Japanese producers play a key role in the analysis. Second, it is a market with no domestic producer and where no country-specific group of producers has a dominant position. More than 80 percent of the Swiss automobile market is shared by producers from only four source

[3] Gagnon and Knetter (1995) and Goldberg (1995) using very different frameworks show that the exchange rate pass-through is endogenous to trade restrictions.

countries.[4] This indicates that oligopolistic competition and thus, interdependence among prices should be explicitly taken into account when investigating the pass-through relationship.

Due to recent developments in time-series econometrics, several of the latest studies of exchange rate pass-through explicitly recognize the fact that exchange rate and price series are often non-stationary and may be cointegrated (see Feenstra, Gagnon and Knetter, 1996, Gross and Schmitt, 1996, Athukorala and Menon, 1995, 1994). To accommodate non-stationarity as well as the simultaneity of pricing, the empirical investigation uses a recent extension of Johansen's (1995) procedure developed in Johansen and Juselius (1995). It is a two-step method which first tests for cointegration in a system of equations and second, after imposing model-relevant restrictions, identifies the short-run structural parameters.

We find that dynamic price interdependence exists in this market both in the short and in the long run. In the short run for instance, we find that German and Japanese producers react to each other's price changes by about 30 percent after one period. The results show however that the degree of pass-through is relatively low compared to other findings. It is also lower in the long run than in the short run. We attribute this last result to the fact that, although some rival producers match long-term price changes, others do not, inducing the producer who faces a change in exchange rate to absorb a greater proportion of the variation with time.

The article is organized as follows. In Section 2, we motivate the empirical analysis by showing how the presence of switching costs and oligopolistic rivalry is consistent with the dynamic analysis of prices and exchange rate pass-through. The data set is described in Section 3 and the empirical implementation is presented in Section 4. The results are discussed in Section 5 and Section 6 concludes.

II. THEORETICAL FRAMEWORK

The aim of the empirical section is to investigate pass-through relationships in an oligopoly setting. As such, it does not test a particular theory but estimates the degree of exchange-rate pass-through in the short and long run, while introducing dynamic adjustments in price setting as well as price interdependence among sellers. The purpose of this theoretical section is to provide a model consistent with the dynamic set-up underlying Johansen's procedure. The implications of the model are then used to assist in the interpretation of the results. The chosen theoretical framework is Froot and Klemperer (1989)' s switching cost model and in

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[4] At the firm level, in 1990, more than 50 percent of all new registrations of passenger cars in Switzerland have been supplied by five producers only, all in our sample of countries: Opel, VW, Toyota, Ford, Peugeot and Renault (MWMA of US, 1992, p. 238). This percentage may underestimate the concentration ratio within different categories of automobiles. For studies of firms' behavior in the automobile industry, see Bresnahan (1981), and Mertens and Ginsburgh (1985).

this section, we first explain why this model is consistent with the empirical set-up and then we derive some of its comparative static properties.

With switching costs, buyers of a product have real or perceived costs of switching to a competing product even if there is no significant difference between products (Klemperer, 1995). Switching costs seem particularly relevant to the automobile market. For instance, consumers may be uncertain about the quality of brands they have not driven. The use of after-sale services can make consumers feel they know better the producer's local representative of the automobile they currently own than other retailers. They may perceive that such a business relationship will bring them a better deal in future transactions than with representatives they do not know. Consumers may also develop a non-economic or a psychological brand loyalty which makes switching brand names difficult.

The key implication is that, if sellers realize that consumers face switching costs, today's market shares determine tomorrow's profit. In other words, price decisions have an intertemporal link that would not exist without them. Of course, there are several other possible channels generating such a link.[5] Switching costs, however, by establishing these links purely on demand considerations, seem especially relevant to analyse the case of a small automobile market without domestic producers.

As an illustration, consider two foreign producers serving a market where there is no domestic production. Assume the foreign firms experience different variations of the exchange rate with respect to the currency of the buyer's market. To simplify notation, let Firm A produce in country A and Firm B, in country B. To establish some of the intertemporal links, we set the model in a two-period game (see Klemperer, 1995). The discounted profit of producer k expressed in its own currency is

$$V^k = e_1^k \Pi_1^k[p_1^A, p_1^B, X_1^k] + \lambda^k e_2^k \Pi_2^k[s^k(p_1^A, p_1^B), p_2^A, p_2^B, X_2^k], \qquad \textbf{(1)}$$

where e_t^k is period t (t=1,2) exchange rate between seller k and the buyer's currency (expressed in units of the seller currency per unit of the buyer's currency); Π_t^k is firm k's profit in period t in the buyer's currency; λ^k is the discount factor faced by k; p_t^k is the price of firm k (in the buyer's currency); s^k is firm k's market share resulting from competition in the first

[5] For instance, limited capacity of production or distribution, adjustment costs in production or in prices such as when contracts exist between producers and retailers (menu costs), or else sunk investments. Limited capacity of distribution does not seem to be a major constraint since most producing firms are also wholesalers on the Swiss market.

period, and X_t^k is the vector of exogenous variables which directly affect profit in period t. For our purpose, the main variables in X_t^k are k's cost of production and k's exchange rate.[6]

Suppose firms choose prices in every period. To find the equilibrium prices, we look for the closed-loop equilibrium. That is, we look for the subgame perfect equilibrium of the game where players can observe and respond to their rival's price at every period. The optimal prices in the second period, conditional on prices realized in period 1, are found by maximizing the second expression in (1) with respect to price. If D_2^k is the demand faced by firm k in period 2 and c^k is k's constant marginal cost expressed in the seller currency, Firm k (k=A,B) chooses p_2^k maximizing $\Pi_2^k=(p_2^k-c^k/e_2^k)D_2^k(s^k(p_1^A, p_1^B),p_2^A,p_2^B)$. The demand D_2^k depends on the market share s^k realized in the first period since some consumers who bought then are now locked in: Prices set in the first period determine the position of the demand in the second period. The price p_2^k satisfies the usual first-order condition,

$$D_2^k + (p_2^k - \frac{c^k}{e_2^k})\frac{\partial D_2^k}{\partial p_2^k} = 0, \qquad k=A,B. \tag{2}$$

Assuming that the second-order conditions hold, the Nash equilibrium in prices for period 2 is

$$\begin{aligned} \hat{p}_2^A &= f_2(p_1^A, p_1^B, X_2^A, X_2^B), \\ \hat{p}_2^B &= g_2(p_1^A, p_1^B, X_2^A, X_2^B), \end{aligned} \tag{3}$$

where X_2^A (X_2^B) is Firm A's (Firm B's) vector of period 2 exogenous variables. As it is apparent from (2), X_2^k (k=A, B) must contain c^k and e_2^k.

A firm which takes into account the role of switching costs considers the effect of its price in period 1 on its profit in period 2. To find the optimal prices in period 1, substitute (3) in (1) so as to write the discounted profit as

$$V^k = e_1^k \Pi_1^k(p_1^A, p_1^B, X_1^k) + \lambda^k e_2^k \pi_2^k(p_1^A, p_1^B, X_2^A, X_2^B), \; k=A,B. \tag{4}$$

where π_2^k is a different function than Π_2^k since it is no longer a function of period 2's prices. The first-order conditions in prices are then

$$V_k^k = e_1^k \frac{\partial \Pi_1^k}{\partial p_1^k} + \lambda^k e_2^k (\frac{\partial \pi_2^k}{\partial p_1^k}) = 0 \qquad k=A,B. \tag{5}$$

[6] Own exchange rate is an element of this vector since costs are in the seller currency while Π_t^k is in the buyer's currency.

