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SYMMETRIC PASS-THROUGH OF TARIFFS AND EXCHANGE RATES UNDER IMPERFECT COMPETITION: AN EMPIRICAL TEST

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ABSTRACT

This paper examines the effect of tariffs and exchange rates on U.S. prices of Japanese cars, trucks and motorcycles. In particular, we test whether the long run pass-through of tariffs and exchange rates are identical: the symmetry hypothesis. We find that this hypothesis is easily accepted in our sample. We also find that the pass-through relation varies across products, ranging from about 0.6 for trucks to unity for motorcycles. These coefficients have very different implications for trade policy. We explain the results based on demand, cost and institutional conditions in each industry. We also find weak evidence that the pass-through of exchange rates has fallen in more recent years.

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1. Introduction

The appreciation of the dollar during 1979-85 has led to renewed interest in the "pass-through" relationship between exchange rates and traded goods prices. A number of recent studies, including Woo (1984), Mann (1986), Dornbusch (1987), Giovannini (1987), and Krugman and Baldwin (1987), have analyzed the extent to which U.S. import prices failed to fall by as much as the value of foreign currencies. These authors generally appeal to some model of imperfect competition, whereby foreign producers increased their profit margins during the 1979-85 period. In contrast, after 1985 import prices have not risen by as much as the dollar depreciation, and foreign profit margins are being squeezed.

Concurrent with these studies, a body of literature has grown on the effects of trade policy under imperfect competition.[1] Katrak (1977), De Meza (1979), Svedberg (1979) and Brander and Spencer (1984) analyze how a foreign monopolist faced with a tariff will optimally change its producer price. If the producer price is lowered then the tariff is less than fully passed through in consumer prices, and the importing country experiences a terms of trade and welfare gain. In this case there is a "terms of trade" argument for import protection, but it is due to imperfect competition abroad rather than the traditional large country model.

In section 2 we demonstrate that a permanent change in exchange rates or ad valorem tariffs have identical effects on the consumer price of imports, with a foreign oligopolistic supplier. The purpose of our paper is to test this hypothesis using data for Japanese car, truck and motorcycle imports. A test of the symmetry hypothesis is of particular interest because there is abundant evidence that exchange rates are less than fully passed through in some markets.[2] If this result also holds for tariffs in our sample, we can appeal to symmetry to suggest that the result may hold for trade policy in other markets. This is a significant step towards establishing the empirical validity of the "terms of trade" argument for tariffs, under imperfect competition.

Our empirical specification is outlined in section 3 and results are presented in section 4. We find that the symmetry hypothesis is easily accepted for the tariff changes which occurred in truck and heavy cycle imports. For trucks the pass-through coefficient which applies to either a permanent change in exchange rates or the tariff is estimated as 0.582 (standard error of 0.062). In contrast, for heavy cycles we obtain a pass-through of unity or even higher. For cars the tariff did not change over the sample period, and the pass-through of exchange rates is 0.725. Thus, while we are able to accept the symmetry hypothesis, we observe a wide range of pass-through coefficients in our sample (this result is also found by Krugman, 1987 using German data). These coefficients have very different implications for trade policy, since a tariff coefficient which is less than (equal to) unity implies a nationally optimal tariff which is positive (zero). Some reasons for this variation in our results across industries are offered in section 4. We also examine whether the pass-through of exchange rates has fallen in more recent years, as suggested by Mann (1986) and Baldwin (1987), and find weak evidence to support this hypothesis. Conclusions are given in section 5.

2. Pricing Under Imperfect Competition

We consider a variant of the model used by Brander and Spencer (1984). The home currency price of the imported and domestic varieties of a good are denoted by p and q, respectively.[3] Import demand is $x(p,q,I)$ where I denotes income. The foreign firm maximizes expected profits in its own currency, treating q as exogenous.[4] We shall assume that its pricing decision must be made _before_ the exchange rate is known with certainty. As discussed by McKinnon (1979, chap. 4), Baron (1976) and Giovannini (1987), the foreign firm then faces a decision as to which currency to use in announcing its price. We shall not analyse this problem, but rely on the fact that 82 percent of Japanese auto sales to the U.S. are invoiced in dollars (Hamada and Horiuchi, 1987, Table 7.6). Since our empirical work deals with this market, we shall simply assume that the foreign firm sets its price in the domestic currency, i.e., chooses p.

