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PRICING TO MARKET IN INTERNATIONAL TRADE: EVIDENCE FROM PANEL DATA ON AUTOMOBILES AND TOTAL MERCHANDISE

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ABSTRACT

This paper focuses on price discrimination in international trade that is associated with movements in exchange rates. This phenomenon is referred to as "pricing to market." We find strong evidence of pricing to market for Japanese exports of automobiles. We find moderate evidence of such behavior for German auto exports, and very little pricing to market for U.S. auto exports. We conjecture that these sharp differences in export pricing behavior may be due to differences in the extent of overseas production by firms based in these countries. Pricing to market may be more important to firms that do not have plants in their target markets.

The patterns observed for automobiles do not hold up for total merchandise exports, where pricing to market varies by both source and destination country. These differences in measured pricing to market may reflect differences in the product mix of trade by source and destination.

Pricing to Market in International Trade:

Evidence from Panel Data on Automobiles and Total Merchandise

Joseph E. Gagnon and Michael M. Knetter[1]

With the extreme fluctuations in currency values since the collapse of the Bretton Woods agreement, firms based in different countries have faced unprecedented shocks to their relative costs of production. In spite of this, it is widely observed that import prices (prices of foreign produced goods in domestic currency) in the United States move very little compared to movements in exchange rates. While this observation is in principle consistent with two quite different models--globally competitive markets in which the United States is a large country and segmented international markets with price discrimination--existing research strongly suggests that these observations are due to price discrimination.[2] Krugman

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1. Board of Governors of the Federal Reserve System, and Dartmouth College, respectively. Gagnon is on leave at the University of California, Berkeley during the academic year 1990-91. Knetter is on leave at Wissenschaftszentrum Berlin for the Fall semester of 1990. This paper represents the views of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or other members of its staff.

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2. See, for example, Knetter (1989) or Marston (1989).

(1987) has referred to such price discrimination, triggered by exchange rate movements, as "pricing to market."

The alternative to pricing-to-market (henceforth PTM) could be thought of as the law of one price: exporters charge identical common currency prices to all buyers. Even if the law of one price does not hold exactly, there may be limits to the extent of price discrimination due to the opportunity for profitable arbitrage. Moreover, arbitrage pressures might grow over time in response to large deviations in sales prices across markets, thus tending to enforce the law of one price in the long run.

The primary focus of this paper is to analyze both long run and short run aspects of PTM behavior using panel data on export unit values in specific categories of automobiles. Although the model developed in this paper is most appropriate for individual differentiated products, we also estimate the model using unit values for total merchandise exports. We consider two alternatives for long run pricing behavior: one which allows price discrimination and one which imposes the law of one price. For each long run model, we then estimate the short run dynamics of export prices. While it is not possible to conduct formal hypothesis tests for the alternative long run models, it is possible to make less formal comparisons of them on the basis of their goodness of fit as well as their implications for short run dynamics.[3] Comparison of the short run and long run behavior also provides indirect evidence on the nature and importance of adjustment costs and currency contracts in international trade.

3. Formal tests are precluded since the variables in the long run regression model are nonstationary, which implies the coefficient estimates have non-standard distributions.

Apart from measuring short run and long run pricing to market, the data sets provide new information on the pattern of PTM. The unit value panels vary by source country and destination country, which allows us to make a number of interesting comparisons of PTM behavior. By observing source and destination effects on pricing behavior, we gain more insight into which economic explanations for PTM are most credible. Finally, because the data sets vary in level of aggregation, we get some idea whether significant results are obscured by averaging over categories of goods.

We find that PTM is pervasive in Japanese auto exports, present for some destinations and categories of German auto exports, and virtually absent from U.S. auto exports. The results are quite robust to alternative specifications and generally are not sensitive to the sample period chosen. In particular, the evidence for PTM in Japanese auto exports to the United States is almost as strong before the imposition of voluntary export restraints as it is after their imposition. There is some evidence of a change in the PTM behavior of Japanese auto exports to Europe, however.

The results are notably different at the aggregate level. For total merchandise exports, the Japanese do not appear to engage in PTM any more than other exporters. Indeed, aggregate U.S. exports to Japan have a higher estimated degree of PTM than aggregate Japanese exports to the United States. Due to aggregation we cannot tell whether this difference in observed PTM behavior reflects differing behavior of U.S. and Japanese firms in each industry, or a different product mix of exports by industry, or both.

Our second main finding is that short run PTM is typically less than long run PTM, indicating that export prices are sticky in the exporter's currency, although there are a few interesting exceptions. This finding would be consistent with export invoicing in the exporter's currency.

The plan of this paper is as follows. In Section 1, we review the price discriminating monopoly model that motivates the empirical model of pricing to market for the cross section data set. Section 2 will discuss the empirical specification of the model and the data sets used in this study. Section 3 covers the estimation of long run and short run PTM behavior. Section 4 explores the impact of aggregation on the results. Section 5 concludes the paper. The appendix shows how the error correction model used to estimate short run PTM may be derived from a model with convex adjustment costs in trade flows.

1. Theory

The motivation for the empirical research to follow can be shown most simply in the context of a price discriminating monopolist.[4] We thus begin with the assumption of segmented markets, although the model does allow integrated world markets as a special case (when demand elasticities are identical and infinite in each destination). The model is partial equilibrium and for simplicity we abstract from adjustment costs and lags between production and sales.[5] Consequently, the theory emphasizes the long run equilibrium relationship we would expect between export prices, costs, and exchange rates for a firm selling to a number of segmented markets in which it may face downward sloping demand schedules. A benefit of

4. This section is based on Knetter (1990).

5. If adjustment costs are important, they should be captured in the error correction equations that are estimated to pick up short run dynamics. The appendix demonstrates how an adjustment cost model of trade can lead to an error correction reduced form for export prices.

the multi-market model is that it allows us to relax assumptions about the cost function that are often found in bilateral pass-through models. In particular, *we make no assumptions about the slope of marginal cost or the effect of exchange rate changes on the cost function*--all that is required is that marginal cost is common across destinations.

Consider a firm that produces goods for sale in *n* separate destination markets. The profits of the firm are given by:

$$(1.1) \qquad \Pi(p_1,\ldots,p_n) = \sum_{i=1}^{n} p_i q_i(e_i p_i) - C\left(\sum_{i=1}^{n} q_i(e_i p_i),\ w\right)$$

where *p* is the export price (i.e., price in the seller's currency), *q* is the quantity demanded which is a function of the price in units of the buyer's currency, *e* is the exchange rate (destination market currency per unit of the seller's currency), *w* is the input price, and $C(\Sigma q,w)$ is the firm's cost function. The first order conditions for profit maximization are simply:

$$(1.2) \qquad \frac{\partial \Pi}{\partial p_i} = p_i q_i'(e_i p_i)e_i + q_i(e_i p_i) - MCq_i'(e_i p_i)e_i = 0;\ i=1,\ldots,n$$

MC is the derivative of the cost function with respect to total quantity, *i.e.*, marginal cost. The arguments of *MC*, total quantity and input price, have been suppressed for simplicity. Manipulation permits (1.2) to be written as the familiar condition that the firm equates the marginal revenue from sales in each market to the common marginal cost. Alternatively, export price to each destination is the product of the common marginal cost and a destination-specific markup:

$$p_i = MC\left(\frac{\eta_i}{\eta_i - 1}\right); \ i=1,\ldots,n \tag{1.3}$$

where η is the absolute value of the elasticity of demand in the foreign market with respect to changes in price. A change in the exchange rate vis-a-vis the currency of country i can affect the price charged to market i in two ways: by affecting marginal cost (through changes in quantity or input price) or the elasticity of import demand. The former effect will spillover to the other destination markets as well, while the latter is idiosyncratic.

These two effects are revealed more clearly by taking the log of (1.3) and totally differentiating the resulting expression with respect to input prices, output prices and exchange rates:

$$dlnp_i = \frac{\partial lnMC}{\partial \Sigma q_j}\left(\Sigma q_j'(p_j de_j + e_j dp_j)\right) + \frac{\partial lnMC}{\partial w}dw - \frac{\eta_i' e_i p_i}{\eta_i(\eta_i - 1)}\left(dln(e_i p_i)\right), \tag{1.4}$$

where the arguments of q' and η' are suppressed and the relation holds for i=1,...,n. Defining:

$$\beta_i = \frac{\dfrac{\partial ln\eta_i}{\partial lnp_i^*}}{(\eta_i - 1) + \dfrac{\partial ln\eta_i}{\partial lnp_i^*}}$$

where $p^* = ep$ is the price in the buyer's currency, enables equation (1.4) to be expressed as:

$$dlnp_i = (1-\beta_i)\ dlnMC - \beta_i\ dlne_i; \ i=1,\ldots,n \tag{1.5}$$

where the term *dlnMC* refers to the *total* change in marginal cost due to both input prices and output volume.

Several important observations can be made about export price behavior on the basis of (1.5). Constant elasticity of demand implies that β equals zero. In that case, export prices change one for one with changes in marginal cost and are invariant to movements along the demand curve due to changes in the exchange rate. In terms of import prices (prices in units of the buyer's currency) marginal cost and exchange rate changes have a symmetric effect--price changes in proportion to both. This may be shown by rewriting (1.5) in terms of import prices:

$$(1.6) \qquad (dlnp_i + dlne_i) = (1-\beta_i)\ dlnMC + (1-\beta_i)\ dlne_i;\ i=1,\ldots,n$$

Other things equal, a 10% increase in marginal cost should leave the exporter in the same position as a 10% exchange rate appreciation with respect to any particular market. The symmetry result holds independently of the shape of the demand schedule as given by β. If demand is less convex than constant elasticity, then β is greater than zero (i.e., markups of price over cost fall with an increase in cost), while if demand is more convex than constant elasticity, β is less than zero.

We believe that demand curves with less convexity than constant elasticity are more plausible than demand curves with greater convexity. Furthermore, the phenomenon of PTM, as described by Krugman, implicitly assumes that demand curves are less convex than constant elasticity. In the empirical results of this paper, we consider estimates of β near 1 to be evidence of complete PTM and estimates of β near 0 to be evidence of no PTM. Values of β greater than 1 are theoretically impossible, since the monopolist must operate where the demand elasticity exceeds 1. Theory does not rule out cases in which β is less than 0, however.

It is important to remember that measures of PTM are not strictly related to measures of the "pass-through" of exchange rates to import prices. Pass-through typically refers to the overall effect of an exchange rate change on a country's import prices. The pass-through of an exchange rate change might be incomplete either because of PTM by foreign producers or because the foreign producers' marginal costs were affected by the exchange rate. In this paper we consider only the former effect.

Before turning to the empirical framework, a few qualifications are in order. First, the monopoly model is certainly an oversimplification of most real markets and therefore the interpretation of β as revealing the convexity of market demand is dubious. The elasticity of export price with respect to exchange rates and marginal cost is likely to depend on characteristics of market demand and the behavior of other firms in the industry. Nonetheless, the empirical specification that follows from the monopoly model may be reasonable. Baker and Bresnahan (1988) find no evidence that oligopoly solution concepts are unstable in their study of the U.S. domestic beer industry. Consequently, even though exporters may face competition within their own country or from producers in other countries, their residual, or perceived, demand curves may in fact be stable. In that case, we would interpret β as revealing information about the convexity of residual demand.

A second qualification is that the analysis that leads to equation (1.6) presumes that price adjustment is instantaneous and costless. To the extent that there are inherent lags and costs in the adjustment process, the relationship in equation (1.6) must be considered a long run equilibrium relationship. Whether it is relevant in the short run will be tested empirically.

Finally, there is no theoretical argument to support the assertion that β is constant over time for general demand functions. (β is constant (at zero) for the class of constant elasticity demand functions, however.) One reason for concern about the constancy of β in the long run is the possibility of arbitrage across markets. If a large deviation in the export price charged to two different markets eventually induced arbitrage across the two markets, the exporter's perceived demand curve could well change in a way that would not leave β constant.

2. Specification and Data

Equation (1.5) describes the optimal price response of the exporter to deviations in the marginal cost and the exchange rate from an initial equilibrium. If the initial equilibrium is taken to be an arbitrary constant for each variable, a natural regression relationship can be obtained by writing equation (1.5) in levels of the variables with an intercept term.

$$(2.1) \qquad \ln p_{it} = \mu_i + (1-\beta_i) \ln MC_t - \beta_i \ln e_{it}$$

Equation (2.1) was estimated using annual export unit values for selected 7-digit categories within the automobile industry as well as total merchandise trade. There are three source countries for auto exports: the United States, Japan, and Germany. For each of these source countries, we have destination-specific f.o.b. values and quantities of exports to several major destinations. The data are taken from government publications of the respective source countries and are typically collected by customs agents in those countries.

There are five source countries for total merchandise trade: Canada, the United States, Japan, Germany, and the United Kingdom. These data are obtained from a tape compiled by the European Economic Community, which is in turn derived from UN and OECD sources.

The exchange rates are annual average spot exchange rates divided by the wholesale price index in each destination market. The rationale for dividing by foreign price levels is that the foreign demand curve, $q(ep)$, is presumably a function of a real price rather than a nominal price. During the sample periods we study there was tremendous variation in exchange rates, which ought to enable us to identify the extent of PTM very precisely.

The marginal costs are not observed directly and no attempt was made to proxy for them with observable series. Rather, the estimation strategy takes advantage of the cross-sectional nature of the available data on export prices. The marginal cost in each time period for each source country is simply estimated (up to a constant scale factor) by the common component in export prices across different destination markets. In other words, a dummy variable is created for each time period and a separate coefficient is estimated for each dummy variable. These "time effects" coefficients are constrained to be identical across the different destination markets for a given source country. Since there is more than one bilateral export price in each time period, the time effects do not exhaust all the degrees of freedom. The advantage of this approach is that it makes the minimum necessary assumptions. The only necessary assumption is that for a given exporter the marginal cost of exporting to different markets is identical.

There are two drawbacks to this estimation strategy. First, it uses up many degrees of freedom. Second, any change in markup that is common across destination markets will be captured by the time effect, so that the time effects may capture more than just marginal cost movements.

3. Estimation and Results

Estimation proceeds by stacking equation (2.1) as follows:

$$(3.1)\qquad \ln p_{1t} = (1-\beta_1)\,\underset{\sim}{\theta}'\,D_t - \beta_1 \ln e_{1t} + u_{1t}$$

$$\ln p_{2t} = \mu_2 + (1-\beta_2)\,\underset{\sim}{\theta}'\,D_t - \beta_2 \ln e_{2t} + u_{2t}$$

$$\vdots$$

$$\ln p_{nt} = \mu_n + (1-\beta_n)\,\underset{\sim}{\theta}'\,D_t - \beta_n \ln e_{nt} + u_{nt}$$

$\underset{\sim}{\theta}$ is a vector of coefficients that controls for effects that vary over time but are the same for all destinations. In terms of the model of section 1, $\underset{\sim}{\theta}$ represents movements in marginal cost, but in a more general oligopoly model it may include changes in industry conduct. D_t is a dummy vector equal to 1 in the t^{th} position and 0 elsewhere. If T is the number of observations in the sample, then there are T dummies, D_t, each of which is a T-vector. $\underset{\sim}{\theta}$ and D_t are defined as follows:

$$\underset{\sim}{\theta}' = [\theta_1 \; \theta_2 \; \ldots \; \theta_T] \qquad \begin{aligned} D_1' &= [1 \; 0 \; \ldots . \; 0] \\ D_2' &= [0 \; 1 \; \ldots . \; 0] \\ &\vdots \\ D_T' &= [0 \; 0 \; \ldots . \; 1] \end{aligned}$$

Equations (3.1) were estimated simultaneously using a Gauss-Newton procedure to minimize the total sum of squared residuals. An intercept term was estimated for each destination to control for factors that are constant over time but differ by country. This term should identify differences in the average quality of goods shipped to different markets as well as differences in the average markups to different markets that do not vary with exchange rates. Because of the complete set of time dummies, the intercept term had to be dropped from one equation. The average level of the θ's thus captures the average quality and markup characteristics of the first destination market.