Since $\partial\pi_2^k/\partial p_1^k<0$ (firms operate in the elastic portion of their demand), $\partial\Pi_1^k/\partial p_1^k>0$, so that prices in period 1 are lower than without switching costs. Assuming the second-order and the stability conditions hold,[7] the perfect Nash equilibrium prices are found by solving the system of best replies (5). The solutions depend only on the exogenous variables of the model:

$$\begin{aligned}\hat{p}_1^A&=f_1(X_1^A,X_1^B,X_2^A,X_2^B),\\ \hat{p}_1^B&=g_1(X_1^A,X_1^B,X_2^A,X_2^B).\end{aligned} \tag{6}$$

For the purpose of this paper, the equilibrium prices (3) and (6) have two important implications. First, a variation in one exchange rate in the first period affects firms' prices over both periods. Consider a change in e_1^A. Since it is an element of X_1^A, both $\hat{p}_1^A$ and $\hat{p}_1^B$ are affected by a change in e_1^A and since these prices determine the equilibrium prices in period 2 (see (3)), $\hat{p}_2^A$ and $\hat{p}_2^B$ are also both affected by a change in e_1^A. Second, in the long run, there is a unique set of prices and thus, a unique set of market shares. If the long-run equilibrium is such that the exogenous variables are constant across periods ($X^A=X_1^A=X_2^A$ and $X^B=X_1^B=X_2^B$), the long-run equilibrium prices (6) reduce to $p^A=f(X^A, X^B)$ and $p^B=g(X^A, X^B)$, and $s^A(p^A, p^B)$ and $s^B=1-s^A$ are uniquely determined. Hence, following an exchange rate variation, prices ultimately converge to the long-run equilibrium consistent with the exogenous variables.

A look at some of the comparative static properties of the model gives some insight into how prices vary following a change in e_1^A (see also Tivig, 1996). Consider the effect of a change in e_1^A on first-period prices, that is the short-run degree of pass-through. Differentiating (5),

$$\begin{aligned}V_{AA}^A dp_1^A+V_{AB}^A dp_1^B&=-V_{Ae}^A de_1^A,\\ V_{AA}^A dp_1^A+V_{AB}^A dp_1^B&=-V_{Ae}^A de_1^A.\end{aligned} \tag{7}$$

Solving (7),

$$\begin{aligned}\frac{dp_1^A}{de_1^A}&=\frac{1}{V}(V_{Be}^B V_{AB}^A-V_{Ae}^A V_{BB}^B),\\ \frac{dp_1^B}{de_1^A}&=\frac{1}{V}(V_{Ae}^A V_{BA}^B-V_{Be}^B V_{AA}^A).\end{aligned} \tag{8}$$

Consider the own price effect. Since $\Pi_1^A=(p_1^A - c^A/e_1^A)D_1^A$, then, using (5), $V_{Ae}^A=D_1^A(1-\epsilon_1^A)$,

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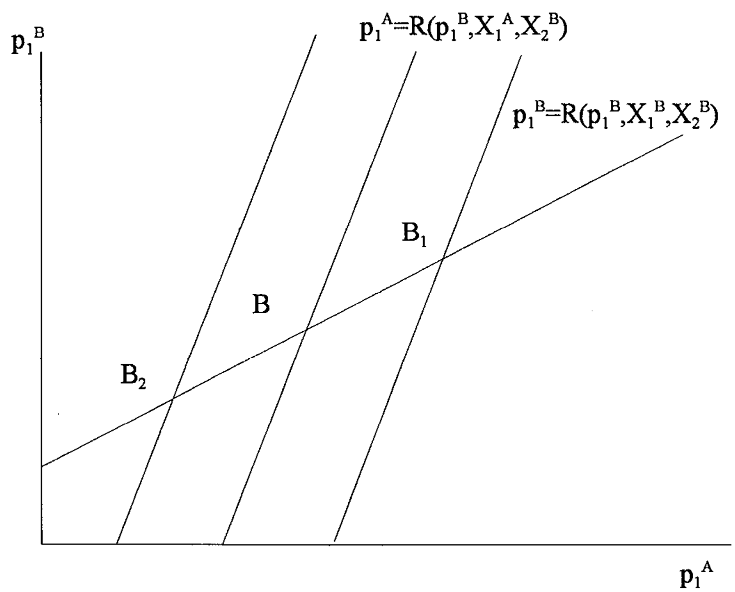
[7] In particular, $V_{kk}^k < 0$ where $V_{kk}^k=\partial V_k^k/\partial p_1^k$, $V=V_{AA}^A\, V_{BB}^B - V_{AB}^A\, V_{BA}^B> 0$, $|V_{kk}^k| > |V_{kj}^k|$, k,j=A,B, j≠k. When prices are strategic complements, we also know that $V_{kj}^k>0$.

where $\epsilon_1^A = -(\partial D_1^A/\partial p_1^A)\, p_1^A/D_1^A$ is the elasticity of the first-period import demand from A (Tivig, 1996). Suppose Firm B's optimal first-period price is not directly affected by a variation of A's exchange rate (i.e. $V_{Be}^B=0$).[8] In this case, substituting V_{Ae}^A in (8),

$$\frac{dp_1^A}{de_1^A} = -\frac{D_1^A}{V}(1-\epsilon_1^A)V_{BB}^B; \qquad \frac{dp_1^B}{de_1^A} = \frac{D_1^A}{V}(1-\epsilon_1^A)V_{BA}^B. \tag{9}$$

Since $V_{BB}^B < 0$ by the second-order conditions, $V > 0$ (see footnote 5) and $V_{BA}^B > 0$ with strategic complementarity, the sign of both expressions is positive if ϵ_1^A is inelastic (perverse case) and negative otherwise (normal case). Moreover, since $|V_{BB}^B| > |V_{BA}^B|$ implies $|dp_1^A/de_1^A| > |dp_1^B/de_1^A|$ and the own price effect is greater than the induced effect on the rival's price.

Figure 1. Short-run rivalry



[8] This assumption is consistent with imperfect capital mobility (see Froot and Klemperer, 1989). Alternatively, the uncovered interest parity condition coupled with the hypothesis introduced later that the exchange rates follow a random walk also lead to $V_{Be}^B=0$.

In Figure 1, B_2 represents the perverse case following an appreciation of A's currency (e_1^A falls) whereas B_1 represents the normal case. Clearly the short-run pass-through can be low even if prices are strategic complements. For this, it suffices ϵ_1^A be close to one. The results are quite different from a simple static Bertrand equilibrium. In a static game, price k is determined by $V_k^k = e^k(\partial\Pi^k/\partial p^k) = 0$ (from (5) without time subscript), so that $V_{Ae}^A = c^A/e^A(\partial D^A/\partial p^A) < 0$. Using (8), since $V_{Be}^B = 0$, the effects of a change in e^A on prices in a static Bertrand game are,

$$\frac{dp^A}{de^A} = -\frac{1}{V}\frac{c^A}{e^A}\frac{\partial D^A}{\partial p^A}V_{BB}^B < 0; \qquad \frac{dp^B}{de^A} = \frac{1}{V}\frac{c^A}{e^A}\frac{\partial D^A}{\partial p^A}V_{BA}^B < 0.$$

There is no perverse effect and the change in price is simply proportional to the cost effect of the change in e^A. Like in (9), $|dp^A/de^A| > |dp^B/de^A|$. There is no guarantee that the degree of pass- through from this static game is higher than the short-run one given by (9). Still, the less elastic the import demand is, the lower the short-run degree of pass-through compared to the static case.

The intuition for these results is clear. With a static Bertrand game, an appreciation of A's currency increases the cost of production expressed in the buyer's currency. As a result, p_1^A increases and, since prices are strategic complements, so does p_1^B. With switching costs, the pass-through effect is richer because of the intertemporal effect. Firm A not only takes into account the effect of a price change on today's profit but also on tomorrow's profit. In particular, if the import demand is inelastic, first-period changes in profit are small and faced with an appreciation of its currency, A chooses to decrease its price instead of increasing it. It does so because future profits have become more important than the immediate cost effect. When the demand is elastic, the effect of a change in e_1^A is normal as the cost effect dominates the intertemporal effect.