Denote the foreign spot price of the domestic currency by s, which is random. Costs in the foreign currency are given by $C(x,w^*)$, where w^* denotes an _aggregate_ of foreign factor prices, and is treated as a scalar. Assuming that costs are homogeneous of degree one in factor prices, they can be written as $C(x,w^*) = \phi(x)w^*$, where $\phi' > 0$ (< 0) indicates rising (falling) marginal costs. The foreign firm's profit maximization problem can then be stated as,

$$\max_{p} E\{spx(p,q,I) - \phi(x)w^*\}, \tag{1}$$

where E denotes expected value. We assume that variables other than the exchange rate are non-random.[5] Letting $e = E(s)$ denote the _expected_ exchange rate, (1) can be rewritten as,

$$\max_{p} \{epx(p,q,I) - \phi(x)w^*\}. \tag{1'}$$

This certainty-equivalent structure to the firm's problem would not arise if the import prices were set in the foreign currency; see Baron (1976) and Giovannini (1987).

The first order condition for (1') is,

$$\phi'(x)(w^*/e) = p[1-(1/\eta)] \equiv r(p,q,I), \tag{2}$$

where $\eta = -x_p p/x$ denotes the (positive) elasticity of demand and $r(p,q,I)$ denotes marginal revenue. Assuming that $[\phi'' x_p(w^*/e) - r_p] \neq 0$, we can invert (2) to obtain the pricing equation,

$$p = \pi(w,q,I), \tag{3}$$

where $w = w^*/e$ denotes foreign factor prices in the domestic currency.

Equation (3) shall be estimated in the following sections. Our purpose here is to show how the elasticities of π depend on the underlying demand and cost functions. A general property of π is that it is <u>homogeneous of degree one</u> in its arguments. That is, increasing w, q and I by the same proportion would increase the optimal import price by that proportion. To see this, note that the demand function $x(p,q,I)$ must be homogeneous of degree zero in its arguments. It follows that the first derivative $x_p(p,q,I)$ is homogeneous of degree negative one. Then it is readily seen that if (p,w,q,I) satisfies the first order condition (2), then $\lambda(p,w,q,I)$ also satisifies this equation, for any $\lambda > 0$. Thus, π is homogeneous of degree one and this property can be empirically tested.

2.1 Changes in the Exchange Rate, Domestic Price, and Income

Totally differentiating (2), we obtain

$$(dp/dw)(w/p) = 1/[(\phi''x/\phi')\eta + (r_p p/r)]. \tag{4}$$

Recalling that $w = w^*/e$, this elasticity gives the change in the import price due to a change in the expected exchange rate or foreign factor prices. The second order condition for profit maximization implies that (4) is positive.[6] Its magnitude, however, depends on the underlying form of demand and costs.

The term $(r_p p/r)$ is the elasticity of marginal revenue with respect to price, which can be written from (2) as,

$$\begin{aligned}(r_p p/r) &= 1 + (p/r)(\eta_p p/\eta^2) \\ &\gtrless 1 \text{ as } \eta_p \gtrless 0.\end{aligned} \tag{5}$$

Along a linear demand curve $\eta_p > 0$ so that $(r_p p/r) > 1$, but the case $\eta_p < 0$ is certainly possible. The term $(\phi''x/\phi')$ is the elasticity of marginal cost with respect to output, and its sign is the same as that of ϕ''.

Summarizing results, the following sufficient conditions establish the magnitude of (4):

$$\eta_p > 0,\ \phi'' > 0 \Rightarrow 0 < (dp/dw)(w/p) < 1, \tag{6a}$$

$$\eta_p < 0,\ \phi'' < 0 \Rightarrow (dp/dw)(w/p) > 1. \tag{6b}$$

(6a) is the "normal" case where a change in the expected exchange rate (or foreign factor prices) is less than fully passed through in import price. It occurs, for example, when demand is close to linear and marginal costs are increasing.[7] However, the unusual case where a change in the expected exchange rate is more than fully passed through cannot be ruled out theoretically. It occurs, for instance, when the elasticity of demand is constant or decreasing in price, and marginal costs are declining.