Since the data are clearly nonstationary, we cannot use standard significance tests on the coefficient estimates from equations (3.1).[6] Moreover, standard cointegration tests do not apply to equations (3.1) because the number of estimated coefficients is of the same order as the

6. Standard Dickey-Fuller and augmented Dickey-Fuller tests were never able to reject a unit root in any of our series. A second unit root was rejected at the 5 percent level in about one-third of the export price and exchange rate series. Given the low power of these tests, we take the evidence as favorable to the hypothesis that all series are integrated of order one.

number of observations.[7] Nevertheless, we would like to treat equations (3.1) as a set of cointegrating relationships, with the time effects taking the role of a generated series with which export prices are cointegrated. We then use the results of the cointegrating regression to estimate an error correction model of export prices. The error correction model provides information on the importance and nature of adjustment costs as well as the appropriateness of the presumed cointegrating relationship.

Equations (3.1) were estimated with the β's unconstrained and under the constraint that $\beta=0$. The constrained regressions were run because of the possibility that arbitrage might work over long horizons to ensure that the law of one price holds. We cannot formally test the restriction that $\beta=0$, but the goodness of fit of both the cointegrating and error correction regressions will provide some information on the two hypotheses.

The general error correction model considered in this paper is given by equation (3.2).[8]

$$(3.2) \qquad \Delta \ln p_{it} = \alpha_i^0 + \alpha_i^1 \Delta \ln p_{it-1} - \alpha_i^2 \Delta \ln e_{it} + \left(1-\alpha_i^2-\alpha_i^3\right)\left(\hat{\theta}_t - \hat{\theta}_{t-1}\right) - \alpha_i^4 \hat{u}_{it-1} + e_{it}$$

Equation (3.2) is regressed using $\hat{\theta}$ and $\hat{u}$ from the results of equation (3.1), where $\hat{\theta}$ is simply the estimated coefficient vector, $\underset{\sim}{\theta}$,

7. It is possible that reasonable asymptotic properties may be obtained by considering the distribution of the estimators as both the number of time periods and the number of destination markets grow. Such analysis is beyond the scope of this paper.

8. Kasa (1989) motivates an error correction representation of export prices with a dynamic optimizing model of trader behavior.

converted into a time series. Since $\hat{\theta}$ and $\hat{u}$ are fixed series in the second regression, a complete set of coefficients may be estimated for every destination market. However, to conserve degrees of freedom, we looked for parameters that could be constrained to zero for every source and destination. In the vast majority of cases, α^0 and α^1 were insignificantly different from zero, so the results focus on the case in which α^0 and α^1 are constrained at zero.

The Automobile Industry

The results for automobile exports are in Tables 1-9. For each of the three source countries--Germany, Japan, and the United States--two long run equations are estimated, one which allows pricing to market and one which imposes the law of one price, *i.e.*, one which constrains β=0. The estimation period is 1973-87 for Japanese and U.S. exports, and 1975-87 for German exports. The first three tables present the unrestricted estimates of long run PTM (β) for each country in turn. The following six tables present the results of the error correction regressions that capture the short run dynamics of export pricing under each of the alternative assumptions about the long run equilibrium.

Table 1 shows the unconstrained estimates of β for equations (3.1) for German exports of autos in three engine sizes--1500-1999 cubic centimeters, 2000-2999 cc, and 3000 cc and over--to six destination markets--Canada, Japan, the United Kingdom, the United States, France, and Sweden. Somewhat surprisingly, the estimated PTM in the long run is most pervasive in the two smaller engine size categories. For exports of small cars to the United States and France, the estimated coefficients are .59 and .52, respectively. The implication is that a 20% depreciation of the dollar, all else constant, would elicit a 12% reduction in the DM price charged to U.S. buyers. The dollar price would rise by only 8%. The DM price faced by other buyers would remain unchanged provided their exchange rates had not changed. The extent of pricing to market is less pronounced for other destinations. In the second category, the estimated magnitude of pricing to market is about 40% for France, Sweden, and the United Kingdom. It is 90% for Japan, and close to zero for the United States and Canada.

For the largest auto category, the point estimates of the PTM coefficient have the perverse sign. For the United States, for example, the estimate of β implies that all else equal, a 20% depreciation of the dollar leads to a 3% increase in the DM price charged to U.S. buyers. Thus, the dollar price rises by 23%--even more than the depreciation itself. For France and Sweden, the effect is even more perverse.

The pattern of export pricing for Japanese auto exports is much clearer. (See Table 2.) All but one coefficient estimate imply that price adjustment will have a stabilizing effect on the price in the buyer's currency in the face of an exchange rate shock. The only exception is small car exports to the United States.[9] The overwhelming majority of the point estimates suggest that 80% to 90% of the impact of exchange rate changes are offset by adjustment of the yen export price for all destinations. In the absence of any change in marginal cost, a 20% appreciation of the yen against the dollar would increase dollar prices by only 4% for autos over one liter in engine size.

9. It should be noted that this category represents a trivial share of U.S. auto imports. In fact, the number of autos exported to the United States never exceeded 16,000 for this category until 1984. In 1987, this category accounted for only 7.6% of total U.S. imports from Japan. See Table 12 for complete information on the quantity shares.

For U.S. auto exports, the long run price adjustment pattern is equally clear--although much different in character. (See Table 3.) The correlation between destination-specific price movements and exchange rates is virtually zero for all destination markets. Export price adjustment appears as likely to increase the variability of price in the buyer's currency as to reduce it. Over half of the estimates lie between -.05 and .05. This stands in stark contrast to the German and Japanese auto exporters' behavior.[10]

We were concerned about the possibility of structural change in these relationships, particularly due to the imposition of voluntary export restraints (VERs) on Japanese auto exports to Canada and the United States in 1981.[11] Due to nonstationarity of the data, we could not run a standard Chow test. However, we did estimate equation (3.1) for Japanese auto

10. We were concerned that low estimates of PTM could be due to a transfer pricing problem. In other words, automakers might ship their cars to a foreign subsidiary at a constant price, and the foreign subsidiary might stabilize the price paid by independent dealers in foreign currency. As far as the firm is concerned, this behavior is pricing to market, but export price data would not identify it as such.

We spoke to executives at General Motors, Daimler-Benz, and Volkswagen. General Motors stated that its domestic exports are shipped directly to independent dealers in foreign countries, except for shipments to Canada, so that transfer pricing is not a problem. Daimler-Benz and Volkswagen both claimed that the prices charged to their American subsidiaries are sensitive to exchange rate movements in an attempt to stabilize the profits of the American subsidiaries. While further PTM might occur between these subsidiaries and their franchised dealers, we find some support for our finding of low PTM in the marked decline in the volume of large German cars exported to the United States over the past few years.

11. The Japanese auto industry has agreed to an informal restriction that limits the Japanese share of the U.K. auto market to 11 percent. This restriction was in place throughout the sample, although we do not know whether it was binding throughout the sample. There has been a binding restriction on Japanese auto exports to Australia throughout the sample. There were no restrictions on Japanese auto exports to the remaining destinations in our sample, and there were no restrictions on U.S. and German auto exports to any of our destinations.

exports over two subsamples, 1973-80 and 1981-87. The results are presented in Table 2A. One would expect that binding quantitative restrictions would be associated with a stable price in the destination market, and hence, a high degree of observed PTM. Indeed, the measure of PTM in Table 2A does tend to rise after 1980. For the United States and Canada, which imposed VERs on Japanese autos, the estimated increase in PTM is modest. For Germany and the United Kingdom, the apparent increase in PTM is striking. We are puzzled by this result. There was no substantial evidence of structural change by U.S. and German exporters over this sample. Further evidence of the constancy of the relationship in equation (3.1) is obtained from the error correction regressions which are presented later. (A large and significant estimate of the error correction coefficient is indicative of a stable relationship in equation (3.1).)

Figures 1-8 provide convincing evidence of the differences in PTM for Japanese exports of autos and U.S. exports of autos. The evidence also shows that the measured export unit values behave quite sensibly given the inflation and exchange rate movements of this period. Figure 1 plots the log of the unit values of Japanese exports of autos between 1000 and 2000 cc to the United States (USP) and Germany (WGP) as well as the estimated time effect (THETA) from the regression of equation (3.1). The time effects behave very much as expected, with their change over time closely approximated by an average of the price changes. The evidence of PTM is quite clear during the 1980s in this figure. The unit value of shipments to the United States rises much more rapidly than the German counterpart during dollar appreciation. Then when the Deutschemark strengthens against the dollar from 1985 onward, German unit values rise abruptly and U.S. unit values fall.

Further evidence can be seen in Figures 2 and 3, which plot the unit values to each destination (USPRICE and WGPRICE, after subtracting their means) against the price-level-adjusted exchange rates (USX and WGX, also net of their means). A fall in the exchange rate series of 0.1 means that with no change in the yen export price, the dollar price would fall by 10%. It is quite clear in Figure 2 that the unit value to the United States is negatively correlated with this exchange rate series. The unit value rises rapidly during periods of dollar appreciation and actually falls during the dollar depreciation of 1986-87. Figure 3 also shows how unit values to Germany rise most when the DM is appreciating and vice-versa. Figure 4 plots the adjusted exchange rate series for several destination markets. The main message of the figure is that the ability of the data to identify PTM is greatest in the 1980s, when there are divergent movements in several of the series.

Figures 5-8 are the corresponding evidence for unit values of U.S. shipments of autos under 6 cylinders to Canada and the United Kingdom (CNP and UKP). The unit value series grow together quite closely, which leaves little scope for PTM. The time effects are centered around the Canadian unit values since Canada is the country without a fixed effect in the regression. Figure 8, which plots the price-level-adjusted exchange rate for several U.S. destination markets shows that the best chance of identifying PTM is during the two swings in the dollar/pound rate in the 1980s. Figure 7 shows there to be very little effect on the upward trend of unit values of shipments to the United Kingdom during either the fall of the pound between 1980 and 1985, or its subsequent rise. Figure 6 plots the rise of unit values to Canada, which proceeds quite steadily, with little apparent relationship to the exchange rate, especially in the 1980s.

Figure 9 illustrates one apparent problem encountered in estimation of the long run model for Japanese exports of large cars. It plots the unit value of exports of cars to the United States and Germany, net of their means, and the time effects net of their mean. The estimated time effects grew at a much faster rate than any of the unit value series. Consequently, it seemed possible that this behavior could account for the high degree of measured PTM, as well as the small amount of variation in PTM by destination. To check the robustness of our results, we estimated a linear model in which the PTM parameter, β, does not interact with the time effects, $\underset{\sim}{\theta}$. (This model is not characterized by symmetry between the effects of cost shocks and exchange rate shocks on export prices in the importer's currency.)

$$
\begin{aligned}
(3.3) \qquad \ln p_{1t} &= \underset{\sim}{\theta}' D_t - \beta_1 \ln e_{1t} + u_{1t} \\
\ln p_{2t} &= \mu_2 + \underset{\sim}{\theta}' D_t - \beta_2 \ln e_{2t} + u_{2t} \\
&\vdots \\
\ln p_{nt} &= \mu_n + \underset{\sim}{\theta}' D_t - \beta_n \ln e_{nt} + u_{nt}
\end{aligned}
$$

The estimated values of β for equation (3.3) were about 1.0 for the United States, Canada, and Germany and about 0.7 for Norway and the United Kingdom. The time effects from this model, net of their mean, are plotted against the de-meaned unit values series for the United States and Germany in Figure 10. The results look more reasonable than those in Figure 9, and the estimated amount of PTM does display more variation. Figures 11 and 12 show the unit value and exchange rate movements for the United States and Germany, respectively. The patterns again are remarkably clear.

We next consider the two sets of error correction results for each of the exporters. Three coefficients are reported for each category of autos. The first gives the short run response of price to exchange rate changes. Comparing it with the long run response reveals whether short run price adjustment is greater or less than long run. The second coefficient estimates the difference in the short run between the effect of changes in the exchange rate and the effect of changes in the estimated time effects. Recall that the theory in the first section of the paper shows that these effects should by symmetric, provided that the time effects are a good measure of marginal cost and all changes are viewed as permanent. The final coefficient measures the response of the export unit value to a deviation from its long run equilibrium value in the previous period. A coefficient of 1 means that any error last period is completely corrected in the current period. An estimate of .5 means that half of last period's deviation is corrected this period.

Table 4 gives the parameter estimates for German autos using the long run equation that allows β to be non-zero. Table 5 is based on the long run equation with β=0. Table 4 (LR betas unconstrained) shows that for Japan, the United Kingdom, France, and Sweden it appears that there is much less evidence of PTM in the short run than in the long run. For Canada and the United States, the opposite seems to hold--export prices seem to overreact in the short run. This would be consistent with invoicing in the buyer's currency for sales to the United States and Canada and invoicing in

Deutschemarks otherwise.[12] The error correction coefficients all have the expected sign, the magnitudes look plausible, and the standard errors are small. The estimated error correction coefficients thus lend some support to the hypothesis of a stable long run relationship in equation (3.1), although an exact statistical test is not possible. Symmetry of the short run responses to exchange rates and marginal costs appears to be rejected in most cases. Departures from symmetric responses are of an unusual character. Short run PTM appears to be more vigorous with respect to changes in marginal cost as captured by the time effects. That is, the sign of α^3 is typically the same as the sign of α^2.

When the law of one price is assumed to characterize export prices in the long run (β=0), there is much less evidence of PTM in the short run as well. Only for exports to the United States do we see a tendency for short term price adjustment to mitigate the impact of exchange rate changes on dollar prices. The error correction coefficients tend to be a bit smaller for this equation, which suggests adjustment to the assumed long run steady state of no price discrimination takes longer.

For Japan, short run price adjustment based on the unconstrained long run model appears to match the long run behavior rather well. When the law of one price is imposed as the steady state, short run behavior still exhibits a good deal of PTM. Once again, the adjustment to past errors is much smaller when price equalization is assumed in the steady state. For the United States, the short run results are very similar for each long run specification reflecting the fact that even in the unconstrained model there was little evidence of PTM. Not surprisingly, the evidence of short run PTM

12. The appendix discusses the empirical implications of export contracts in different currencies.

is minimal. Adjustment to the steady state appears rapid, with error correction coefficients clustered around 1.