In addition of changing the nature of the first-period effect, switching costs have a dynamic effect on prices that a static Bertrand game cannot generate. As a result, the long-term degree of pass-through may differ from the short-run one. To get some intuition into the effects of a change in e_1^A on period 2 prices, we use (3),

$$\frac{dp_2^A}{de_1^A} = \frac{\partial f_2}{\partial p_1^A}\frac{\partial p_1^A}{\partial e_1^A} + \frac{\partial f_2}{\partial p_1^B}\frac{\partial p_1^B}{\partial e_1^A};$$
$$\frac{dp_2^B}{de_1^A} = \frac{\partial g_2}{\partial p_1^A}\frac{\partial p_1^A}{\partial e_1^A} + \frac{\partial g_2}{\partial p_1^B}\frac{\partial p_1^B}{\partial e_1^A}.$$

The price reaction in period 2 is composed of two effects: The strategic and the own price effects. Suppose the own price effects $\partial f_2/\partial p_1^A$ and $\partial g_2/\partial p_1^B$ are negative (i.e., a high own price yesterday induces a lower own price today) and $\partial f_2/\partial p_1^B$ and $\partial g_2/\partial p_1^A$ are positive (i.e., prices are strategic complements across periods). In this case, the sign of dp_2^A/de_1^A and dp_2^B/de_1^A are uncertain. Consider then two possibilities. First, suppose $\partial p_1^B/\partial e_1^A \approx 0$ so that a change in e_1^A has no immediate effect on the rival's price. In this case, the second-period price effects have opposite signs. Second, suppose $\partial p_1^B/\partial e_1^A \neq 0$ but $\partial f_2/\partial p_1^k$ and $\partial g_2/\partial p_1^k$(k=A,B) have the same (absolute) value.[9] In this case, $dp_2^A/de_1^A = -dp_2^B/de_1^A$ and period 2 prices also change in opposite directions. In short, there are reasonable conditions for which the price effects can be expected to have opposite signs in the second period. These changes are also smaller than in the first period since both changes are functions of first period price changes.

Combined with (9), these results also tell us that it is sufficient to have a normal effect in period 1 for the long-run degree of pass-through to be smaller than the short-run one: With a normal effect in period 1, an appreciation of A's currency leads to a loss of market share for Firm A since $dp_1^A/de_1^A > dp_1^B/de_1^A > 0$. If period 2 effects have opposite signs then $dp_2^B/de_1^A > 0$ and $dp_2^A/de_1^A < 0$ to avoid further losses in A's market share which decreases the degree of pass-through.

Even if this two-period model captures well the intertemporal effects of switching costs on the pricing decision in any period, it masks the fact that there is typically no first period to the game. In other word, in every period, firms inherit consumers from past periods and must decide which price to set, not only to lock in new consumers but also to exploit old consumers who have been locked in the past. This balancing act implies that (1) can be written as

$$V_t^k = e_t^k \Pi_t^k(s^k(p_{t-1}^A, p_{t-1}^B), p_t^A, p_t^B, X_t^k) + \lambda e_{t-1}^k \Pi_{t-1}^k(s^k(p_t^A, p_t^B), p_{t-1}^A, p_{t-1}^B, X_{t-1}^k) \ .$$

The best reply for Firm k in period t, taking into account the optimal prices in period t+1, can be written as

$$p_t^k = R_t(p_t^j, p_{t-1}^A, p_{t-1}^B, X_t^k, X_{t+1}^k) \ , \qquad \textbf{(10)}$$

for k,j=A,B; j≠k. The intertemporal effect is now mitigated by the incentive to exploit old consumers.[10] Equation (10) is the reference equation for the empirical implementation of the model.

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[9] Tivig (1996) shows this is the case in a Hotelling model and thus, in a model where demand for a product depends on the price difference between imperfect substitutes. This is a reasonable representation of a mature market like the automobile market since it implies that no firm can increase its market share without decreasing its rival's share.

[10] Klemperer (1995) argues that, because of the incentive to exploit old customers, the presence of switching costs increases the equilibrium prices with respect to the equilibrium without switching costs.

III. THE DATA

The empirical investigation concentrates on the pricing behavior of automobile producers from *several* source countries in a *single* export market from 1977 to 1994.[11] By considering a case involving a single export market, the analysis focuses on the exchange rate pass-through namely, how much of the exchange rate variations are passed on prices by producers. Switzerland is the export market. One advantage of this market is that, during the sample period, there was no quantitative restriction and, when in place, tariffs were very low.[12] In other studies, quantitative restrictions on Japanese automobiles either have resulted in the exclusion of this major producer (see Feenstra, Gagnon, Knetter, 1996) or have been shown to affect the size of the pass-through parameter (Gagnon and Knetter, 1995, Goldberg, 1995).

The source countries under consideration are Belgium (B), France (F), Germany (G) and Japan (J).[13] These countries capture 85 percent to 90 percent of the market over the sample period (see Gross and Schmitt , 1996, for individual market shares). Other source countries have been excluded from the analysis because they are marginal exporters to the Swiss market (for examples, Sweden and the United States) or because they entered the market only recently (like Korea and the Czech Republic). In both cases, observations are often zero at the quarterly frequency. One relatively important exclusion is Italy for which unit-values are highly variable due to the heterogeneity of the products.

The prices (P^B, P^F, P^G, P^J) are measured as unit-values (at the border before tariffs).[14] They are in currency of the destination country (i.e., Swiss Francs). Two categories are analyzed separately, small-size and medium-size automobiles. Border records categorize automobiles according to the weight. Thus, in the small-size category are vehicles between 800 and 1,200

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[11] The classification of automobiles changed at the end of 1976 and earlier series are not consistent with the new definitions.

[12] Until 1987, the tariffs per 100 kilograms for non-EC/EFTA automobiles were SFr96 or less and for EC/EFTA automobiles, SFr81. After 1987, there was no tariff on EC/EFTA automobiles and a uniform per-unit tariff of SFr81 for non EC/EFTA automobiles (*Statistique Mensuelle du Commerce Extérieur de la Suisse,* Direction Générale des Douanes, Bern). The change in tariffs has been shown to have no significant impact on the pricing strategy (Gross and Schmitt, 1996).

[13] In our trade statistics, the source country is the country of production and, in addition to the well known brand names, Germany is the source country for Opel (GM Europe). Ford's production originates either from Germany or from Belgium (Commission of the EC, 1983, p.61).

[14] Information on the total values and number of units are from the *Statistique Mensuelle du Commerce Extérieur de la Suisse*, Direction Générale des Douanes, Bern.

kilograms and, in the medium-size category, vehicles between 1,200 to 1,600 kilograms. We excluded the category above 1,600 kilograms because the characteristics of the products are likely to have changed significantly during the period (for example with the development of the market for four-wheel drive recreational vehicles).

The remaining variables are the cost of production in the source countries (C^B, C^F, C^G, C^J) which is approximated by unit labor cost in manufacturing in all countries except Belgium for which the producer price index in manufacturing is used. Clearly, the unit labor cost is only a proxy for the marginal cost of production since it omits raw materials for instance (see Goldberg and Knetter, 1997).[15] We expect this omission not to create biased exchange rate coefficients since the Dollar, the main currency for pricing raw materials, is not in our specifications. The measure we use is partly determined by data availability for long time-series and partly by our empirical approach. The choice to focus on rivalry and estimate dynamic simultaneous best reply functions has at least two implications: First, we cannot estimate from our dataset more precise cost measures as it is done in the PTM literature.[16] Second, in the time-series context the number of degrees of freedom shrinks rapidly, we therefore also assume (and test) non-endogeneity of marginal cost determination with respect to prices. This amounts to assuming that the marginal cost in producing countries is constant within the range of variation of their market shares in Switzerland.[17] The nominal exchange

[15] Labor compensation is the largest component of value-added in the two-digit level industry including automobile production in all countries. In 1983, for our countries, it was between 62 percent and 80 percent of added-value and in 1991, between 56 percent and 76 percent (*National Accounts Statistics. Main Aggregates and Detailed Tables*. United Nations. New-York). In Germany and in France, for the automobile industry alone labor compensation represented approximately 76 percent on average between 1983 and 1989 (Williams et al., 1994, Table 9.7).

[16] Gagnon and Knetter (1995), for example, can exploit the time-series and cross-sectional dimensions of their panel to estimate costs as the common component in prices charged by one source country to several destination markets for a given product.