Turning to the effect of a change in income, from (2) we calculate

$$(dp/dI)(I/p) = [(\phi''x/\phi')m - (r_I I/r)] / [(\phi''x/\phi')\eta + (r_p p/r)], \tag{7}$$

where $m = x_I I/x$ is the income elasticity of demand for the import. The denominator in (7) is the same as in (4), which is positive. To sign the numerator, consider the case where m is positive and constant for all prices. This implies that demand is of the general form $\ln x = \ln f(p,q) + m\ln I$, or $x(p,q,I) = f(p,q)I^m$. Calculating the elasticity of demand we obtain $\eta = -f_p p/f$ which is independent of income. Thus, marginal revenue does not depend on income and $r_I = 0$. The sign of (7) then depends on whether marginal costs are increasing or decreasing:

$$m > 0 \text{ constant}, \ \phi'' > 0 \Rightarrow (dp/dI)(I/p) > 0, \tag{8a}$$

$$m > 0 \text{ constant}, \ \phi'' < 0 \Rightarrow (dp/dI)(I/p) < 0. \tag{8b}$$

Lastly, we consider the elasticity of p with respect to the price of the domestic variety q. Since π in (3) is homogeneous of degree one its elasticities must sum to unity, and so

$$\begin{aligned}(dp/dq)(q/p) &= 1 - (dp/dw)(w/p) - (dp/dI)(I/p) \\ &= [(\phi''x/\phi')(\eta - m) + (p\eta_p/\eta^2)(p/r) + (r_I I/r)] \\ &\qquad / [(\phi''x/\phi')\eta + (r_p p/r)],\end{aligned} \tag{9}$$

using (4), (5) and (7). The magnitude of (9) can be inferred from our earlier results. For example, if (8a) holds then the first equality in (9) immediately implies that $(dp/dq)(q/p) < 1$, so an increase in the domestic price is less than fully matched in the import price. If (6a) and (8a) both hold and $\eta > m$, then the second equality in (9) implies that $(dp/dq)(q/p) > 0$. By also examining the converse case, we obtain the following results:

$$(6a),\ (8a) \text{ and } \eta > m \Rightarrow 0 < (dp/dq)(q/p) < 1, \tag{10a}$$

$$(6b),\ (8b) \text{ and } \eta > m \Rightarrow (dp/dq)(q/p) < 0. \tag{10b}$$

In (10a) a rise in the domestic price is less than fully matched by an increase in the import price. In contrast, under (10b) the import price would fall. This occurs, for example, when prices are a constant markup over marginal costs ($\eta_p = 0$) and the latter are declining as imports rise ($\phi'' < 0$ with $\eta > m$).

Note that in the analysis above we have treated the domestic price q as a parameter. It would not difficult to treat q as endogenous by adding a condition analogous to (2) for the domestic price. Then the effects of changing the expected exchange rate or factor prices would be obtained by comparative statistics on the two equation system. We do not report these results here, since our empirical specification will focus directly on the pricing equation (3), the properties of which we have characterized above.

2.2 Changes in the Ad Valorem Tariff

Suppose that an ad valorem tariff of τ is applied to imports. Let p denote the consumer price of imports, so the foreign firm receives $p/(1+\tau)$. With e still denoting the expected exchange rate, the foreign profit maximization problem can be stated as,

$$\max_p \{e[p/(1+\tau)]x(p,q,I) - \phi(x)w^*\}$$

$$= [e/(1+\tau)] \max_p \{px(p,q,I) - \phi(x)[w^*(1+\tau)/e]\}. \tag{2'}$$

The structure of (2') makes it clear that a change in the tariff $(1+\tau)$ has the same effect on the consumer price p as a change in (w^*/e). This point is also seen by writing the first order condition for (2') as $\phi'(x)[w^*(1+\tau)/e] = r(p,q,I)$, and inverting to obtain

$$p = \pi[w(1+\tau),q,I], \tag{3'}$$

where $w = w^*/e$ as before. Any changes in the tariff or expected exchange rate which have the same effect on $(1+\tau)/e$ will have identical pass-through on the consumer price of imports: this is the <u>symmetry</u> hypothesis.