Total Merchandise

The regression results for total merchandise are presented in Tables 10 and 11. The estimation period is 1968-87. According to Table 10, Canadian exports are characterized by a very large degree of PTM. With the exception of U.S.-Canadian trade, Japan is the destination market with the largest β's. However, Japanese exports of merchandise are characterized by only a modest degree of PTM. This finding is strikingly different from the results in Table 2 for Japanese auto exports, suggesting that Japanese exporters in industries other than automobiles engage in far less PTM. Due to aggregation across products we cannot tell whether these discrepancies in overall PTM behavior reflect differing behavior of countries' firms in each industry, or a different product mix of exports for each country, or both. While U.S. exporters appear to price to the Canadian and Japanese markets, there is no evidence of PTM by U.S. exporters to Germany and the United Kingdom.[13]

Figure 13 plots the estimated time effects for U.S. exports (UTHETA) along with the export unit values for U.S. exports to Canada (UPC) and the United Kingdom (UPE). (All of the series were de-meaned before

13. The apparently high degree of U.S. pricing to the Canadian market does not necessarily imply that the two markets are not integrated. It may be the case that prices charged to Canada are nearly the same as those for the United States and that the U.S.-Canadian exchange rate divided by the Canadian wholesale price level is a good proxy for the inverse of the U.S. wholesale price level, which is not included in the regression. Indeed, the observed low variability of the U.S.-Canadian real exchange rate may be *prima facie* evidence of U.S.-Canadian integration.

plotting.) Figure 14 shows the strong negative correlation between U.S. export unit values to Canada and the price-level-adjusted U.S.-Canadian exchange rate (UEXC). Figure 15 shows that the correlation is less pronounced between U.S. exports to the U.K. and the U.S.-U.K. exchange rate (UEXE). The estimated time effects were always well-behaved for total merchandise, in that they grew at roughly the same rate as the average export unit values.

Table 11 displays the error correction results for total merchandise. It appears that the estimate of α^3 is very close to zero in the vast majority of cases, regardless of which first stage regression is run. Thus, these data are sympathetic to the symmetry hypothesis even in the short run. The only exceptions occur in German exports when long run β is unconstrained and in Canadian exports to the United States.

The estimates of α^2 tend to be much larger in magnitude and more often significant when β is unconstrained in the long run than when β is constrained at zero. The estimates of α^2 when β is unconstrained are very highly correlated with the unconstrained estimates of β. The most notable difference between the estimates of β and the corresponding estimates of α^2 is that the estimates of α^2 tend to lie between β and 0. The only exceptions to this pattern are Japanese exports to Canada and the United States.

Recall that α^2 represents short run PTM and β represents long run PTM. Lower short run PTM than long run PTM is consistent with the hypothesis that currency contracts are typically in the exporter's currency. Thus, the evidence of Table 11 supports the hypothesis of export price contracts in exporter currency except for Japanese exports to Canada and the United States. These results are similar to those for automobile exports.

The error correction coefficients are almost always larger and more significant when β is unconstrained in the long run. This result supports the hypothesis of non-zero long run PTM, but the statistical significance of this support is uncertain due to nonstationarity of the data.

4. Effects of Aggregation on Estimation of Pricing to Market

Existing studies of exchange rate pass-through and PTM vary in many dimensions - export country, destination market, sample period, industry category, and industry aggregation. This paper has provided a great deal of variation in sources and destinations, and it has covered a reasonably long sample period. We now turn to a detailed consideration of how aggregation affects estimates of PTM, focusing on Japanese auto exports.

Japanese auto exports consist of three 7-digit categories based on engine size (1 liter or less, over 1 liter but not more than 2 liters, and over 2 liters). Table 12 reports the annual quantity of autos by category exported to Canada, the United Kingdom, the United States, and Germany, respectively. (Note that Canada did not receive enough shipments to be included in the sample of destinations for less than 1 liter engines.) The United States is by far the largest buyer -- importing more cars in each category than the other countries combined. The two larger categories make up the bulk of auto exports. Japan's exports have grown most rapidly to West Germany, followed by the United States, over the 1973-1987 period. It should also be noted at the outset that we made no attempt to control for the effect of VERs on auto exports to the United States and Canada. It has been well documented by Feenstra (1988) that these quantity constraints led to quality upgrading. This is likely to show up as a shift toward larger autos, as well as an increase in unit values by category during the period in which the VER is binding. If the VER binds during periods of strong dollar (say, 1981 to 1985), then we might expect to see an upward bias in measured PTM to the United States. However, as was discussed in the previous section, there is only mild evidence of an increase in PTM by Japanese exporters to the United States over this period.

We report the estimated value of β by destination for each separate category of auto exports, for total autos (obtained by summing the value and quantity for each category) and for a model that uses the disaggregated categories but constrains β to be common across categories within a particular destination. (This constrained model is only estimated in the long-run equation.)

Turning first to the results for the long run regressions (equations 3.1)) in Table 13, we see remarkably similar coefficient estimates for the two larger categories within the disaggregated group. Coefficient estimates suggest that export price adjustment offset about 75 percent of the effect of exchange rate changes for each destination. Results for small cars are quite disparate -- with no measured PTM to the United States and virtually complete PTM to the United Kingdom. It is worth noting that the United Kingdom buys relatively more small cars than either the United States or West Germany.

The constrained regression behaves rather poorly in the sense that the standard errors of the coefficient estimates are large. The estimated coefficient for Canada falls well outside the range spanned by the coefficients obtained by estimating the two categories separately. In addition, the U.S. coefficient of -0.24 is much closer in magnitude to the result for small cars than the other two categories. This is troubling since small cars are a trivial share of total U.S. imports, yet they appear to drive the results of the constrained equation.

The long run results using total cars match much better with the underlying behavior implied by the individual categories. The parameter estimates range from .8 to just over 1, suggesting nearly complete long run pricing to market. The estimated coefficient for the United States is nearly identical to those for the United Kingdom and West Germany.

The results for the error correction model (equation (3.2)) are in Tables 14 and 15, using the unconstrained long run equation and the long run with $\beta = 0$, respectively. (These tables report only the short run PTM coefficient, α^2, and none of the other coefficients.) The estimates of short run PTM tend to be much smaller when long run PTM is constrained to zero.

The effect of aggregating to total cars is peculiar for Canada in both cases -- with the β for total cars showing less short run PTM than any individual categories, although it is not clear that the difference is statistically significant. Apart from that, the estimated short run PTM does not appear to be distorted much by aggregating over categories. It does, however, mask strange behavior for small cars -- remarkably little PTM and virtually none to the United States.

6. Conclusion

This paper has attempted to provide several different perspectives on pricing to market. We have attempted to distinguish long run behavior from short run disequilibrium dynamics, make comparisons across sources and destinations, and examine the impact of aggregation on our results.

We have considered two hypotheses about PTM in the long run. One hypothesis is that PTM persists indefinitely and may be different for different exporters, destinations, and industries. The other hypothesis is that PTM cannot persist in the long run, presumably because arbitrage eventually prevails. It is very difficult to choose between the two long run models, perhaps because the exchange rate is itself cointegrated with marginal costs. (Recall that the exchange rate is deflated by the foreign price level, so it is essentially the reciprocal of the foreign price level converted into the exporter's currency.) Unfortunately, our use of estimated time effects to control for marginal cost make it impossible to use standard cointegration tests on this hypothesis.

If the exchange rate is cointegrated with marginal cost as captured by $\underset{\sim}{\theta}$, then β is simply the coefficient on deviations from the cointegrating relationship between the exchange rate and marginal cost, and it may be subjected to standard hypothesis testing. Standard tests of the statistical significance of β almost always reject the null hypothesis that β=0. One interpretation of these results is that PTM can persist in the long run within certain bounds. Beyond these bounds arbitrage pressures may operate to keep foreign prices from deviating too far from domestic costs. Figures 1 and 9 certainly support the view that PTM can persist over the longer term and that the bounds may be quite wide.

For both long run models, the short run dynamics appear reasonable. The source of short term disequilibrium seems to involve price rigidity in the exporter's currency and not in the importer's currency. The notable exceptions to this pattern were German exports of autos to the United States and Canada, and Japanese exports of autos and total merchandise to the United States and Canada.

Source and destination effects are harder to label. Our prior beliefs, influenced by previous research and popular press accounts, were that foreign exporters tend to price to the U.S. market more than to other destinations, and that Japan and Germany do more pricing to market than other source countries. The data presented here do not provide much support for this view. While it does seem true that Japan does more pricing to market in autos than other suppliers, the United States does not stand out as a destination market for autos that is characterized by an unusual degree of PTM. Even more surprising, we find that for total merchandise exports, the United States appears to price to the Japanese market more than the Japanese price to the U.S. market. One hypothesis that we would like to explore further is that groups of countries may form integrated markets, such as the countries of the European Economic Community or the United States and Canada.

Finally, we saw that aggregation over product categories does not seem to obscure the broad patterns in the data, but may neglect some interesting heterogeneity in behavior in specific subcategories. In order to draw inferences about industry behavior, 7-digit data may be superior. To gauge source and destination market effects across a wide range of industries, more aggregated categories may be sufficient.

This paper falls short of providing a meaningful explanation for the stark differences in PTM across source countries. For total merchandise these differences may reflect differences in the types of goods that the countries export. However, PTM behavior remains different across source countries even within specific categories of automobile exports. Given the number of destination markets each of the sellers has in common, it is hard to argue that the differences could be due to different demand characteristics or greater barriers to arbitrage. The market share argument is not convincing since PTM is so pronounced for all Japanese destination markets. The persistence of price differentials seems to rule out invoicing asymmetries as well.

One potential explanation for differences in PTM behavior across source countries is that PTM is an essential strategy for firms that do not have production facilities in their target markets. U.S. automakers tend to serve foreign markets primarily out of foreign production. Their domestic exports are largely specialty items that command minuscule market shares abroad. The Japanese had very limited foreign production over this sample. In order to maintain market share, the Japanese were forced to price to market. One puzzle for this hypothesis is that German exporters of large cars apparently do very little PTM.

Appendix: Adjustment Costs, Export Invoicing, and Error Correction

It is well-known that optimal control problems with quadratic adjustment costs typically yield reduced form equations of the error correction class.[14] This appendix demonstrates that an adjustment cost model of trader behavior may give rise to an error correction mechanism in export prices. The model used here assumes that there is a one period lag between the decision to export and the time of shipment, there are costs of adjusting the volume of trade from period to period, and there are no inventories. Within this framework we consider the implications of price contracts and different currencies of invoice for the short run dynamics of export prices.

The overall conclusions of this appendix may be stated as follows: First, when there are no price contracts in effect at the time of shipment, the short run impact of the exchange rate on the export price is greater than the long run impact. The same result holds true when the export price is contracted in the importer's currency. When the export price is contracted in the exporter's currency prior to the time of shipment, the short run impact of the exchange rate is less than the long run impact.

We begin with the case in which there are no price contracts and the trader simply sells the export good in the importing country when it arrives. The trader is assumed to maximize the real discounted flow of future profits from exports.[15] His objective is given by (A.1) and he is assumed to face market demand in the importing country given by (A.2). He

14. See Nickell (1985).

15. The model of this appendix is adapted from Gagnon (1989).

faces a cost of adjusting the volume of trade that is quadratic in the size of the adjustment. This adjustment cost is assumed to occur in the importing country, but the qualitative nature of the results is not affected by assuming that the adjustment cost occurs in the exporting country.

$$\text{(A.1)} \quad \underset{q_t}{\text{Max}}\ E_{t-1} \sum_{i=0}^{\infty} \theta^i \left\{ p_{t+i}q_{t+i} - c_{t+i}q_{t+i} - \frac{f\Pi^*_{t+i}(q_{t+i}-q_{t+i-1})^2}{e_t} \right\} / \Pi_{t+i}.$$

$$\text{(A.2)} \quad p^*_t = \left(a - bq_t\right)\Pi^*_t.$$

$$\text{(A.3)} \quad p^*_t = e_t p_t.$$

$$\text{(A.4)} \quad c_t = \Pi_t.$$

Here q represents the volume of exports, p is the price of exports in the exporter's currency, c is the unit cost of exports in the exporter's currency, e is the exchange rate between exporter and importer currency, Π is the aggregate price level in the exporting country, p^* is the price of exports in the importer's currency, Π^* is the aggregate price level in the importing country, and E_t is the expectations operator conditional on period t information. The real discount factor is θ. The adjustment cost parameter is f. The parameters of the demand curve are a and b. Equation (A.3) is an identity that states that the export price in the importer's currency is simply the export price in the exporter's currency times the exchange rate. Equation (A.4) represents a simplifying assumption that costs are proportional to the aggregate price level in the exporting country. Without loss of generality we can normalize the factor of proportionality at unity.

By substituting equations (A.2), (A.3), and (A.4) into (A.1) and solving for the first-order condition we obtain the following decision rule for the trader:

$$\text{(A.5)} \qquad q_t = \gamma + \alpha q_{t-1} - E_{t-1} \sum_{i=0}^{\infty} (\alpha\theta)^i \delta e_{t+i} \Pi_{t+i}/\Pi^*_{t+i}.$$

Here γ, α, and δ are (positive) functions of the underlying parameters a, b, f, and θ. We assume that the exchange rate and the aggregate price level in each country are exogenous with respect to the trader. To obtain an estimable reduced form we must first posit a stochastic process for the real exchange rate, $e\Pi/\Pi^*$. If the real exchange rate follows the first-order autoregression given by (A.6) it is easy to show that the reduced form equation for the volume of exports is (A.7). (Note that the real exchange rate may follow a random walk, in which case ρ=1.)

$$\text{(A.6)} \qquad e_t \Pi_t/\Pi^*_t = \rho e_{t-1} \Pi_{t-1}/\Pi^*_{t-1} + u_t.$$

$$\text{(A.7)} \qquad q_t = \gamma + \alpha q_{t-1} - \frac{\delta e_{t-1} \Pi_{t-1}}{(1-\rho\alpha\theta)\Pi^*_{t-1}}.$$

By substituting (A.2) and (A.3) into (A.7) we obtain the corresponding reduced form for the export price in terms of the importer's currency relative to the importer's aggregate price level:

$$\text{(A.8)} \qquad \frac{e_t p_t}{\Pi^*_t} = \left\{ (a-b\gamma-\alpha a) + \frac{\alpha e_{t-1} p_{t-1}}{\Pi^*_{t-1}} + \frac{\delta b e_{t-1} \Pi_{t-1}}{(1-\rho\alpha\theta)\Pi^*_{t-1}} \right\}.$$

According to equation (A.8) the ratio of the export price to the importer's aggregate price is stationary if and only if the real exchange rate is stationary. Equation (A.8) can be expressed as a simple error-correction model:

$$(A.9) \qquad \Delta\frac{e_t p_t}{\Pi_t^*} = -(1-\alpha)\left\{\frac{e_{t-1}p_{t-1}}{\Pi_{t-1}^*} - \Gamma - \frac{\Phi e_{t-1}\Pi_{t-1}}{\Pi_{t-1}^*}\right\}.$$

where $\Gamma = (a-b\gamma-\alpha a)/(1-\alpha)$ and $\Phi = \delta b/((1-\alpha)(1-\rho\alpha\theta))$. The long run equilibrium for the export price is obtained by setting the expression inside the braces equal to zero and rearranging terms.

$$(A.10) \qquad p_t = \frac{\Gamma\Pi_t^*}{e_t} + \Phi\Pi_t.$$

Equation (A.10) is quite similar to equation (2.1) in that the export price responds negatively to the exchange rate and positively to marginal cost. (Recall that marginal cost in this case is simply Π.) The constant term has been lost due to a normalization. The symmetry between exchange rate and marginal cost effects is present only for particular combinations of the underlying parameters. Although it is not readily apparent from the definitions of Γ and Φ, it can be shown that both of these coefficients must lie between 0 and 1. Thus, an exchange rate change has a less than proportional effect on the export price in the long run.