[17] In 1988, for example, total exports of German automobiles to Switzerland represented 2.9 percent of total production. For France and Japan they represented 1.4 percent and 1.2 percent respectively. (*Automobile Revue*, Special annual issue published on the occasion of the 64th International Geneva Motor Show. Hallwag, Bern. 1990). Goldberg and Knetter (1997) rightly points out that costs might be endogenous to price in the case of larger markets.

rates (E^B, E^F, E^G, E^J) are defined as source-country currency per Swiss Franc.[18] The variables are observed at the quarterly frequency between 1977 and 1994. All the variables are in logarithms and seasonally unadjusted except the unit labor costs.

Pre-testing indicates that all the series are non-stationary. The results of Augmented Dickey-Fuller (ADF) tests in Table 1 show that for all series the hypothesis of integration of order 1 cannot be rejected while that of integration of order 2 can be rejected. Hence, the specification as well as the estimating methodology must take this feature into account and both are detailed in the next section.

[18] The costs series are from *Main Economic Indicators,* Organisation for Economic Cooperation and Development, Paris. The exchange rate series are from the *Bulletin Mensuel,* Swiss National Bank, Zurich.

Table 1. Unit-Root Tests

	I[1]		I[2]
	ADF(6)	ADF(8)	ADF(n)
	α=0		c=α=0
BELGIUM			
P^B (small)	-2.28	-2.41	-4.80(2)**
P^B (medium)	-1.78	-2.23	-5.30(2)**
C^B	-2.58	-2.49	-2.25(2)**
E^B	-1.31	-2.28	-3.51(4)**
FRANCE			
P^F (small)	-1.06	-1.54	-5.32(2)**
P^F (medium)	-0.71	-1.16	-4.13(2)**
C^F	-2.73*	-2.31	-2.87(0)**a
E^F	-1.48	-2.34	-3.27(4)**
GERMANY			
P^G (small)	-1.68	-1.56	-5.15(2)**
P^G (medium)	-1.20	-1.36	-5.48(2)**
C^G	-3.15*b	-2.50b	-3.95(2)**
E^G	-1.44	-2.06	-4.63(4)**
JAPAN			
P^J (small)	-2.74*	-2.26	-4.54(2)**
P^J (medium)	-0.80	-1.42	-4.60(2)**
C^J	-2.39b	-2.43b	-2.77(4)**
E^J	-2.89b	-2.96b	-4.58(4)**

*Significant at 10 percent, ** significant at 5 percent.
The null hypothesis is the series is I[1] or I[2].
a With constant.
b With time trend.

IV. EMPIRICAL INVESTIGATION

The main theoretical implications of the model in Section 2 are threefold: First, exchange rate variations have intertemporal effects on price-setting; second, rivalry may generate a lower exchange rate pass-through in the long-run than in the short-run; third, in the long-run there is a stable relationship between rivals' prices. The goal of this empirical section is to investigate the role of rivalry on the magnitude of the exchange rate pass-through, in the long run and in the short run, in light of these implications.

Based on (10), the model to be estimated is a set of best reply functions for producers from four source countries where price levels are determined simultaneously. The presence of future values requires an hypothesis about expectation formation. It is assumed that costs and exchanges rates follow a random walk so that if producers form rational expectations, each pricing equation becomes solely a function of present and past values,

$$p_t^k = f[c_t^k, c_{t-i}^k, e_t^k, e_{t-i}^k, p_t^j, p_{t-i}^j] \ , \tag{11}$$

where k,j=B, F, G, J, j≠k and I=1 to n.

Since all the series were found to be integrated of order 1, the system of simultaneous pricing consistent with (11) must be specified in first differences. Nevertheless, if the series are cointegrated, a long-run relationship between the levels can still be identified. The appropriate econometric methodology to identify cointegration in systems is developed in Johansen (1995).[19] We use a recent extension of that methodology to estimate the short-run structural parameters of the best reply functions (11) as well as test for cointegration (see, Johansen and Juselius, 1995). It is a two-step procedure such that, in the first step, the presence of cointegrating vectors is tested using a fully *unrestricted* model. This step tests for cointegration between the four source-country price levels in each category. It is therefore consistent with the existence of a stable long-run relationship between prices (and thus, market shares) as established in the previous section. In the second step, structural parameters for the pricing equations in first differences are estimated using the cointegration relationship between prices as an error-correction mechanism. Restrictions consistent with (10) are imposed such that, producers in each source country react to rivals′ prices but not to rivals' costs and exchange rates.

A. First-Step: Cointegration Analysis

The Johansen methodology requires the specification of a VAR model. It is well known however, that characteristics such as the number of lags or the presence of a vector of deterministic variables (constant, seasonal variables) may affect the results (see Hargreaves, 1994). Also, in VAR specifications, the number of degrees of freedom shrinks very rapidly.

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[19] See also Banerjee et al. (1993), Engle and Granger (1991).

Our strategy is therefore to design the most parsimonious specification consistent with normality of residuals since the results are valid only for well-behaved errors.

The starting number of lags is set equal to 2 for level variables. A vector of constants is introduced and tested for significance. The specification for each automobile category includes four fully endogenous prices (P^B, P^F, P^G, P^J), four weakly exogenous cost variables (C^B, C^F, C^G, C^J) and four weakly exogenous exchange rate variables (E^B, E^F, E^G, E^J). Following footnote 15, we expect costs not to be endogenous to the prices of automobiles. Also, the unit labor cost is unlikely to be sensitive to exchange rate variations. The restriction of weak exogeneity for costs and exchange rates is imposed at the outset and formally tested after the determination of the cointegrating relationships.[20] The stability of the parameters is tested over the last 24 observations of the sample (1989.1-1994.4). In this first step of the analysis, the identification of cointegration between prices is based on an unrestricted model, i.e. all the variables enter all price equations and there is no restriction on the parameter values.

For the small-size category normality of the residuals is achieved with a constant, three seasonals and two lags (see Table 2, upper panel).

Table 2. Tests for the Residuals of the VAR Estimations

	P^B	P^F	P^G	P^J
Small Automobiles: 1977.1-1991.3				
Chi²(2) for normality[a]	.794	.158	.086	1.678
Chi²(16) for serial correlation[b]	18.995	19.328	14.274	13.503
Medium Automobiles: 1977.1-1994.4				
Chi²(2) for normality	2.531	.742	1.630	.911
Chi²(16) for serial correlation	19.886	15.309	19.036	21.074

[a] χ^2= [(T-m)/6]*[SK²+(1/4)EK²] where m is the number of regressors, SK is skewness and EK is excess kurtosis. Jarque and Bera. (1980)
[b] $\chi^2 = T\Sigma r_j^2$ for j=1,2,..., 16. Box and Pierce. (1970).
The critical values at 5 percent (10 percent) are 5.991 (9.210) for 2 degrees of freedom and 26.296 (31.999) for 16 degrees of freedom respectively.

The period however is shortened to 1977.1 until 1991.3 since values of predictive Chow-tests over the last 24 observations indicate a clear break in several quarters after 1991.3. For all

[20] As the number of variables is large (12 in each category of automobiles), we follow Johansen (1992a, 1992b) who shows that the full model can be divided into a conditional model and a marginal model where weak exogeneity becomes a linear restriction on the marginal model. The test is then an F-test in the marginal model such that the coefficient on the cointegrating vector identified by the reduced-rank analysis is zero.

source countries, but especially for France and Germany, the model cannot predict the pricing of small automobiles beyond 1991 without systematic errors (parameter constancy between 1991.4 and 1994.4 is rejected such that $F(52,29) = 5.15$; p-value=.000). The break may be associated with the introduction in 1992 of the single European market. Although Switzerland is not a member of the European Union, the pricing in the Swiss market may still have been affected when producers adjusted their strategies for the European market.

In the medium-size category of automobiles, normality of the residuals is achieved with two lags and no constant. The hypothesis of no constant could not be rejected by the L.R. test at 5 percent level significance ($\chi^2[3]=4.102$). In that category, forecasts over the last 24 observations could not identify a major break in pricing.