Because of symmetry, the elasticity of the consumer import price with respect to $(1+\tau)$ is given by (4) and (6). From (6a), if the elasticity of demand is increasing in price (as with linear demand) and marginal costs are rising, then the tariff is less than fully passed through in the import price. This means that $p/(1+\tau)$, which is the foreign producer price, is reduced. This rise in the terms of trade corresponds to an increase in welfare for the domestic country (Katrak, 1977; Svedberg, 1979; Brander and Spencer, 1984). Thus, under conditions (6a) there is a "terms of trade" argument for import protection, due to oligopolistic pricing abroad.

If (6b) holds then nationally optimal intervention instead takes the form of an import subsidy (DeMeza, 1979; Brander and Spencer, 1984). The negative value of τ is more than fully passed through in the import price, so $p/(1+\tau)$ is reduced. This again corresponds to a rise in the terms of trade and increase in welfare for the importing country.

Note that the welfare results stated above apply when q is held constant and profits of the domestic firm are not affected by the tariff. If instead q is endogenous and profits change, then the welfare analysis is more complex. First, we must consider the impact of the tariff on profits of domestic firms, which depends on the assumed market structure, i.e., on the conjectural variations and possibility of entry (Brander and Spencer, 1981; Eaton and Grossman, 1986; Horstmann and Markusen, 1986). Second, we should recognize that a tariff which increases domestic output can reduce the divergence between marginal cost and marginal utility of the domestic variety (Eaton and Grossman, 1986). Third, we could consider the effect of the tariff on the number of varieties produced by domestic and foreign firms (Feenstra and Judd, 1982; Feenstra, 1988b). These various effects could be combined in a computable model, as done by Dixit (1988) for the U.S. auto industry. Our goal in this paper is more limited, however, and we shall just evaluate the "terms of trade" effect by estimating the pass-through of tariffs and the exchange rate, using the pricing equation (3').

3. Regression Specification and Data

We shall use a log-linear specification for (3'):

$$\ln p_t = c_t + \alpha \ln(w_t^*/e_t) + \beta \ln(1+\tau_t) + \gamma \ln q_t + \delta \ln I_t + \varepsilon_t, \tag{11}$$

where $c_t = c_0 + c_1 t + c_2 t^2$ is a time trend and ε_t is a random error. To estimate (11) we must specify how the expected exchange rate is determined. We shall suppose that the expected rate in each quarter is a log-linear function of the current and past quarterly-average spot rates:

$$\ln e_t = \sum_{i=0}^{k} \Theta_i \ln s_{t-i}. \tag{12}$$

The coefficients Θ_i in (12) would depend on the time-series properties of exchange rates. For example, if the spot rates follow a random walk then rational expectations would be formed with $\Theta_1 = 1$ and $\Theta_i = 0$ for $i \neq 1$. Other cases are considered by Frankel and Froot (1987). Including the current quarterly-average exchange rate s_t in (12) is meant to reflect information received by firms <u>within</u> a quarter which is then immediately reflected in prices. If our units of time were smaller, such as monthly or weekly, then omitting the current spot rate would be appropriate.[8]

Substituting (12) into (11), we shall estimate the pricing equation (3') as follows:

$$\ln p_t = c_t + \sum_{i=0}^{k} \beta_i \ln(w_t^*/s_{t-i}) + \beta \ln(1+\tau_t) + \gamma \ln q_t + \delta \ln I_t + \varepsilon_t, \tag{13}$$

where $\beta_i = \theta_i \alpha$ and $\sum_{i=0}^{k} \theta_i = 1$ is assumed. The errors in (13) can arise from inaccurate measurement of prices (see below) or price presetting for periods exceeding one quarter (see Giovannini, 1987). When estimating (13) we shall check for autocorrelation in ε_t. Quarterly dummies are also used when needed.