For values of ep/Π^* close to unity--a harmless normalization--we can write equation (A.9) in terms of the export price in the exporter's currency.

$$\text{(A.11)} \quad \frac{\Delta p_t}{p_t} = \frac{\Delta \Pi_t^*}{\Pi_t^*} - \frac{\Delta e_t}{e_t} - (1-\alpha)\left\{\frac{e_{t-1}p_{t-1}}{\Pi_{t-1}^*} - \Gamma - \frac{\Phi e_{t-1}\Pi_{t-1}}{\Pi_{t-1}^*}\right\}.$$

According to equation (A.11) the short run effect of a change in the exchange rate is to lower the export price proportionally. Thus the short run effect of an exchange rate movement on the export price is **greater** than the long run effect.

If consumers of the export good contract a fixed price in the importing country currency, the demand curve faced by the exporter would be as follows:

$$\text{(A.12)} \quad e_t p_t = \left(a - bq_t\right) E_{t-1}\left(\Pi_t^*\right).$$

If the importer's aggregate price level follows a random walk, then (A.12) can be rewritten as (A.13). (Note that this assumption implies restrictions on the exporter's aggregate price level and the exchange rate to ensure that (A.6) holds.)

$$\text{(A.13)} \quad p_t = \left(a - bq_t\right)\left(\frac{\Pi_{t-1}^*}{e_t}\right).$$

Use of demand curve (A.13) does not affect the trader's optimal decision rule for quantities (A.7). The associated dynamic equation for the export price is affected, however. Substitution of (A.13) into (A.7) and rearranging terms yields the following error correction equation for the export price:

$$(A.14)\qquad \frac{\Delta p_t}{p_t} = \frac{\Delta \Pi^*_{t-1}}{\Pi^*_t} - \frac{\Delta e_t}{e_t} - (1-\alpha)\left\{\frac{e_{t-1}p_{t-1}}{\Pi^*_{t-2}} - \Gamma - \frac{\Phi e_{t-1}\Pi_{t-1}}{\Pi^*_{t-1}}\right\}.$$

Equation (A.14) is characterized by nearly the same long run equilibrium export price as equation (A.11). In fact, since Π^* follows a random walk, it will be very difficult empirically to distinguish between the equilibrium relationships embodied in (A.11) and (A.14). The short run response of the export price to the exchange rate is also the same in equation (A.14) as in equation (A.11). The major difference is that changes in the importer's aggregate price have no short run effect on the export price. Since movements in the exchange rate are typically much larger and more unpredictable than movements in the aggregate price, equation (A.14) is likely to be empirically indistinguishable from equation (A.11).

Now suppose that consumers must contract in advance to pay for the export good in the exporter's currency at a fixed price. The demand curve faced by the exporter is given in (A.15).

$$(A.15)\qquad p_t = \left(a - bq_t\right)E_{t-1}\left(\frac{\Pi^*_t}{e_t}\right).$$

If the importer's aggregate price level converted to exporter currency follows a random walk, then (A.15) can be rewritten as (A.16). (Once again, this assumption implies restrictions on the exporter's aggregate price level to ensure that (A.6) holds.)

$$(A.16)\qquad p_t = \left(a - bq_t\right)\left(\frac{\Pi^*_{t-1}}{e_{t-1}}\right).$$

Use of demand curve (A.16) does not affect the trader's optimal decision rule for quantities (A.7), but it does affect the associated dynamic equation for the export price. Substitution of (A.16) into (A.7) and rearranging terms yields the following error correction equation for the export price:

$$\text{(A.17)} \quad \frac{\Delta p_t}{p_t} = \frac{\Delta \Pi^*_{t-1}}{\Pi^*_t} - \frac{\Delta e_{t-1}}{e_t} - (1-\alpha)\left\{\frac{e_{t-2}p_{t-1}}{\Pi^*_{t-2}} - \Gamma - \frac{\Phi e_{t-1}\Pi_{t-1}}{\Pi^*_{t-1}}\right\}.$$

Once again, equation (A.17) is characterized by nearly the same long run equilibrium export price as equation (A.11). However, the short run response of the export price to the exchange rate is zero in equation (A.17). According to (A.17) the lagged exchange rate change ought to enter the error correction equation when the export price is contracted in the exporter's currency. Empirically, the coefficient on the lagged exchange rate was usually much less significant than the coefficient on the contemporaneous exchange rate, so it was not reported in the paper.

The insignificant coefficients on lagged exchange rates are probably due to temporal aggregation in the data. When the period of observation is longer than the period of the export price contract, the different empirical implications of contracts in exporter currency and contracts in importer currency will be blurred, as some of the adjustment toward long run equilibrium will take place within the observation period. This point is relevant because the data used in this paper have an annual frequency and Magee (1974) reports that currency contracts typically last 3 months. In such a case, the coefficient on short run adjustment will be biased toward the value of long run adjustment. With contracts in the importer's currency, this implies a value of α^2 between β and 1. With contracts in the exporter's currency this implies a value of α^2 between β and 0.

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Table 1
Estimates of β in Equation (3.1)
German Exports of Automobiles

Destination	1500-1999 cc	2000-2999 cc	3000 cc and over
Canada	0.24	-0.08	-0.20
United States	0.59	0.12	-0.15
Japan	-0.64	0.90	0.09
United Kingdom	0.05	0.37	0.00
France	0.52	0.45	-0.79
Sweden	0.31	0.44	-0.68

Table 2
Estimates of β in Equation (3.1)
Japanese Exports of Automobiles

Destination	0-1000 cc	1001-2000 cc	2001 cc and over
Canada		0.79	0.80
United States	-1.66	0.79	0.81
United Kingdom	0.95	0.82	0.91
Germany	1.03	0.83	0.91
Norway			0.86
Sweden		0.82	
Switzerland	0.98		
Australia	0.35		

Table 3
Estimates of β in Equation (3.1)
U.S. Exports of Automobiles

Destination	6 cylinders or less	over 6 cylinders
Canada	-0.04	0.11
Japan	0.05	-0.02
Germany	-0.03	0.00
United Kingdom	0.16	-0.21
Sweden	0.10	0.00
Italy		0.02
Australia		0.24

Table 2A
Split Sample Estimates of β in Equation (3.1)
Japanese Exports of Automobiles

Destination	1001-2000 cc		2001 cc and over	
	73-80	81-87	73-80	81-87
Canada	0.58	0.86	0.66	0.69
United States	0.53	0.85	0.48	0.67
United Kingdom	0.19	0.86	-0.28	0.76
Germany	0.02	0.89	0.18	0.75
Norway			0.67	0.60
Sweden	0.44	0.85		

Table 4
Estimates of Equation (3.2)
German Exports of Automobiles
β unconstrained

Destination	1500-1999 cc			2000-2999 cc			3000 cc and over		
	α^2	α^3	α^4	α^2	α^3	α^4	α^2	α^3	α^4
Canada	.44	-.01	.55	-.01	-.63*	.47	-.31*	.56*	.72*
	(.23)	(.91)	(.30)	(.14)	(.19)	(.30)	(.11)	(.18)	(.31)
United States	.72*	.27	.24	.21*	.17	.76*	-.13	.63*	.64
	(.11)	(.48)	(.16)	(.05)	(.13)	(.09)	(.23)	(.15)	(.44)
Japan	-.66*	-.40*	1.08*	.01	.85*	.23*	-.06	.54*	.46*
	(.07)	(.18)	(.15)	(.18)	(.31)	(.15)	(.23)	(.18)	(.23)
United Kingdom	-.24	-.02	.71*	.06	.55*	.74*	-.25*	.49*	.46*
	(.37)	(.60)	(.22)	(.15)	(.14)	(.14)	(.12)	(.18)	(.17)
France	.06	.79*	.27	.13	.75*	.62	-.84*	-.50*	.71*
	(.22)	(.23)	(.19)	(.24)	(.23)	(.32)	(.28)	(.16)	(.19)
Sweden	.11	.65*	.38*	.08	.35	.23	-.50	-.94*	.66*
	(.12)	(.21)	(.10)	(.21)	(.39)	(.31)	(.29)	(.25)	(.20)

Table 5
Estimates of Equation (3.2)
German Exports of Automobiles
$\beta = 0$

Destination	1500-1999 cc			2000-2999 cc			3000 cc and over		
	α^2	α^3	α^4	α^2	α^3	α^4	α^2	α^3	α^4
Canada	-.07 (.25)	-1.58* (.65)	.58* (.16)	-.21 (.20)	-.54 (.47)	.48 (.28)	-.10* (.05)	.48* (.24)	.38 (.22)
United States	.54* (.23)	-.30 (.68)	.17 (.12)	.12 (.04)	.15 (.17)	.44* (.06)	-.06 (.10)	.44* (.20)	.76* (.33)
Japan	-.46* (.22)	.66 (.36)	.33 (.28)	-.08 (.18)	.17 (.45)	.24 (.15)	.05 (.16)	.36* (.18)	.46* (.10)
United Kingdom	-.04 (.41)	.73 (.48)	.69* (.17)	.13 (.31)	.16 (.23)	.45 (.36)	-.23 (.12)	.47 (.27)	.05 (.17)
France	-.03 (.18)	.93* (.25)	.04 (.09)	.06 (.28)	.32* (.10)	.59 (.37)	-.18 (.41)	-.75* (.28)	.23* (.11)
Sweden	.13 (.14)	.75* (.16)	.32* (.10)	-.15 (.15)	-.03 (.31)	.29 (.15)	-.08 (.22)	-.86 (.57)	.35* (.13)

Table 6
Estimates of Equation (3.2)
Japanese Exports of Automobiles
β unconstrained

Destination	0-1000 cc			1001-2000 cc			2001 cc and over		
	α^2	α^3	α^4	α^2	α^3	α^4	α^2	α^3	α^4
Canada				.76*	.09	.33	.94*	-.12	.35*
				(.08)	(.11)	(.17)	(.06)	(.09)	(.17)
United States	-1.67*	-.17	.89*	.69*	.20	.28	.88*	-.05	.65*
	(.11)	(.10)	(.38)	(.10)	(.14)	(.19)	(.06)	(.07)	(.23)
Germany	.45*	.63*	.24	.63*	.22	.60*	.85*	.08	.64*
	(.14)	(.10)	(.16)	(.13)	(.17)	(.18)	(.19)	(.23)	(.30)
United Kingdom	.43*	.77*	.47*	.46*	.41	.41*	.40	.53	.47*
	(.08)	(.08)	(.10)	(.17)	(.22)	(.17)	(.29)	(.40)	(.13)
Norway							1.00*	-.25	.29
							(.17)	(.22)	(.20)
Sweden				1.06*	-.36*	.62*			
				(.12)	(.17)	(.24)			
Switzerland	.44*	.64*	.30*						
	(.16)	(.13)	(.19)						
Australia	.27	.69	.88*						
	(.53)	(.44)	(.39)						

Table 7
Estimates of Equation (3.2)
Japanese Exports of Automobiles
$\beta = 0$

Destination	0-1000 cc			1001-2000 cc			2001 cc and over		
	α^2	α^3	α^4	α^2	α^3	α^4	α^2	α^3	α^4
Canada				.40*	.37	.07	.46*	.08	.05
				(.07)	(.23)	(.11)	(.15)	(.21)	(.08)
United States	-.32	-.76	.37*	.49*	.49*	-.05	.61*	.45	.02
	(.24)	(.70)	(.17)	(.08)	(.14)	(.12)	(.16)	(.24)	(.10)
Germany	.26*	.47*	.22*	.11	.33	.21	.73*	.14	.22
	(.09)	(.11)	(.06)	(.17)	(.41)	(.16)	(.22)	(.17)	(.13)
United Kingdom	.08	.72*	.22*	-.01	.48*	.38*	-.10	.23	.31
	(.07)	(.12)	(.07)	(.08)	(.23)	(.14)	(.17)	(.22)	(.19)
Norway							.00	-.12	.31
							(.19)	(.22)	(.16)
Sweden				.02	-.19	.21			
				(.10)	(.32)	(.11)			
Switzerland	.17	.55*	.28*						
	(.14)	(.19)	(.07)						
Australia	.25	-1.57*	.09						
	(.27)	(.53)	(.31)						

Table 8
Estimates of Equation (3.2)
U.S. Exports of Automobiles
β unconstrained

Destination	6 cylinders or less			over 6 cylinders		
	α^2	α^3	α^4	α^2	α^3	α^4
Canada	.25*	.37	-.05	.24*	.48*	.42*
	(.09)	(.12)	(.13)	(.10)	(.48)	(.15)
Japan	.05*	.34*	.98*	-.12	.44*	1.05*
	(.12)	(.15)	(.23)	(.17)	(.20)	(.25)
Germany	-.19	.15	.99*	.07	.44	1.26*
	(.10)	(.08)	(.24)	(.23)	(.28)	(.27)
United Kingdom	.18	.03	1.60	-.23	-.27	1.25*
	(.12)	(.31)	(.22)	(.32)	(.45)	(.24)
Sweden	-.03	-.49	1.31*	.08	.60	.96*
	(.16)	(.45)	(.26)	(.21	(.43)	(.34)
Italy				-.30	-1.71*	1.04*
				(.69)	(.55)	(.16)
Australia				-.06	.43	1.11*
				(.99)	(.84)	(.29)

Table 9
Estimates of Equation (3.2)
U.S. Exports of Automobiles
$\beta = 0$

Destination	6 cylinders or less			over 6 cylinders		
	α^2	α^3	α^4	α^2	α^3	α^4
Canada	.25 (.84)	.36 (.91)	-.09 (.10)	.31* (.10)	.40* (.14)	.38* (.16)
Japan	.03 (.12)	.34* (.15)	.96* (.22)	-.13 (.17)	.53* (.20)	1.07* (.19)
Germany	-.23 (.69)	.41* (.08)	1.04* (.24)	.04 (.21)	.49 (.28)	1.25* (.24)
United Kingdom	.06 (.12)	-.40 (.32)	1.61* (.22)	-.20 (.37)	.23 (.67)	1.10* (.26)
Sweden	-.06 (.16)	-.73 (.40)	1.21* (.22)	.04 (.20)	.52 (.44)	.96* (.34)
Italy				-.14 (.68)	-1.75* (.70)	1.04* (.19)
Australia				.00 (1.02)	-.18 (.87)	1.04* (.35)

Table 10
Estimates of β in Equation (3.1)
Total Merchandise

Source	Destination				
	Canada	U.S.	Japan	Germany	England
Canada		1.311	.629	.582	.618
U.S.	.932		.522	-.184	-.073
Japan	.290	.252		.303	-.064
Germany	.582	.339	.772		-.411
England	.093	.112	.456	.269	

Table 11
Estimates of Equation (3.2)
Total Merchandise

Destination	$\beta = 0$			β unconstrained		
	α^{2}	α^{3}	α^{4}	α^{2}	α^{3}	α^{4}
			Canadian Exports			
United States	0.38 (.24)	-0.38* (.15)	-0.03 (.07)	1.34* (.22)	-0.52* (.21)	0.24 (.15)
Japan	0.02 (.09)	-0.01 (.09)	0.21 (.16)	0.52* (.07)	-0.11 (.10)	0.16 (.14)
Germany	0.11 (.09)	0.04 (.10)	0.03 (.10)	0.51* (.06)	0.02 (.09)	0.96* (.25)
United Kingdom	0.06 (.07)	0.09 (.07)	0.07 (.13)	0.59* (.04)	-0.08 (.06)	1.08 (.20)
			U.S. Exports			
Canada	0.30 (.20)	0.15 (.13)	0.22 (.13)	0.56* (.17)	0.20 (.13)	0.45* (.18)
Japan	0.22* (.07)	-0.12 (.08)	0.16* (.08)	0.33* (.08)	-0.03 (.10)	0.24 (.15)
Germany	-0.04 (.13)	-0.09 (.13)	0.08 (.19)	0.03 (.08)	-0.04 (.09)	0.52* (.20)
United Kingdom	-0.21* (.10)	-0.04 (.11)	0.62* (.19)	-0.20* (.09)	0.04 (.09)	0.58* (.20)
			Japanese Exports			
Canada	0.29* (.05)	-0.08 (.05)	0.13 (.13)	0.42* (.05)	-0.06 (.05)	0.23 (.14)
United States	0.26* (.06)	-0.04 (.06)	0.02 (.11)	0.38* (.06)	-0.02 (.06)	0.11 (.12)
Germany	0.11 (.09)	-0.01 (.10	0.33 (.18)	0.28* (.08)	-0.01 (.08)	0.52 (.27)
United Kingdom	-0.06 (.11)	0.01 (.11)	0.47 (.17)	-0.03 (.07)	0.00 (.08)	0.34* (.14)