The results of the tests for cointegration are presented in Tables 3 along with the 5 percent critical level of significance.

Table 3. System Estimation and Cointegration Tests

	Small-size automobiles: 1977.1-1991.3			*Medium-size automobiles: 1977.1-1994.4*		
	λ_{max} [a]	Trace[b]	Critical value	λ_{max}	Trace	Critical value
$r \leq 3$	2.986	2.986	3.84	1.667	1.667	9.13
$r \leq 2$	6.406	9.392	15.34	7.931	9.598	19.99
$r \leq 1$	13.071	22.464	29.38	16.044	25.642	34.80
$r=0$	41.101	63.565	47.21	27.683	53.325	53.42

Critical values from Johansen (1995), Table 15.2, 15.3. Results derived with PcGive (Hendry, 1989).
[a] $-T\log(1-\mu_i)$ where μ_i is the eigenvalue i. H_0: $r=r_0$ is tested against H_A: $r=r_0+1$.
[b] $-T\Sigma og(1-\mu_i)$ where μ_i is the eigenvalue i. H_0: $r\leq r_0$ is tested against H_A: $r>r_0$. Johansen and Juselius (1990).

Starting with the small-size automobiles, the hypothesis of one cointegrating vector (i.e. $r\leq 1$) cannot be rejected and the valid cointegrating relationship is

$$P^G + .33328P^J - .26233P^B - 1.21247P^F = 0 \ . \qquad \textbf{(12)}$$

The case of the medium-size category is less clear-cut. The value of the Trace test is slightly below the critical value for r=0. The hypothesis of no cointegration is however rejected at

10 percent significance.[21] Also, following Granger (1997), in a preliminary analysis, we investigated cointegration in pairs of prices. In both categories, the hypothesis of no-cointegration is rejected for all pairs of prices.[22] Therefore, we consider that the presence of one cointegrating vector cannot be rejected in the medium-size category; it is

$$P^{G} - .54824P^{J} - .0746P^{B} - .40936P^{F} = 0 . \qquad \textbf{(13)}$$

Once the cointegrating vectors are identified, the hypothesis of weak exogeneity of costs and exchanges rates is formally tested. The F-values are 0.54 (p-value=.82) for the small-size category and 1.28 (p-value=.27) for the medium-size category so that the weak exogeneity of costs and exchange rates cannot be rejected.

The cointegrating vectors (12) and (13) identify a stable relationship between price levels by producers from the four source countries per category of automobiles. The existence of these relationships is consistent with the results of the theoretical section. In effect, we have established that, in a game among interdependent producers, the long-run equilibrium prices (and thus market shares) are unique. This implies the existence, but not the uniqueness, of a relationship between these prices. The fact that there is at least one cointegrating vector defining a stable relationship between long-term prices indicates that sellers may indeed aim at long-run market shares.

B. Second-Step: Estimation of the short-run parameters

The second step is the estimation of the structural parameters for the price-setting equations in differences. It requires taking into account the restrictions imposed by the theoretical framework. In particular, each source country prices automobiles according to its own cost, own exchange rate and competitors' prices. The constraints that the coefficients on other countries' cost and exchange rates are zero are thus explicitly introduced. There are two exceptions which relate to the fact that most of the automobiles exported from Belgium are produced by German owned plants. We let the German cost and exchange rate variables enter unrestrictedly the pricing equation from Belgium and the Belgian cost and exchange rate

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[21] These tests, like unit-root tests, have low power in small samples and it is recommended to use 10 percent significance level (see Dickey and Rossana, 1994, Section IIIB).

[22] We used the Engle-Granger (1987) procedure to test for cointegration in pairs of prices. The results of the cointegration tests are, for the medium-size/small-size automobiles, t(D,J)=-4.39/-2.40, t(D,B)=-3.09/-2.0, t(D,F)=-2.49/-4.0, t(J,B)=-3.58/-2.55, t(J,F)=-2.40/-2.28, t(B,F)=-3.24/-3.20. The critical value at 5 percent is t=-1.95.

variables enter unrestrictedly the German pricing equation.[23] Also, in this second step, following Johansen and Juselius (1995), we fix the coefficients of the cointegrating vectors (12) and (13) and treat them as error-correction terms.[24] In addition, the final specifications are obtained by following the general to specific procedure thereby constraining to zero the coefficients with very low level of significance (t-value below 1.0). In both categories imposing cross-country restrictions on costs and exchange rates led to serial correlation in some equations. We therefore introduced one more lag in the dependent variables to be able to interpret the final results. The final system is estimated by Three-Stage-Least-Squares and the results are given in Table 4 for small-size automobiles, and in Table 5 for medium-size automobiles.

Before analyzing the results, a few comments about their robustness are in order. First, in both categories the residuals are normally distributed and serial correlation is absent. Second, in spite of the numerous restrictions dictated by the theoretical model, our final system does not appear to be misspecified (the values for the L.R. test for overidentifying restrictions are reasonably close to the critical values). Third, the coefficients on the ecm terms (i.e., the cointegrating vectors with fixed parameters) are remarkably similar to those found in the unrestricted estimation. In both categories these coefficients are somewhat smaller in the second stage. This can be attributed to the introduction of an additional lag in the dependent variable for some equations to correct for serial correlation in the residuals. The similarity of values suggests that the restrictions of zero-effect by other countries' costs and exchange rates are well supported by the data. Hence, source-country producers follow a strategy where in addition to competitors' prices, they care only about their own costs and their own exchange rate. Finally, the significant constants in the small-size category indicate some quality changes may have occurred during the period.

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[23] The relevance of the German-Belgium cross-effect (as opposed to other country combinations) is supported by evidence on the ratio of the value of foreign vs domestic parts in each automobile. For examples, in 1979, the ratio of German parts in a French automobile was only 9.3 percent but it was 106 percent in a Belgian automobile. The ratio of French parts in a German automobile was 6.2 percent (own calculations from Commission of the EC, 1983, p. 120 ff.).

[24] The other cointegrating vectors are constrained to have zero coefficients. This does not appear to be in violation to the unrestricted specification since their respective weights in the first-step are all very small.

Table 4. Pricing Equations: Small-Size Automobiles, 1977.1-1991.3

	ΔP^k_t			
	Belgium (k=B)	*France (k=F)*	*Germany (k=G)*	*Japan (k=J)*
constant	-.125 (0.8)	.230 (2.3)	-.358 (2.4)	-.229 (2.2)
ΔC^k_{t-i}	-	-	-	i=0 -.186 (1.2)
ΔE^k_{t-i}	i=1 .327 (2.6)	-	i=0 -.228 (1.4)	i=1 -.283 (4.4)
ΔP^B_{t-i}	i=1 -.396 (3.5)	i=0 .249 (2.2)	-	i=1 -.281 (3.9)
ΔP^F_{t-i}	-	-	-	-
ΔP^G_{t-i}	i=1 .283 (2.2)	i=1 -.205 (2.4)	i=1,2 -.271 (2.1) -.278 (2.7)	i=1 .242 (2.7)
ΔP^J_{t-i}	-	i=1 -.159 (1.5)	i=1 .286 (2.0)	i=1 -.404 (3.8)
ecm_{t-1}	- .114 (0.9)	.213 (2.6)	-.320 (2.5)	-.208 (2.3)
ΔC^G_{t-i}	i=1 .511 (1.6)	n.a.	n.a.	n.a.
ΔC^B_{t-i}	n.a.	n.a.	i=1 -.608 (2.6)	n.a
n	59	59	59	59
Corr. actual and predicted	.490	.400	.590	.628
Norm.resid. $\chi^2(2)$	2.601	1.047	2.408	2.100
4-lag serial corr. $\chi^2(16)$	14.360	16.712	17.018	20.953
L.R. test for overindentifying restrictions $\chi^2(67)$=92.99				

Absolute t-value in parentheses.

Table 5. Pricing Equations: Medium-Size Automobiles: 1977.1-1994.4

	ΔP^k_t			
	Belgium (k=B)	*France (k=F)*	*Germany (k=G)*	*Japan (k=J)*
ΔC^k_{t-i}	i=0 .529 (1.0)	i=1 -1.399 (3.7)	i=1 .351 (1.3)	i=1 -.601 (1.4)
ΔE^k_{t-i}	-	i=1 .325 (2.1)	i=0 -.514 (2.6)	i=0 -.210 (1.2)
ΔP^B_{t-i}	i=1 -.412 (3.8)	-	-	-
ΔP^F_{t-i}	-	i=1 -.411 (4.2)	-	i=1 .274 (1.6)
ΔP^G_{t-i}	-	-	i=1,2 -.368 (3.3) -.322 (3.0)	-
ΔP^J_{t-i}	-	i=1,2 .116 (1.9) .175 (2.8)	-	i=2 -.196 (1.7)
ecm_{t-1}	.044 (0.9)	.191 (4.1)	.054 (1.8)	.086 (1.9)
ΔC^G_{t-i}	(k=G) i=0 .943 (2.1)	n.a.	-	n.a.
n	70	70	70	70
Corr. actual and predicted	.433	.596	.460	.383
Norm.resid. $\chi^2(2)$	2.126	1.044	1.218	8.873
Ser. Corr. $\chi^2(16)$	16.680	8.161	11.584	11.479
L.R. test for overindentifying restrictions $\chi^2(67)$=85.62				
Tests of parameter constancy over 1992.1-1994.4 restricted specification: Cumulative test: $\chi^2(48)/48$ =.998. Forecast F-form: F(48,52) = .910 (p-value=.63).				

Absolute t-value in parentheses.

V. RESULTS

We divide the discussion of the results in three parts. First, we look at the short-run impact of changes in exchange rate on producers' own prices. Second, we look at price rivalry among source countries. Third, we investigate the dynamics of permanent and temporary exchange rate variations. We pay particular attention to the long-run degree of exchange rate pass-through (that is, once all adjustments have taken place) and the dynamics of the price rivalry between source countries. When possible, we compare our results with those of other studies.

A. Exchange Rate Pass-Through

We start with the short-run exchange rate pass-through which in Tables 4 and 5 are measured by the coefficient for changes in exchange rate. For both categories of automobiles they are negative in the German and Japanese equations. Thus, starting from an equilibrium, an appreciation of the German Mark (DM) or of the Japanese Yen leads to an increase in the corresponding price of the source country. Note however, that the speed of response varies. The response is immediate for German producers in both categories and for Japanese producers in the medium-size category. It is delayed by one period in the Japanese small-size category. In most instances the effect is small and sometimes it is weak (Japan, small-size category). At best, the short-run effect of an exchange rate shock is 50 percent and most coefficients are around 25 percent. German producers have a noticeably different rate of pass-through across categories. While it is only 23 percent in the small-size category, it is 51 percent in the medium-size category. Gagnon and Knetter (1995) identifies a similar behavior on the US market where the pass-through for large Mercedes is almost full and around 20 percent for smaller BMW.

These results reflect the normal case when the exchange rate changes. According to Section 2, the differences in price responses indicate that the nature of the import demand may be different for the small-size category and for the medium-size category.

In the pricing of small automobiles from Belgium, the exchange rate variable reflects the perverse effect arising from a low elasticity of import demand. However, it can also result from the particular nature of Belgian production (see footnote 11). In effect, an appreciation of the exchange rate between the Swiss and the Belgian Franc ($\Delta E^B<0$), *ceteris paribus*, necessarily implies that German producers are facing an appreciation of the Belgian Franc (BF) vis-a-vis the German currency (i.e., $E^B=(BF/DM)*E^G$ with E^G constant). Thus, costs originating in Germany (C^G which appears significantly in the equations for Belgium) are falling when measured in Belgian Francs and, as a consequence, producers in Belgium drop the price of automobiles exported to Switzerland. We tested this hypothesis by estimating the final complete pricing model with the German cost variable transformed in local currency in the Belgian equation. The exchange rate E^B looses all significance and the coefficient on C^G

hardly changes with increased significance supporting the role of German costs in Belgian Francs.[25]

While the positive exchange rate sign can be related to German owned production in Belgium, the negative sign in the French medium-size equation can only be attributed to the perverse effect arising from inelastic demand for imports.

Finally, unlike other studies of pass-through, we did not impose constraints on the cost elasticities. There is no clear evidence that the data would support the restriction of symmetric pass-through of exchange rate and costs as there is a significant difference between their respective coefficients. This suggests that prices may not be set in the buyer's currency (see Gross and Schmitt, 1996, and Gagnon and Knetter, 1995, for further evidence).

B. Price Rivalry

Consider now the short-term price rivalry among producers. A striking feature of the results is that pricing-rivalry is source-country specific and product-category specific. There are two clear cases of pair wise rivalry. In the small-size category, German and Japanese producers react to each other's price changes. Moreover, there is a remarkable symmetry in the size (around 25 percent) of the one-period lag response. In the medium-size category the dual rivalry is between French and Japanese producers. In this case almost 30 percent of the price change is taken into account by each rival although French producers need two quarters to adjust fully. In both instances, the coefficients are positive so that each producer has an incentive to follow its rival's price change. In terms of Section 2, prices are strategic complements (i.e., $\partial^2 V^k/\partial p_1^k \partial p_1^j > 0$, k=A,B; j≠k in (5)).

Interestingly, these results of pair wise rivalry between source countries are consistent with the evidence on substitution at the product level found by Berry, Levinsohn and Pakes (1995) (B-L-P, hereafter). In their study of the U.S. automobile market, they find that Mazda 323 and Nissan Sentra are close substitutes for Ford Escort and Chevy Cavalier (Opel Kadett for GM Germany). Our findings that Japanese and German producers compete directly on the small-size market are consistent with this product substitution since, in our sample, all the Nissan,

[25] More specifically, the Belgian pricing equation becomes,

$$\Delta P_t^B = -.166 \underset{(3.5)}{-.383\Delta P_{t-1}^B} \underset{(0.3)}{+.048\Delta E_{t-1}^B} \underset{(2.0)}{+.254\Delta P_{t-1}^G} \underset{(1.2)}{-.149 ecm_{t-1}} \underset{(2.7)}{+.514\Delta \hat{C}_{t-1}^G},$$

where $\Delta\hat{C}^G$ is German unit labor cost in Belgian Francs and absolute t-statistics are in parentheses.

Mazda, Ford models and 95 percent of the Opel models are in the small category.[26] Moreover, in 1988, these models had very similar listed price ranges (i.e., between SFr15,000 and SFr23,000 depending on the options). Hence, the models found to be substitutes in B-L-P compete in the same market niche in our sample.

In the medium-size category, the link between the B-L-P results and our sample is less direct because our results show pair wise rivalry between French and Japanese and B-L-P do not have French automobiles in their sample. However, they find that in the larger category, Honda Accord and Ford Taurus (Ford Sierra in Germany) are close substitutes. In 1988, the listed prices of these models were in the same brackets as those of two relatively popular French models, Peugeot 505 and Renault 25. Again, it appears that the models identified as likely substitutes in B-L-P and exported to Switzerland, compete not only in the same category but also in the same price range.

Finally, we note that in all cases but two, automobiles from different source-countries are strategic complements. Also, producers from Belgium do not seem to care about pricing rivalry. Their only reaction is in the small-size category and to German pricing. This is of course consistent with the fact that most of the Belgian production is by German owned firms. This conclusion is reinforced by the strong cost complementarity between Belgian and German production (i.e., cost variations in Germany are passed onto Belgian prices). There is no link, however, between German and Belgian prices in the medium-size category. The difference in the unit-values of automobiles produced in the two countries is also much larger in that category.[27] As it will become clear in the simulation results, the absence of rivals' prices in the Belgian equations is consistent with strategic pricing by Germany.

Finally, in all cases except Belgium the coefficient for the error-correction mechanism is significant. The speed of adjustment, however, is much larger in the small-size category than in the medium-size category. Note that given the way the cointegrating vector is specified, we expect a negative sign for Germany and a positive sign in all other cases.