Table 11 (cont'd)

Destination	$\beta = 0$ α^2	α^3	α^4	β free α^2	α^3	α^4
			German Exports			
Canada	0.22*	0.16	0.23*	0.37*	0.31*	0.34
	(.07)	(.11)	(.09)	(.08)	(.11)	(.18)
United States	0.19*	0.14	0.37*	0.30*	0.32*	0.42*
	(.09)	(.13)	(.18)	(.09)	(.12)	(.20)
Japan	-0.04	0.35*	0.28*	0.19	0.49*	0.17
	(.12)	(.14)	(.12)	(.14)	(.15)	(.13)
United Kingdom	-0.19	-0.64	0.35	-0.35*	-0.22*	0.38
	(.22)	(.23)	(.12)	(.09)	(.10)	(.20)
			U.K. Exports			
Canada	0.10	-0.08	0.45	0.12	-0.06	0.44
	(.15)	(.09)	(.26)	(.10)	(.07)	(.24)
United States	0.10	0.06	0.40	0.11	0.07	0.28
	(.14)	(.08)	(.21)	(.10)	(.06)	(.21)
Japan	-0.03	-0.02	0.99*	0.21	0.01	1.39
	(.19)	(.15)	(.30)	(.12)	(.11)	(.22)
Germany	0.03	0.02	0.48	0.25*	0.02	0.81*
	(.09)	(.08)	(.24)	(.07)	(.06)	(.26)

Table 12
Japanese Passenger Car Exports by Engine Size
(Numbers of Vehicles)

Year	Canada	United Kingdom	United States	Germany
		Small Cars: ≤ 1.0 L		
1973		17486	1011	8422
1974		26573	219	6828
1975		28156	4	15589
1976		20502	2514	8686
1977		29364	12290	5161
1978		19859	13105	7794
1979		25157	9705	9549
1980		16979	3768	19320
1981		12338	7034	25034
1982		9432	15404	17168
1983		20630	7211	35553
1984		27016	111796	41074
1985		37475	282275	48392
1986		39532	269953	61099
1987		53470	186168	60721
		Medium Cars: > 1.0 L, ≤ 2.0 L		
1973	69898	58945	551342	25326
1974	109120	64286	634597	11634
1975	68771	90040	585283	32211
1976	104214	98332	855863	44531
1977	113380	129984	1035550	62579
1978	106025	140580	1131490	112730
1979	54895	157871	1320072	165613
1980	145920	145827	1632307	237823
1981	179665	144272	1492303	219832
1982	138824	143255	1391454	177042
1983	153175	168239	1443281	217645
1984	165637	157634	1594958	257845
1985	196973	149057	1937732	249290
1986	212392	156743	2005540	297813
1987	208178	135306	1809098	342283

Table 12 (cont'd)

Year	Canada	United Kingdom	United States	Germany
		Large Cars: > 2.0 L		
1973	6471	5410	73629	356
1974	6625	2610	96829	138
1975	7042	2545	153812	925
1976	12446	1907	238548	1377
1977	13473	3726	344087	1211
1978	18415	3011	376813	2987
1979	8606	8699	312558	5782
1980	15408	3220	284807	5233
1981	25179	5751	341674	17854
1982	19575	5379	387094	8482
1983	17933	6516	347182	16091
1984	18109	6749	327679	25311
1985	16855	7586	331977	28822
1986	19754	7855	410762	38607
1987	29345	11690	449710	49905

Table 13
Japanese Auto Exports
Long-Run PTM: β

Destination	≤1.0L	>1.0L, ≤2.0L	>2.0L	Constrained	Total
Canada		1.06	0.71	0.37	1.07
United Kingdom	1.06	0.76	0.86	0.84	0.82
United States	0.02	0.71	0.72	0.24	0.79
West Germany	0.66	0.78	0.87	0.85	0.80

Table 14
Japanese Auto Exports

Short Run PTM: α^2
(Long Run β Unconstrained)

Destination	≤1.0L	>1.0L, ≤2.0L	>2.0L	Total
Canada		0.75 (.09)	0.79 (.05)	0.47 (.42)
United Kingdom	0.23 (.14)	0.57 (.12)	0.38 (.19)	0.74 (.10)
United States	-0.04 (.09)	0.55 (.06)	0.79 (.06)	0.62 (.08)
West Germany	0.40 (.11)	0.76 (.09)	0.71 (.19)	0.60 (.08)

Table 15
Japanse Auto Exports

Short Run PTM: α^2
(Long Run $\beta = 0$)

Destination	≤1.0L	>1.0L, ≤2.0L	>2.0L	Total
Canada		0.27 (.08)	0.27 (.17)	-0.14 (.16)
United Kingdom	-0.03 (.12)	0.00 (.10)	-0.08 (.19)	0.03 (.11)
United States	-0.22 (.10)	0.38 (.09)	0.37 (.18)	0.34 (.06)
West Germany	0.34 (.11)	0.39 (.13)	0.83 (.26)	0.26 (.11)

Figure 1

Data from "JAPAN CARS, 1000-2000 cc"

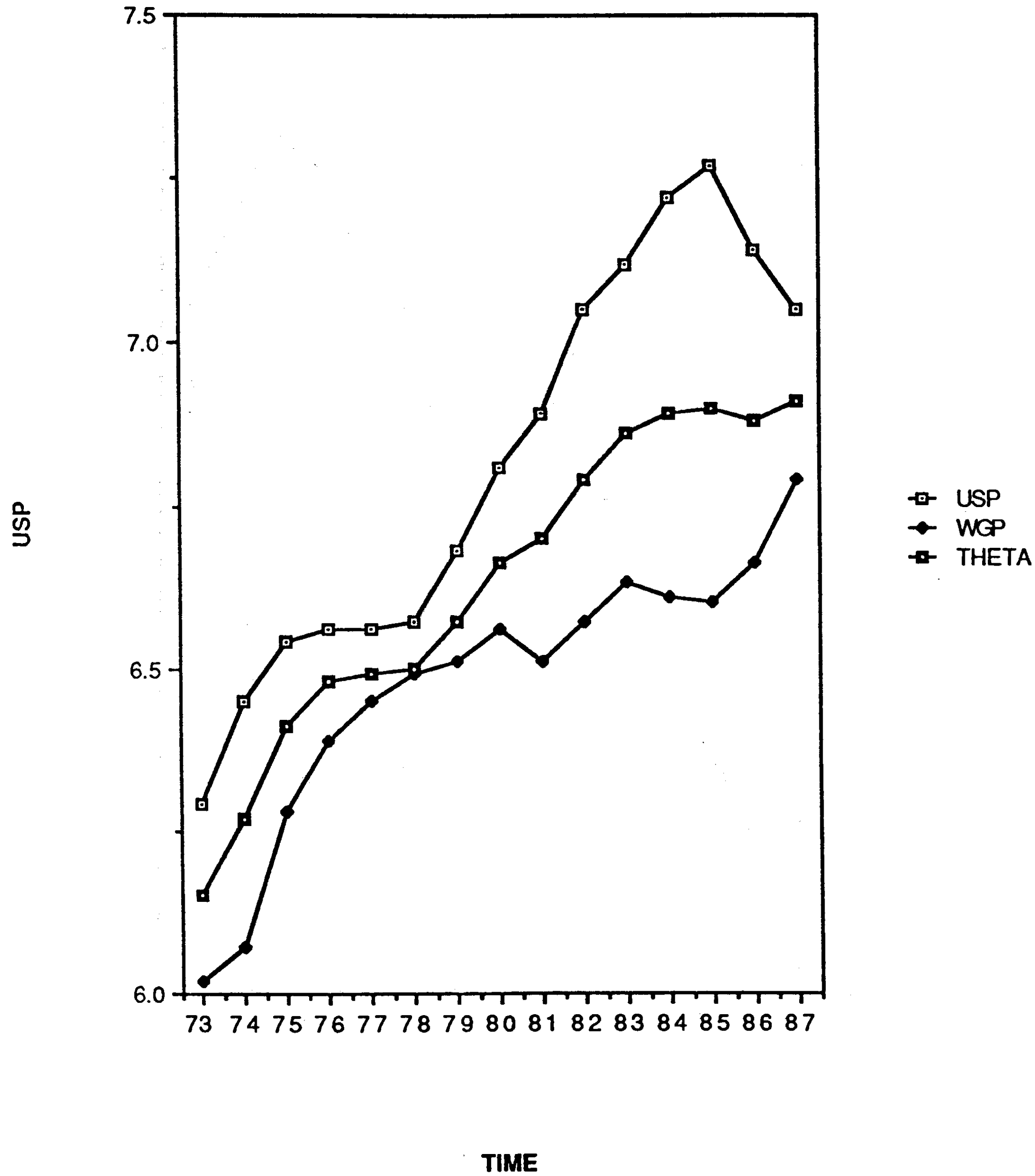


Figure 2

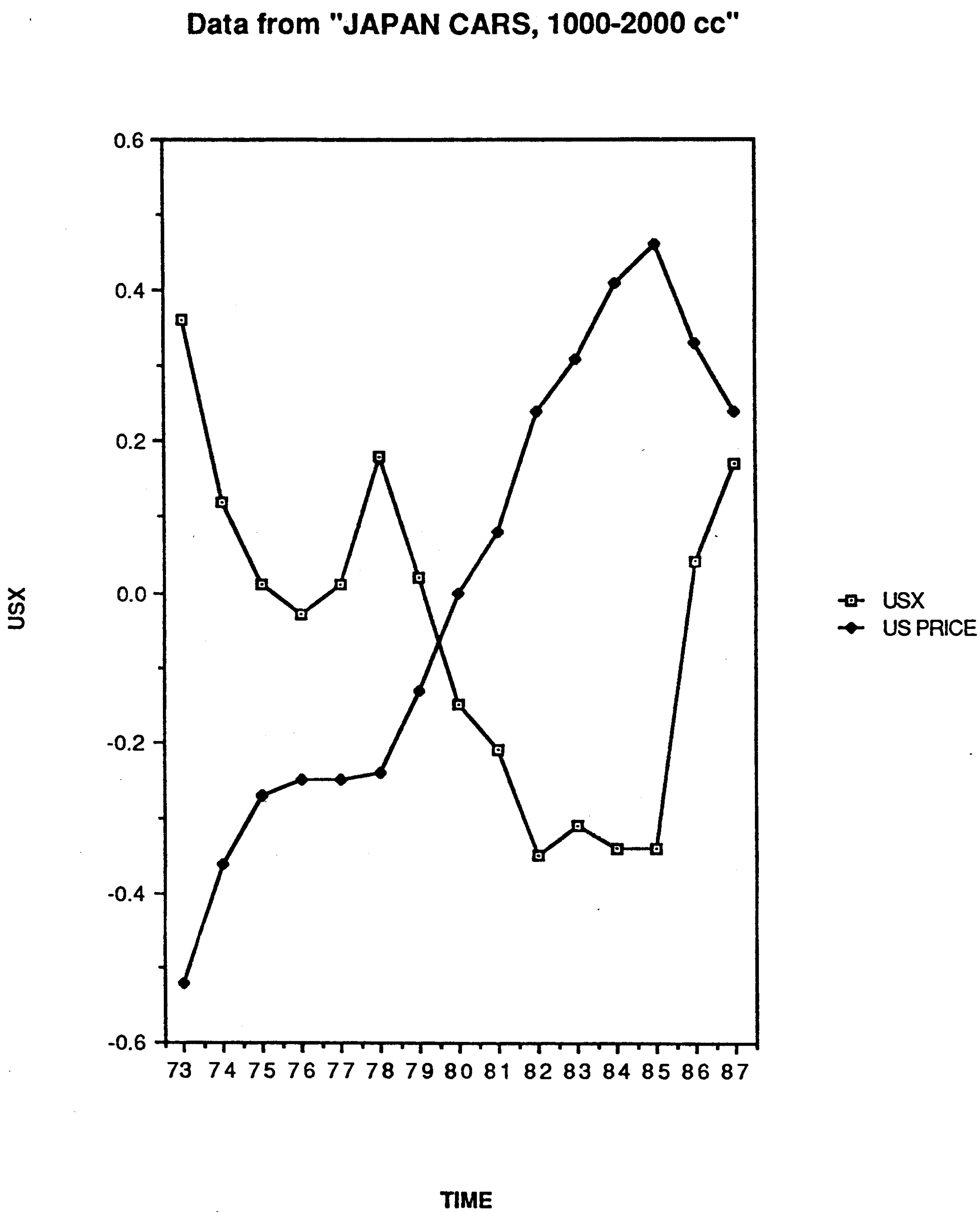

Data from "JAPAN CARS, 1000-2000 cc"
0.6
0.4
0.2
0.0
-0.2
-0.4
-0.6
USX
73 74 75 76 77 78 79 80 81 82 83 84 85 86 87
TIME
USX
US PRICE