C. Price Dynamics

To gain a better understanding of the dynamics of pricing as well as to assess the degree of pass-through in the long-run, we have run simulations for exchange rate variations in the two currencies for which the pass-through effect matters the most, the Japanese Yen and the

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[26] Price and weight for each model are given in *Automobile Revue*, Special annual issue published on the occasion of the 64th International Geneva Motor Show. Hallwag, Bern. March, 1988.

[27] The average unit-values, in Swiss Francs, over the sample are the following: For small-size automobiles P^B=12,996, P^F=11,154, P^G=14,182, P^J=10,176 and for medium-size automobiles, P^B=19,308, P^F=18,968, P^G=26,971, P^J=15,815.

German Mark. Two types of shocks are considered, a permanent and a temporary appreciation, both ex-post observations since we assume exchange rates follow a random walk. Consider first the impact of a permanent shock, namely a once for all change in exchange rate.

Figure 2 traces the period-by-period *cumulative* impact of a 10 percent change in currency for the price of each source country. Three results stand out. First, period 2 prices change in opposite directions. While the producer directly affected by the currency appreciation decreases its price, the rival producer increases it. Second, there is a systematic attempt by the producer directly affected by the appreciation to regain the market share lost with the initial shock. As a result, the long-term degree of exchange rate pass-through (that is, the price effect once all the adjustments have taken place) is lower than the short-term degree of pass-through. Third, rival prices may increase, but they may also decrease permanently.

The case of a permanent exchange rate shock is interesting because the impact on relative prices is rather different for a shock on the Yen and a shock on the German Mark. Moreover, the dynamics varies from one category to the other. The percentage changes in equilibrium prices are given in Table 6.

Table 6. Percentage changes in equilibrium prices
following 10 percent appreciation in exchange rates

	German Mark Appreciation		Japanese Yen Appreciation	
	Small	Medium	Small	Medium
Belgium	-0.11 percent	0.99 percent	-0.27 percent	-0.38 percent
France	0.55 percent	4.98 percent	0.30 percent	-1.52 percent
Germany	**0.75 percent**	**4.05 percent**	-0.25 percent	-0.39 percent
Japan	-0.32 percent	3.40 percent	**1.63 percent**	**0.54 percent**

A permanent appreciation in the Mark provokes an increase in most prices. In particular, German producers manage to induce their French rivals to almost match their increase in the small category (0.55 percent vs 0.75 percent) and to more than match it in the medium category (4.98 percent vs 4.05 percent). In both categories, automobiles priced in Belgium react very little thereby suggesting that German producers use their offshore production to control their rivals, in particular Japanese producers who do not react as much as the French. As indicated earlier, this behavior is consistent with the fact that Belgian-produced automobiles are in more direct competition with rivals than German-produced automobiles.

Figure 2. Cumulative Effect of 10 percent Appreciation

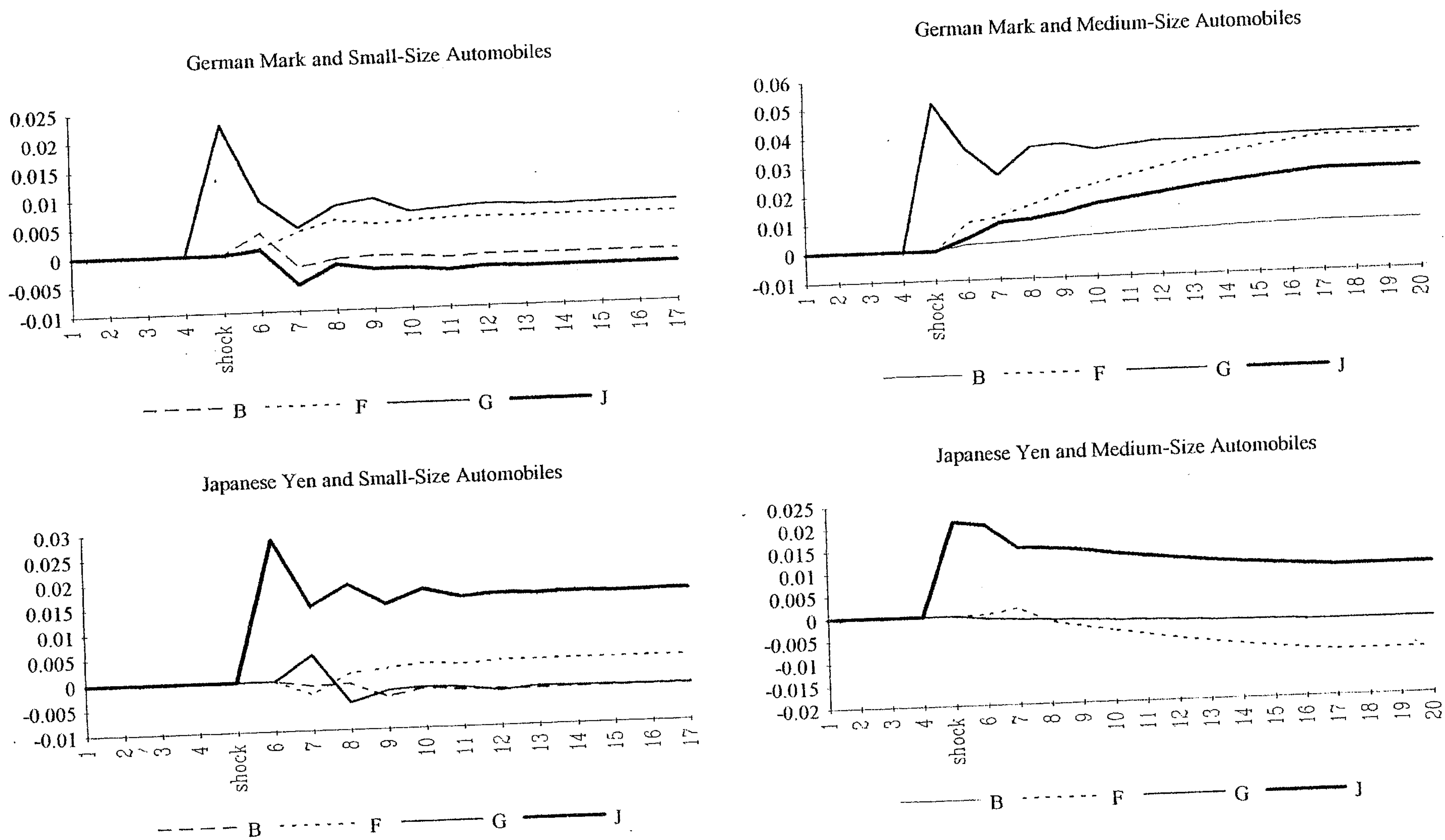
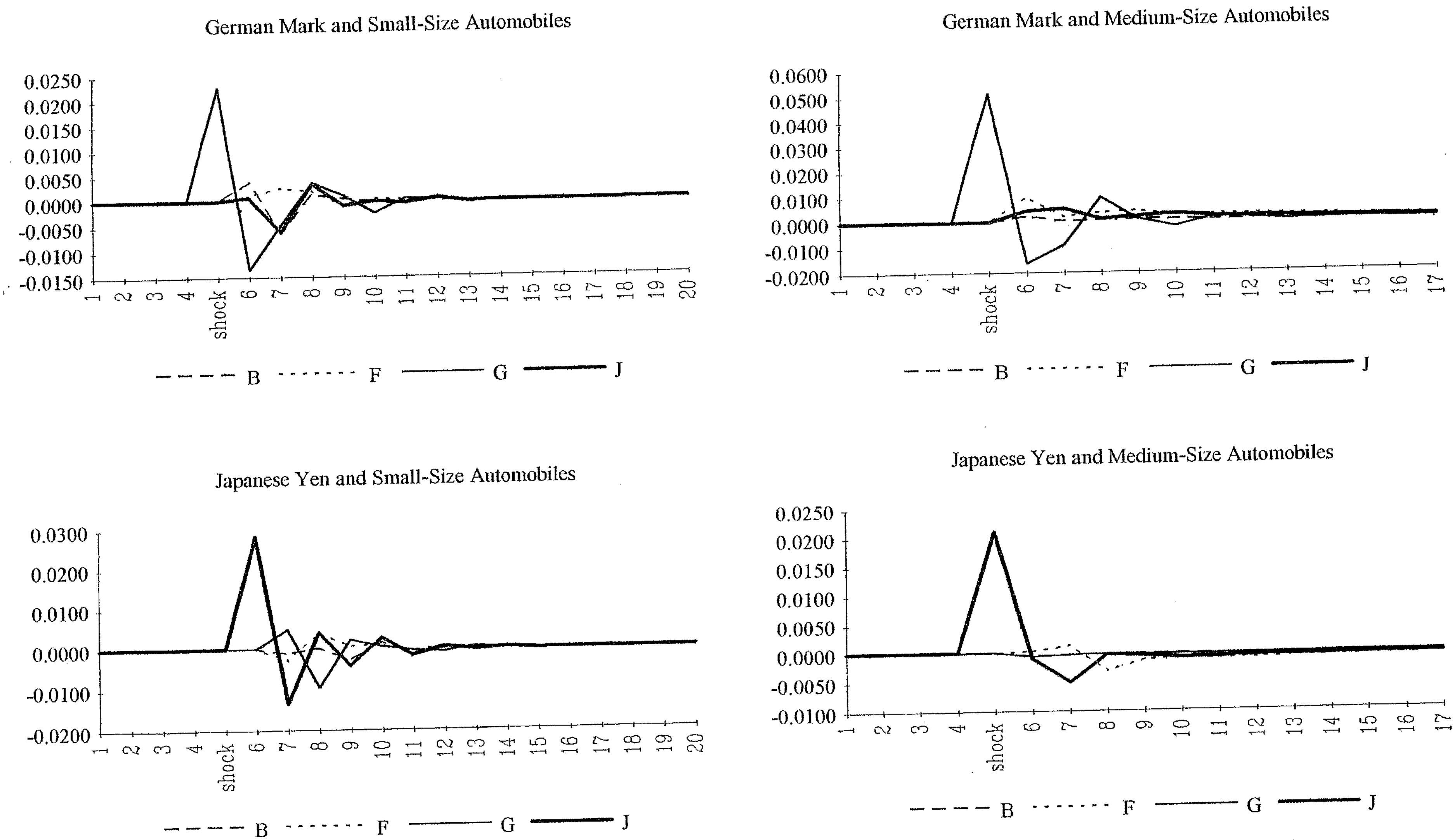