Figure 3

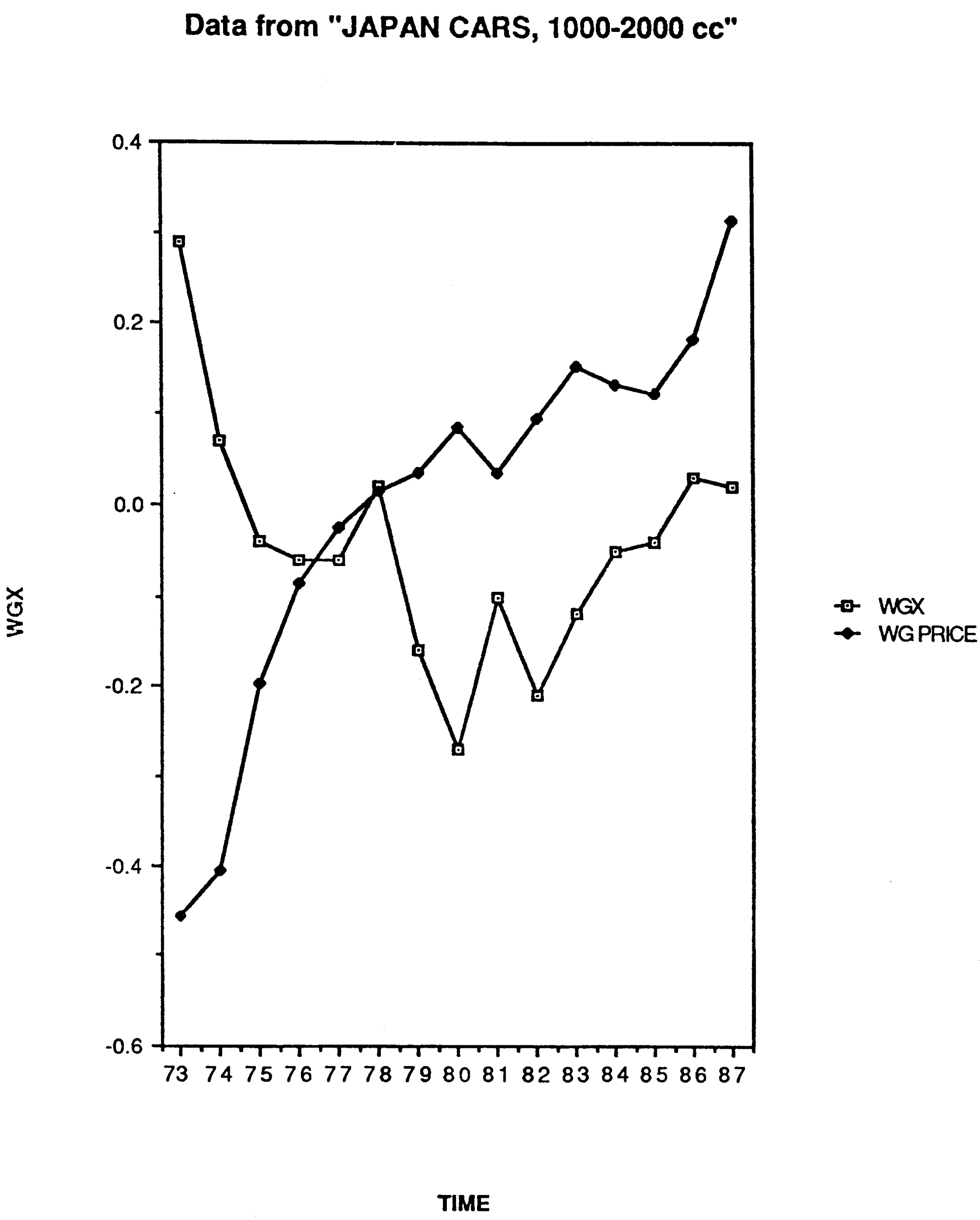

Data from "JAPAN CARS, 1000-2000 cc"
0.4
0.2
0.0
-0.2
-0.4
-0.6
WGX
73 74 75 76 77 78 79 80 81 82 83 84 85 86 87
TIME
WGX
WG PRICE

Figure 4

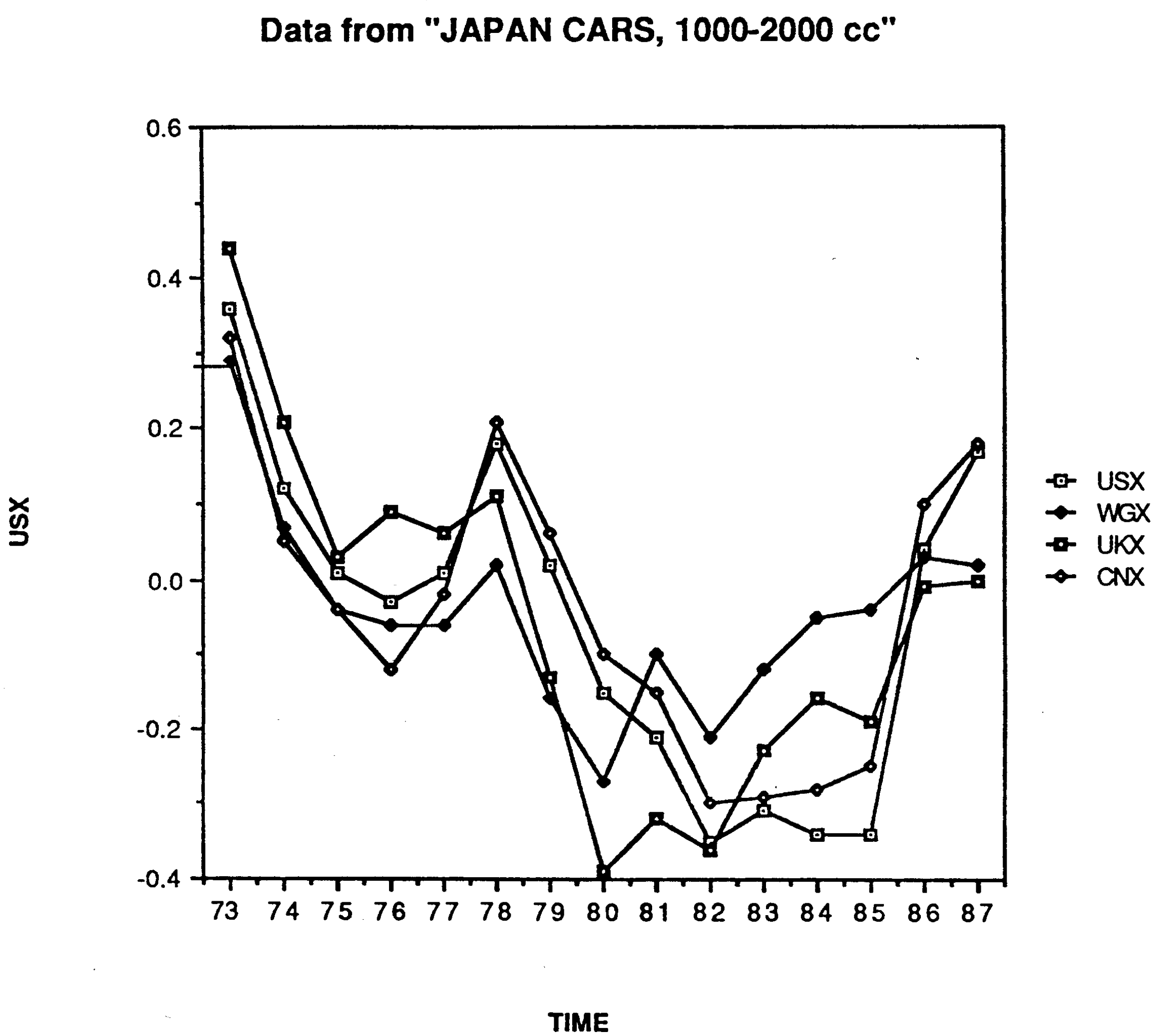

Data from "JAPAN CARS, 1000-2000 cc"
USX
0.6
0.4
0.2
0.0
-0.2
-0.4
73 74 75 76 77 78 79 80 81 82 83 84 85 86 87
TIME
USX
WGX
UKX
CNX

Figure 5

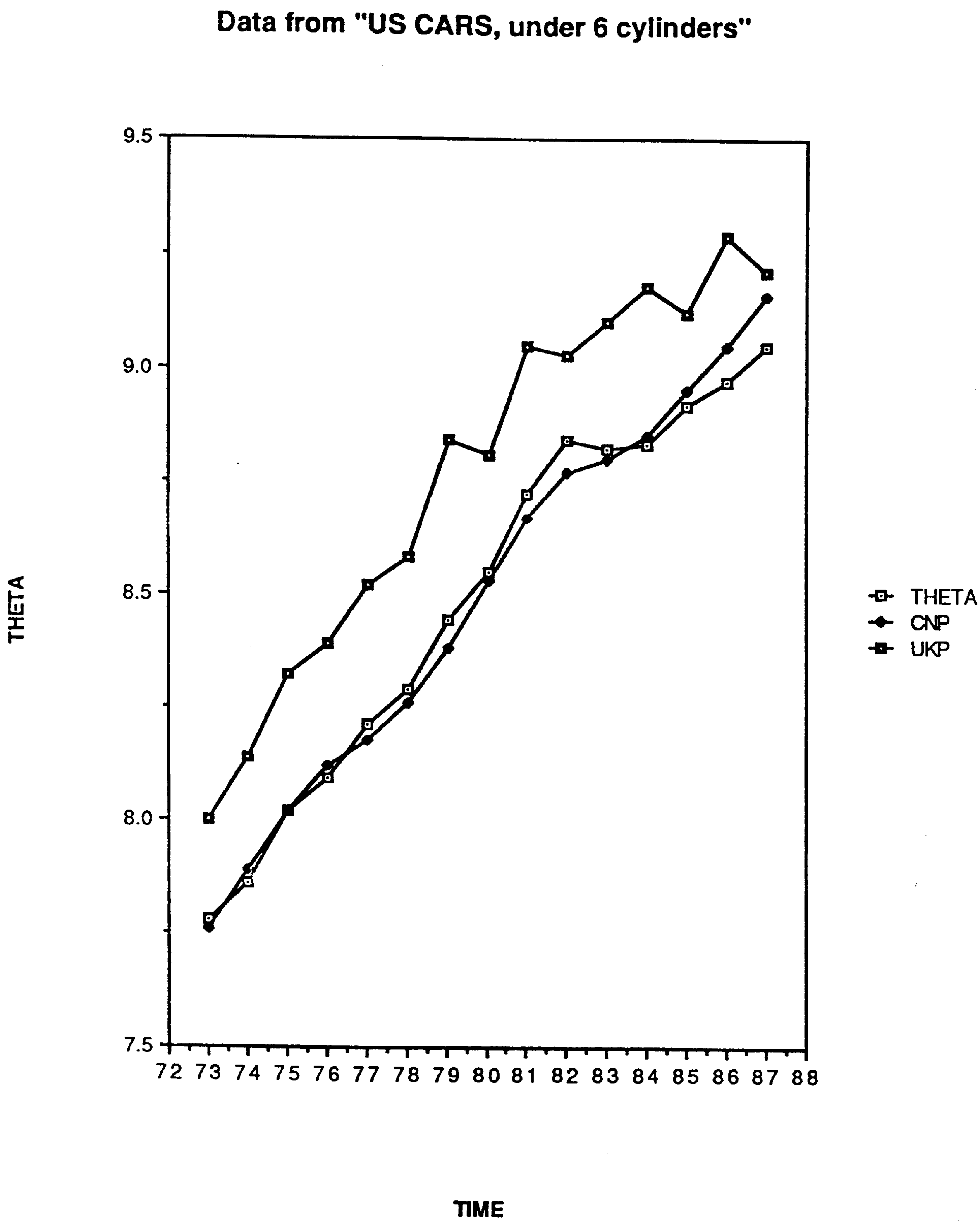

Data from "US CARS, under 6 cylinders"
9.5
9.0
8.5
8.0
7.5
THETA
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88
TIME
THETA
CNP
UKP

Figure 6

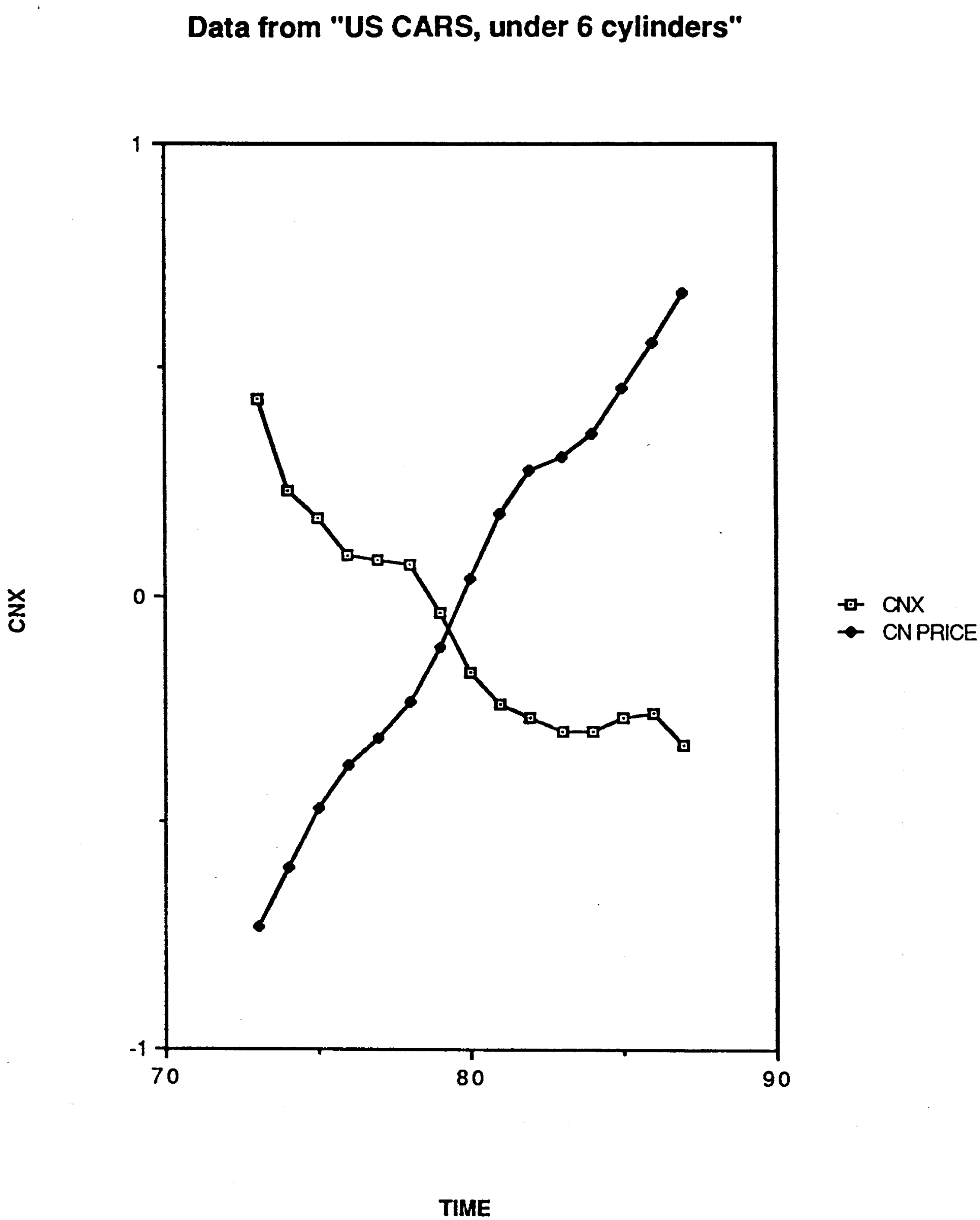

Data from "US CARS, under 6 cylinders"
1
0
-1
70
80
90
CNX
CNX
CN PRICE
TIME

Figure 7

Data from "US CARS, under 6 cylinders"

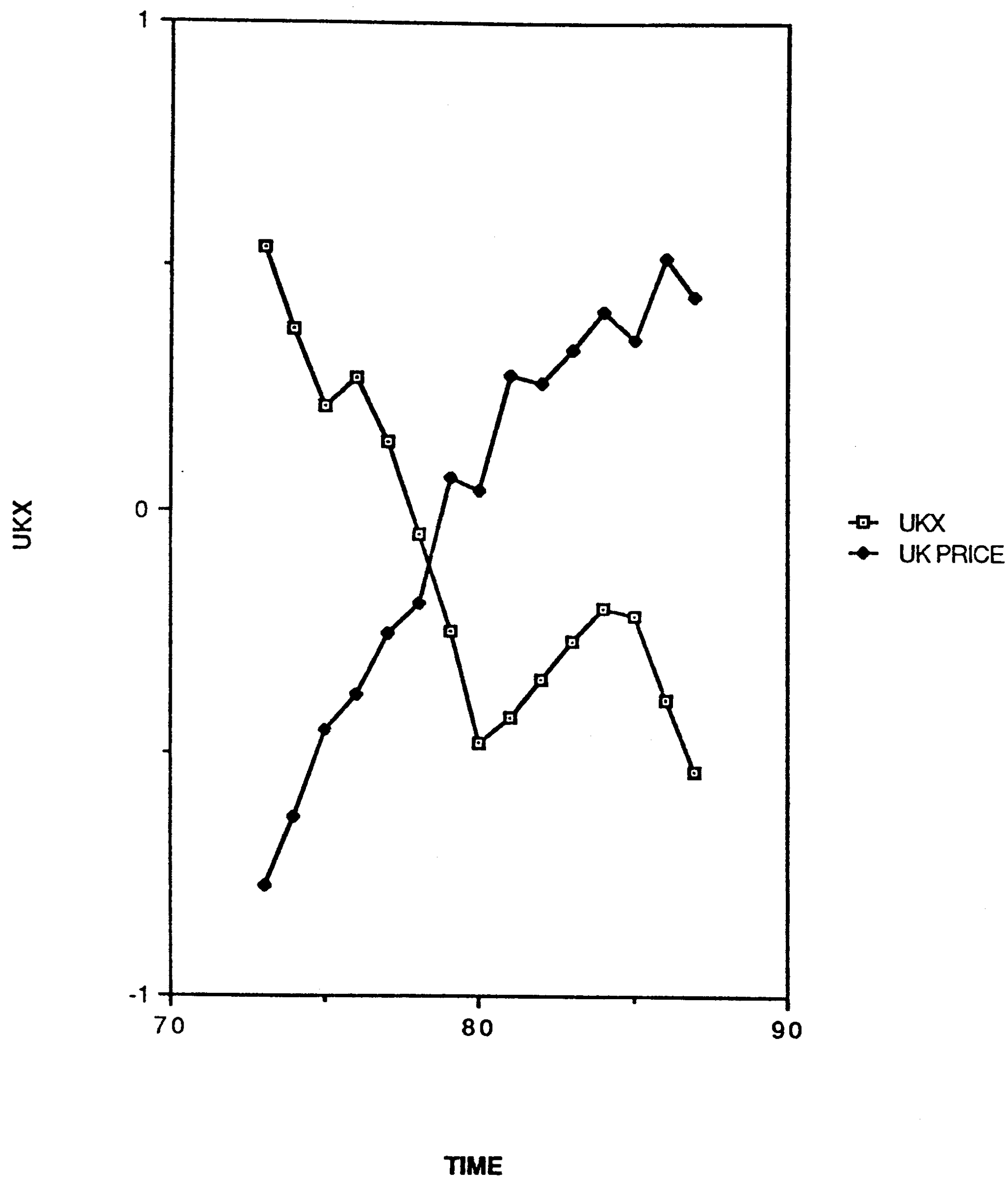
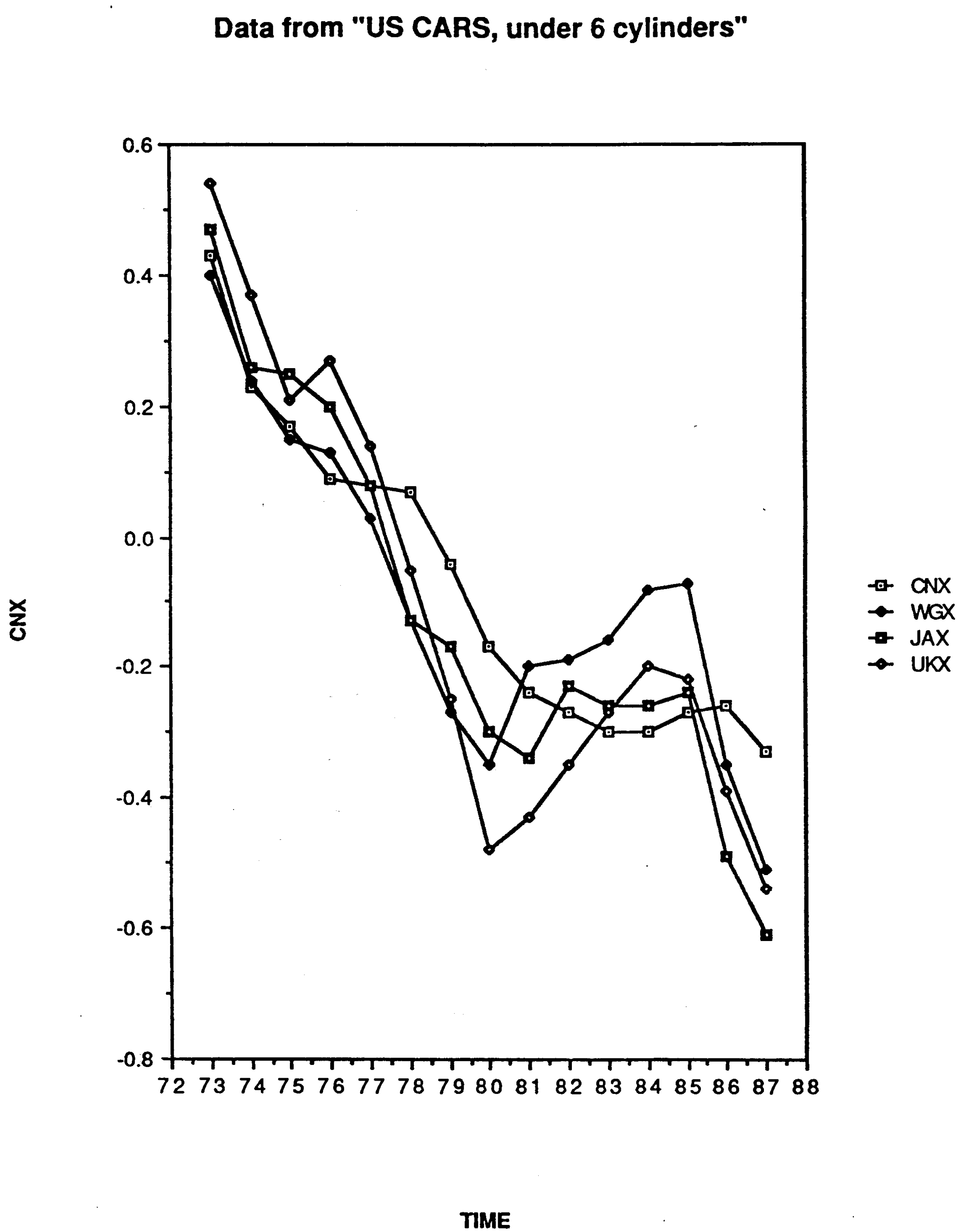


Figure 8


Data from "US CARS, under 6 cylinders"
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0.4
0.2
0.0
-0.2
-0.4
-0.6
-0.8
CNX
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88
TIME
CNX
WGX
JAX
UKX

Figure 9

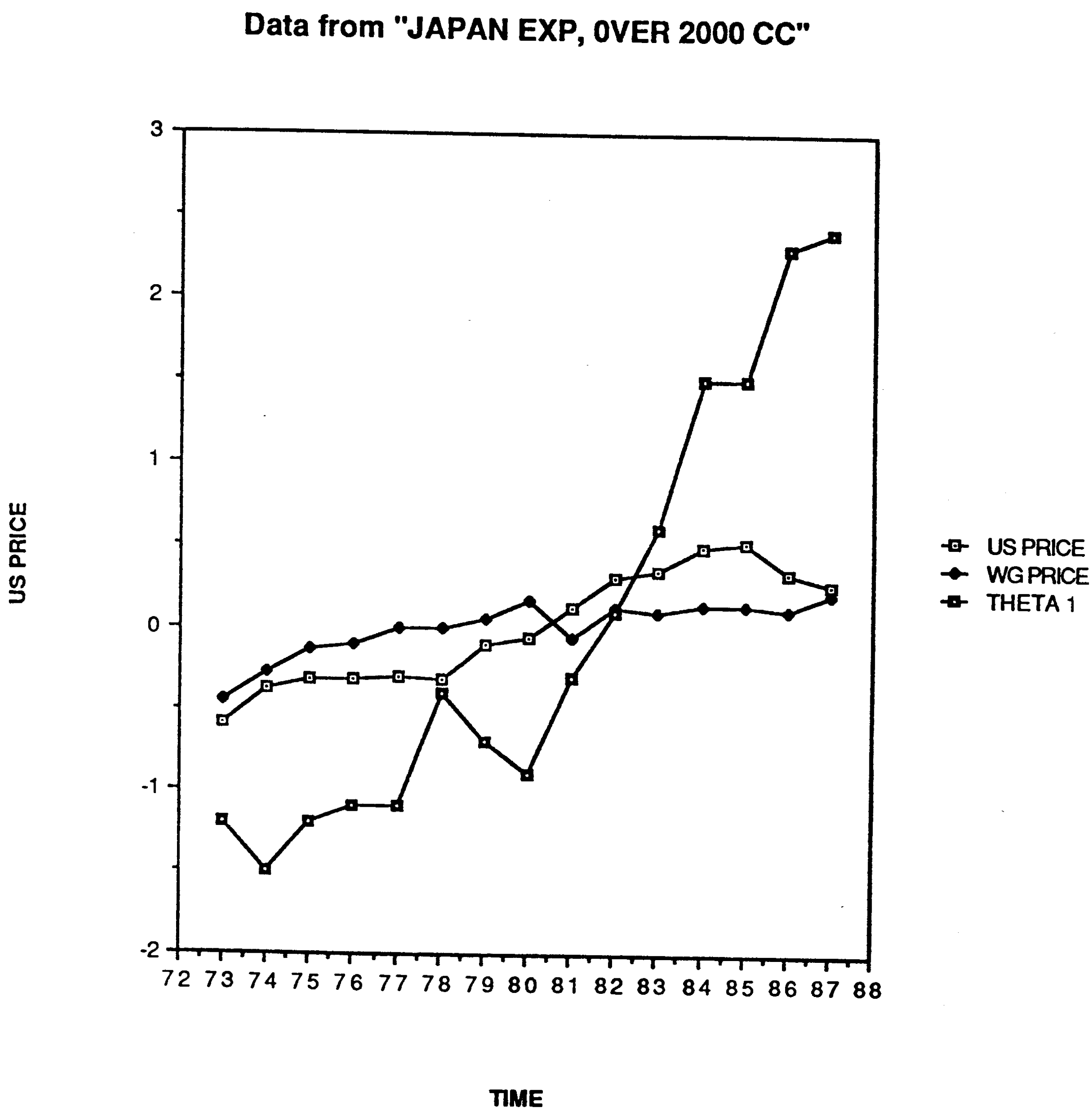

Data from "JAPAN EXP, 0VER 2000 CC"
US PRICE
3
2
1
0
-1
-2
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88
TIME
US PRICE
WG PRICE
THETA 1

Figure 10

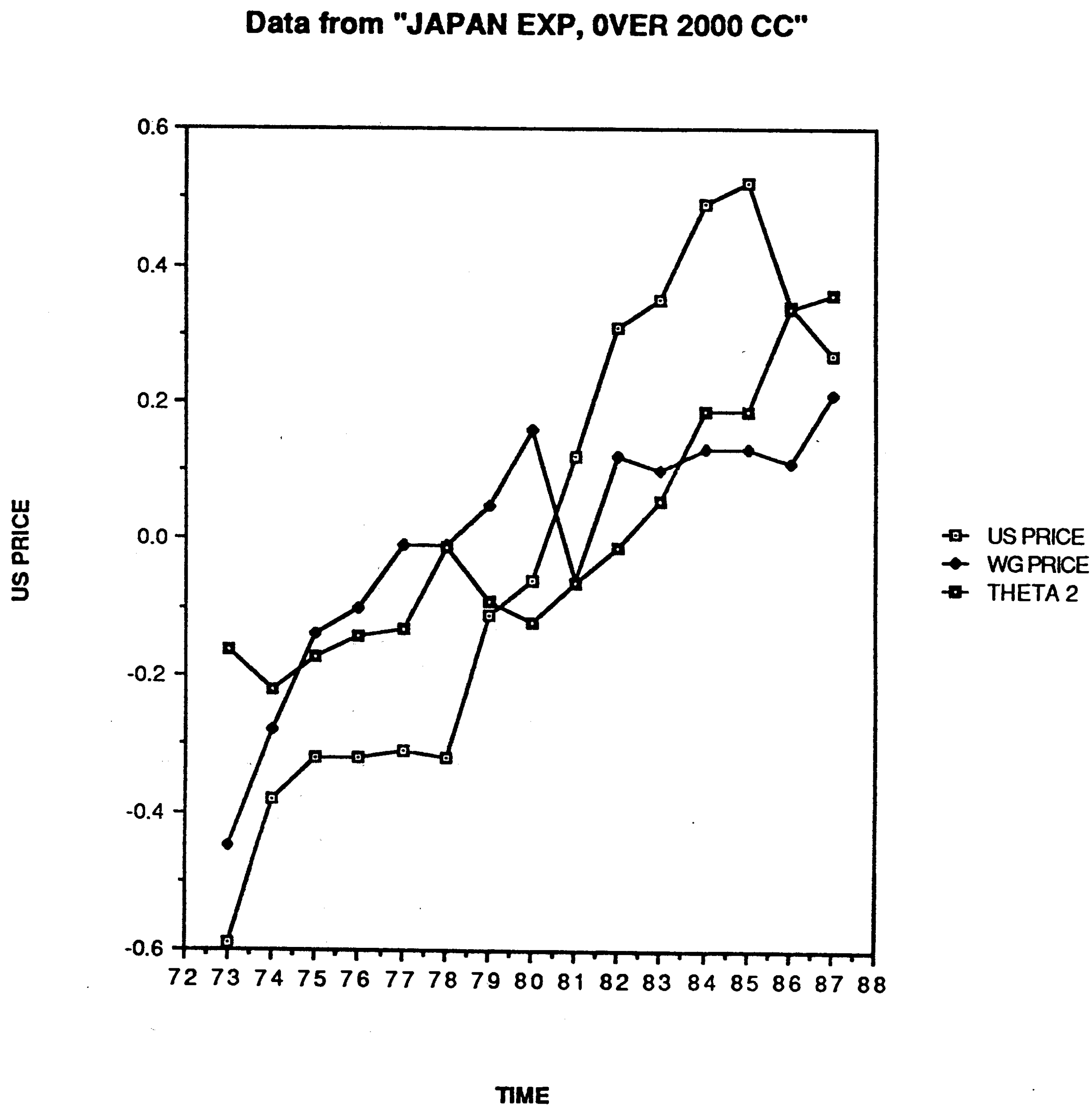

Data from "JAPAN EXP, 0VER 2000 CC"
0.6
0.4
0.2
0.0
-0.2
-0.4
-0.6
US PRICE
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88
TIME
US PRICE
WG PRICE
THETA 2