Figure 3. Period-to-Period Effect of 10 percent Temporary Appreciation



An appreciation of the Yen hardly leads to any reaction by competitors of Japanese products, immediately or in the long-run. Thus, Japanese producers are forced to decrease their price from the initial impact causing rivals to follow suit. As a result, Japanese producers have difficulties inducing their rivals to match their increase in prices in either category. In fact in all but one case, competitors lower their prices. The exception is French producers in the small category, who increase their price by 0.30 percent while the exchange rate shock raises the Japanese price by 1.63 percent. Note that the convergence toward equilibrium prices is much slower in the medium-size than in the small-size since the ecm coefficients are very small.

Consider now the impact of a temporary shock namely, a single period change followed by a reversal of the change in the next period. Figure 3 illustrates the period-by-period dynamics of the price changes of the four source countries following a temporary appreciation of the Japanese Yen and of the German Mark.

For both categories and independently of which currency appreciates, the second-period reversal of the first-period exchange rate appreciation leads to a price overshooting. This indicates that the country directly affected by the first-period appreciation corrects its price in order to regain market shares lost in the first period. The remaining oscillatory movements are driven by the negative autocorrelation in the first difference in prices. In the medium-size category, complete stability in prices is reached after 20 periods while it is reached after 12 periods in the small-size category. Note that temporary shocks have only temporary effects on prices and therefore the mean of the series is zero.

Clearly, there is more price interdependence in the small-size than in the medium-size category. To measure the relative variability of the series following a shock, we have computed the ratio of the variance of each rival's price series against the variance of the prices of the producer directly affected by the shock. The maximum relative variability in the small-size category is against Belgium for the Mark appreciation [Var(B)/Var(G)=0.074] and against Germany for the Yen appreciation [Var(G)/Var(J)=0.123]. In the medium-size category, the maximum relative variation is 0.021 for Mark appreciation and 0.025 for Yen appreciation against France in both cases. Thus, competition is stronger in the small-size category than in the medium-size one and, variations in the Yen lead to more turbulence in prices than variations in the Mark. Nevertheless, the overall variability remains small for temporary shocks. This can be attributed to two factors: The small degree of price interdependence in the short run and the low adjustment coefficient of the ecm term. Still the pair wise competition between Japanese and French producers in the medium-size category and especially the pair wise competition between Japanese and German producers in the small-size category are quite apparent.

Finally, the difference between the short-run and the long-run degree of pass-through is striking. Some have suggested this is consistent, among other things, with invoicing in the buyer's currency (Gagnon and Knetter, 1995). We find the result is also consistent with rivalry between producers which can generate the following general pattern. Suppose an appreciation of its currency increases Firm A's price in period 1. Whether or not the rival producer's reacts immediately to this price change, Firm A's price increases more than Firm B's. As a result, Firm A loses market share. In period 2, Firm A's reacts by decreasing its price while Firm B goes in the opposite direction and increases its price. Firm B can afford to increase its price because it has gained market share in the first period. Firms then continue to react to each other's price until convergence to the initial (for temporary shock) or to the new (for permanent shock) long-run equilibrium.

The dynamic adjustments just described above also result in a generally lower long-run degree of pass-through than what is found in analyses without endogenous rivalry. Our findings can be compared to two other studies which estimate long-run pass-through without simultaneous pricing: Gagnon and Knetter (1995) which studies pricing by Japanese and German producers in a variety of destination markets and Gross and Schmitt (1996) which studies the pricing on the Swiss market with a dataset similar to ours. Gagnon and Knetter find that the long-run degree of pass-through by Japanese producers is zero for small automobiles on the Swiss market and around 20 percent for larger automobiles across most other destination markets. German producers exhibit approximately 50 percent pass-through across various destinations. Gross and Schmitt find a zero pass-through for German automobiles while it is 37 percent and 67 percent for small- and medium-size Japanese automobiles. The long-run degrees of pass-through obtained in this study are equal or lower than those found in these two papers.

VI. CONCLUSION

In this paper we have investigated the role of rivalry on pricing in general and on the exchange rate pass-through in particular. We chose the Swiss market for automobiles as the import market because there is no dominant producer and it is consistent with an oligopoly setting with four source-countries covering a large fraction of the market. Another important feature of this market is the absence of significant distortionary trade barriers against Japanese automobiles.

The main message of this paper is that price interdependence matters and that exchange rate variations can have significant feedback effects on rivals' prices leading to smaller pass-through in the long-run than in the short-run. Using recently developed econometric techniques for system estimations when series are integrated, we have shown that rivals take this inter-dependence into account in two ways. First, we have uncovered a stable long-run relationship among prices which is consistent with long-run equilibrium oligopolistic prices. Second, in the short-run, we have found that rivalry is source-country specific and category-specific. In particular, two clear cases of pair-wise rivalry have been found: Japan and Germany in the small-size automobile market, and Japan and France in the medium-size automobile market. Both give rise to significant changes in prices whether the shock is temporary or permanent. For instance, we found that some rivals match permanent price changes.

As for the degree of exchange rate pass-through, we discovered that, in short run, it is relatively small since it is always less than 50 percent. With permanent changes in exchange rates, the producer directly affected by exchange rate variations absorbs an increasing proportion of the exchange rate change through time because the main rivals maintain or even decrease their price. As a consequence, the long-run degree of pass-through is very low. These results are remarkable given the fact that the price series are unit-values for country-specific sources and not product-specific prices.

This study shows that the time-series approach to pricing can be a useful tool to investigate price interdependence in general. The analysis of exchange rate changes offers the advantage of frequent and relatively large variations compared to other price components. Hence, our approach may contribute to a better understanding of the role of rivalry in particular and market structures in general.

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