Figure 11

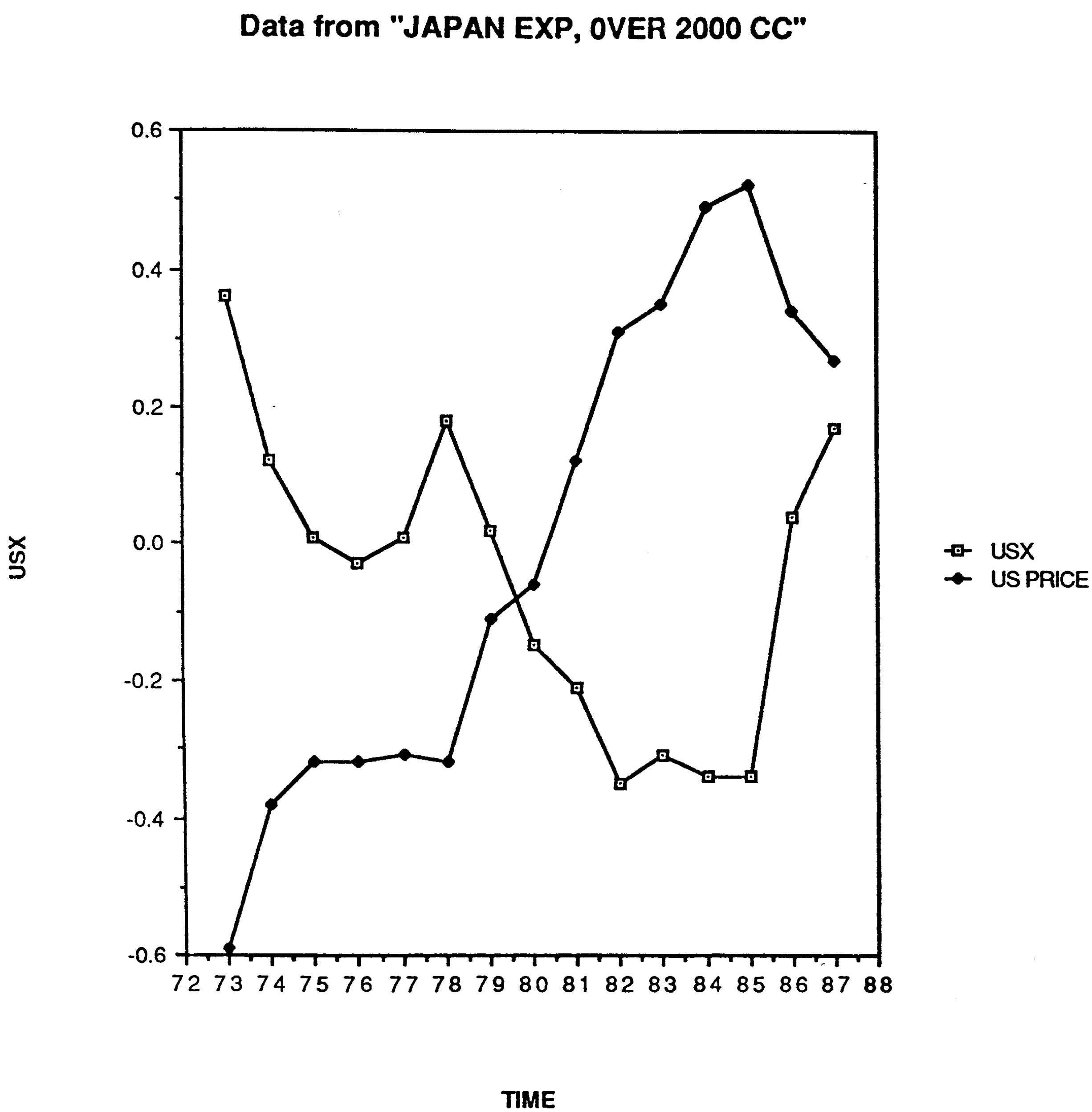
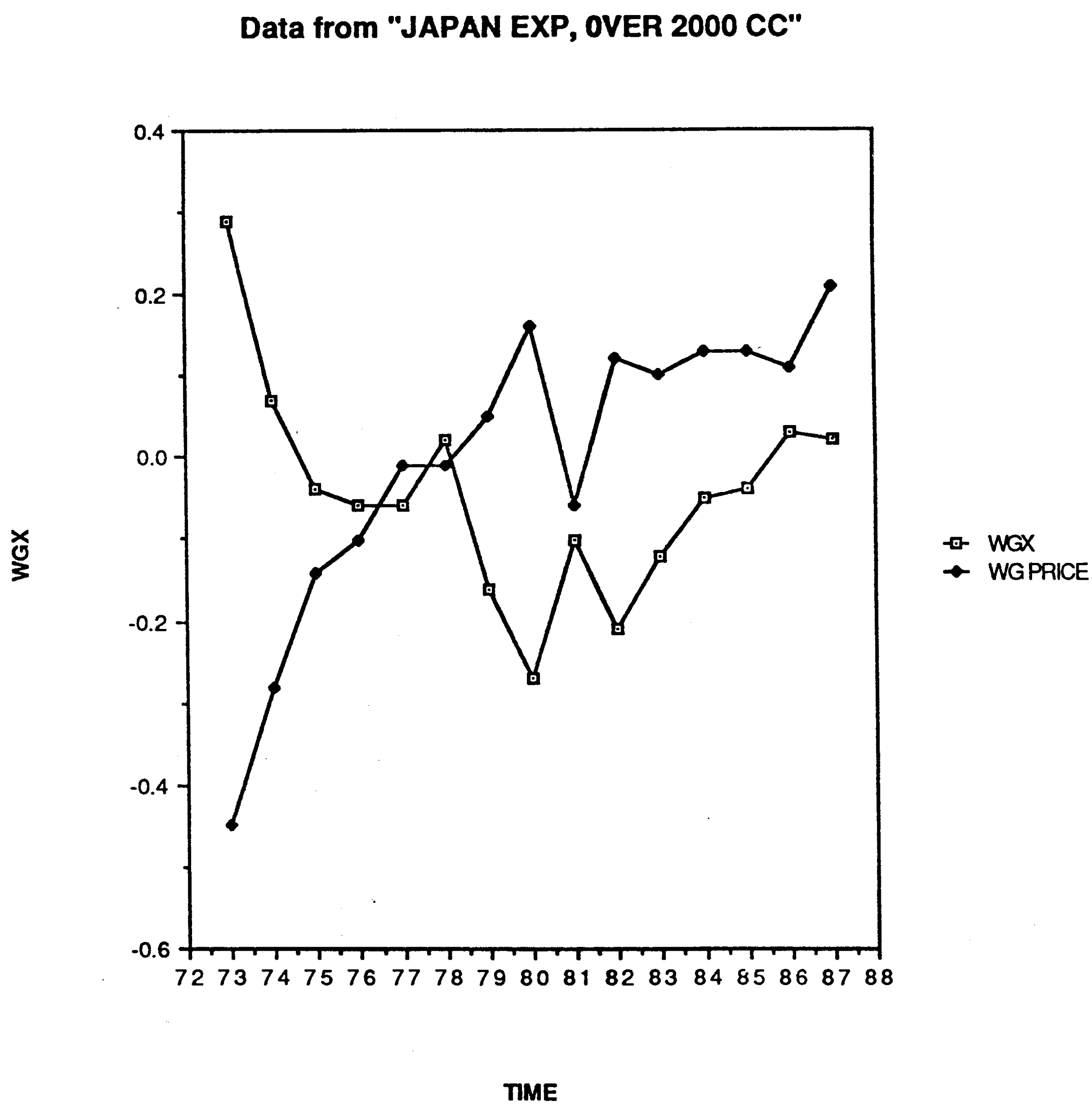

Data from "JAPAN EXP, 0VER 2000 CC"
0.6
0.4
0.2
0.0
-0.2
-0.4
-0.6
USX
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88
TIME
USX
US PRICE

Figure 12


Data from "JAPAN EXP, 0VER 2000 CC"
WGX
0.4
0.2
0.0
-0.2
-0.4
-0.6
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88
TIME
WGX
WG PRICE

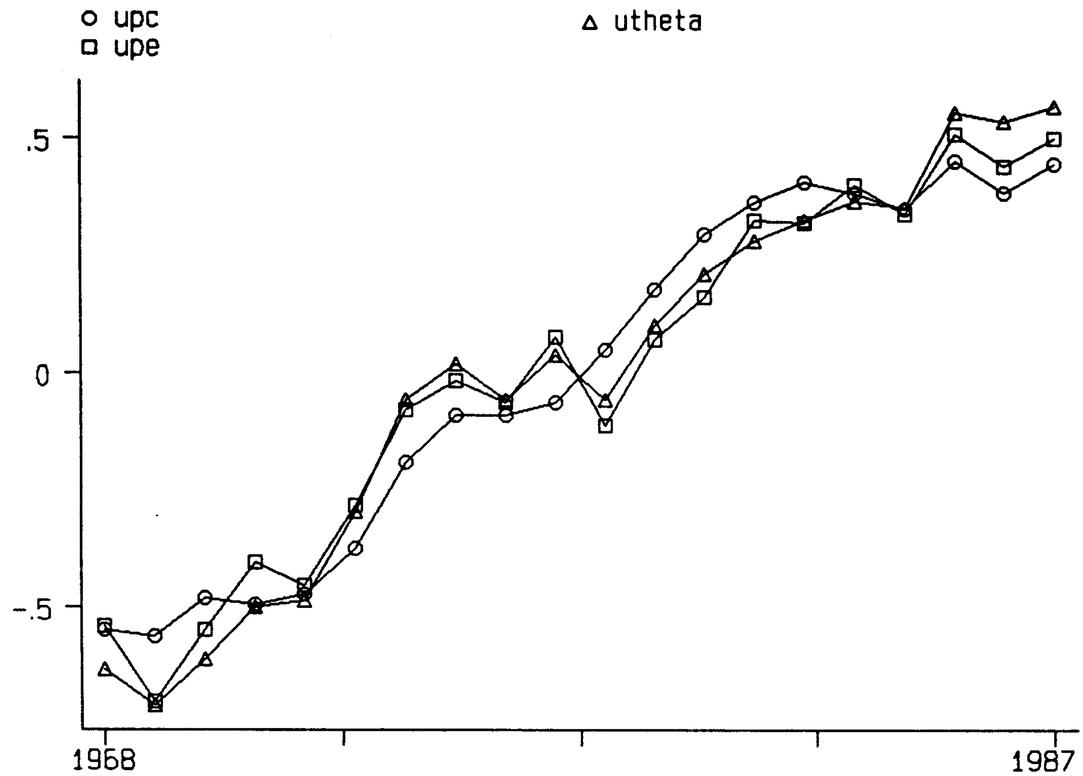


Figure 13
U.S. Total Merchandise Exports

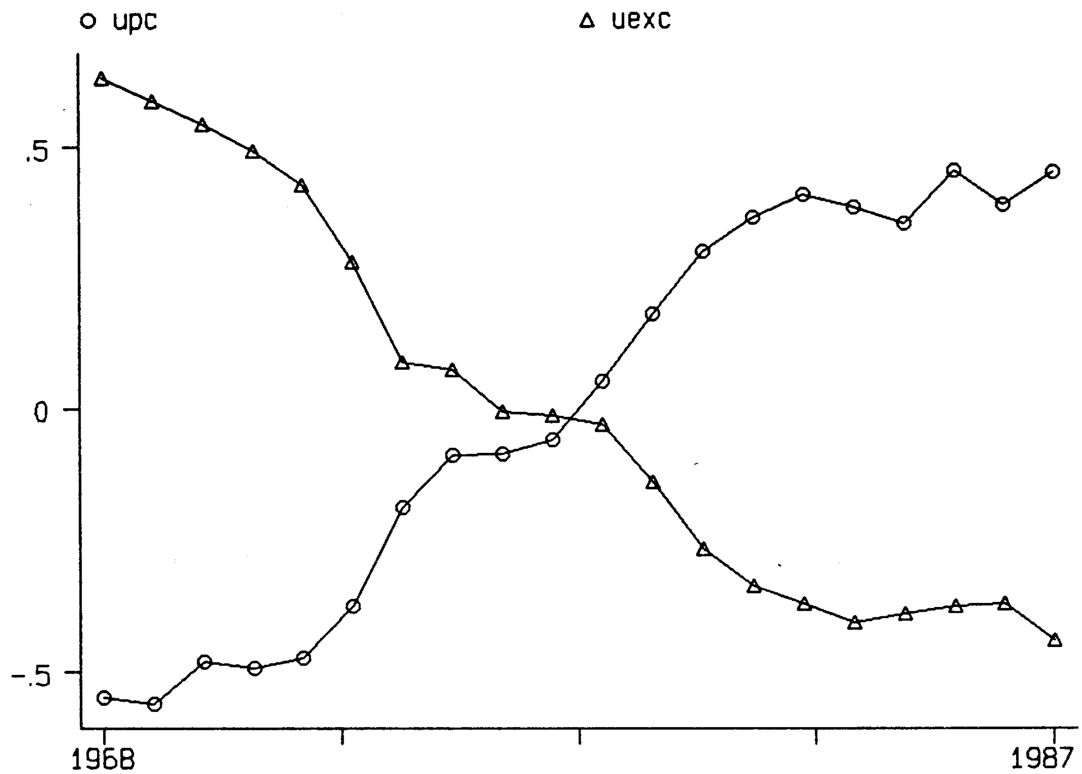


Figure 14
U.S. Total Merchandise Exports: Canada

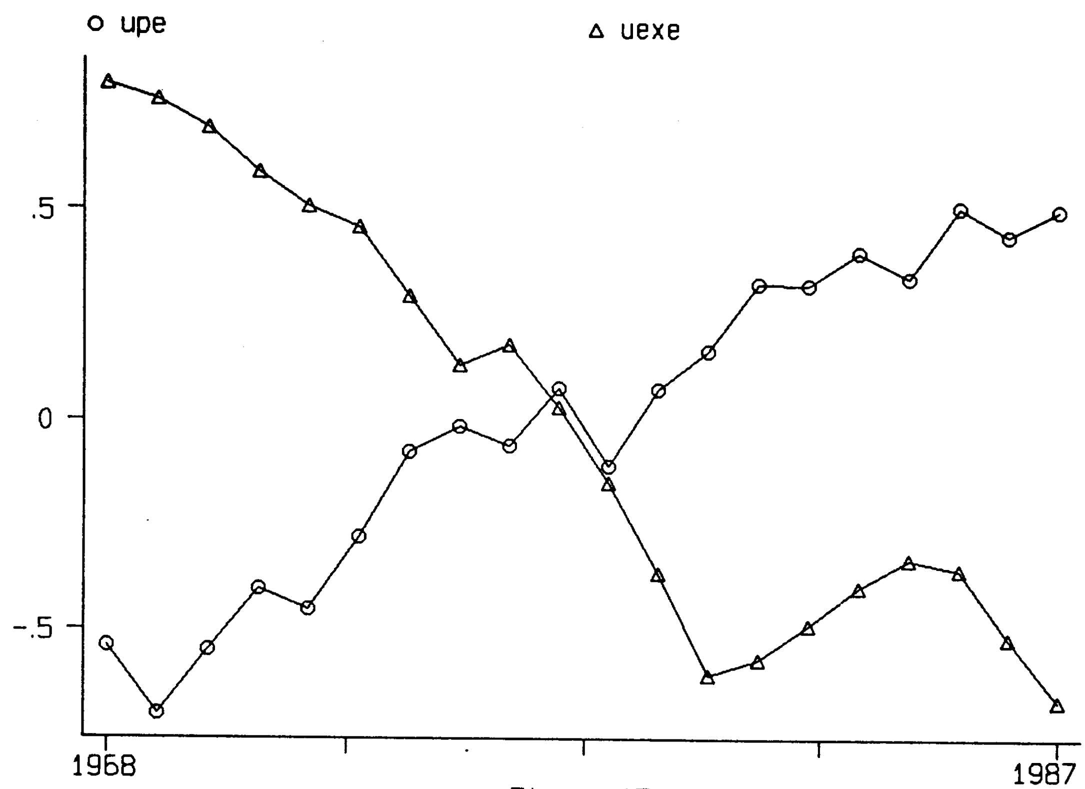


Figure 15
U.S. Total Merchandise Exports: U.K.

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