Our primary reason for estimating (13) is to test for symmetric pass-through of tariffs and a permanent change in exchange rates. This symmetry hypothesis is stated as

$$\beta = \sum_{i=0}^{k} \beta_i , \tag{14a}$$

which is a linear constraint on the coefficients in (13). We also wish to test that the pricing equation is homogeneous of degree one, which can be stated as

$$\sum_{i=0}^{k} \beta_i + \gamma + \delta = 1. \tag{14b}$$

This homogeneity test is a check for the overall specification of the pricing equation, and is analogous to testing that a conventional demand system is homogeneous of degree zero in prices and income. It will be of particular interest to test (14a) and (14b) simultaneously.

Of secondary interest in (13) is the time pattern of the coefficients β_i, which indicate the rate at which exchange rate changes are passed though in import prices. To reduce erratic behavior of the coefficient estimates, we shall use a second-order polynomial lag on β_i, so that $\beta_i = a + bi + ci^2$. For $k = 4$, as we shall use, we can readily calculate that

$$\sum_{i=0}^{4} \beta_i = 5a + 10b + 30c. \tag{15}$$

Substituting (15) into (14) shows how symmetry and homogeneity can be tested when using the polynomial lag.

We shall estimate the pricing equation separately for U.S. imports of Japanese cars, compact trucks and heavy motorcycles (greater than 700 cc). The sample period lies between 1974:1 and 1987:1, depending on data availability. Japanese trucks experienced an increase in their tariff from 4 to 25 percent, effective August 21, 1980. Heavy motorcycles had a tariff of 45 percent imposed on April 16, 1983, falling annually to 35, 20, 15 and 10 percent, and ending in October 1987.[9] The ad valorem tariff on Japanese cars did not change during the sample period, so this regression is estimated to observe the pass-through of exchange rates without testing symmetry.

The most important features of the data are summarized here. The import prices are either wholesale (c.i.f.) unit-values inclusive of duty, or Divisia indexes of several dissaggregate wholesale unit-values. They are obtained from U.S. Bureau of the Census (1974-1987). In addition to the well known problems with using unit-values, these data series suffer from not correcting for quality change. The most pronounced quality change occurred in Japanese cars (Feenstra 1984, 1988a), which were subject to a "voluntary export restraint" since April 1981. This restraint changes the nature of the optimal pricing decision for importing firms, so (3') no longer applies. For this reason we omit the period of the trade restraint when estimating the car regression. Other upgrading which may have occurred in any of the products can be reflected in the time trend c_t and random error ε_t in (13).

Among the independent variables, the spot exchange rate (measured as a quarterly average) shows great fluctuation, ranging between 300 and 150 yen/dollar during the sample period. The factor price aggregates w_t^* are measured as the Japanese domestic wholesale prices for each product, available from the Bank of Japan (1974-1986). They are very stable over time. For cars and trucks a quarterly price deflator for U.S. absorption (q_t), and total U.S. expenditure on each product (I_t), are available from the U.S. Dept. of Commerce (1974-1987). For cars we also included the unit-value of German imports (from U.S. Bureau of the Census) as another competing price. Since the variables q_t and I_t are endogenous, all the car, truck and cycle regressions were estimated with instrumental variables.[10]

For heavy cycle imports, we had two sources of unit-value data. The first was interview data reported by the U.S. International Trade Commission (1983, Table 8, and 1983-1984). The advantage of this data is that it gives the unit-value of imports for consumption, inclusive of duty, for the major Japanese importers (Honda, Suzuki and Yamaha). However, the disadvantage is that the data ends in 1984:4, and it includes German heavy cycles (imported by BMW) within the reported unit-value.[11] A second source was unpublished data from the U.S. Dept. of Commerce giving the unit-value of import shipments, distinguishing Japanese and German heavy cycles up to 1987:1.[12] The disadvantage of this data is that import shipments include sales to inventories, which are very erratic in quantity and to a lesser extent in unit-value. These data had to be adjusted to include the tariff.[13] We shall use both sources of import price data when estimating the heavy cycle regression. Finally, data on the price of U.S. heavy cycles was not available, so the U.S. price of steel was used as a proxy.

4. Estimation Results

To determine the appropriate lag length for the exchange rate, we first estimated (13) with 2, 3, 4 and 5 unconstrained lags. The sum of the estimated coefficients β_i from this exercise are reported in Table 1. These estimates measure the total pass-through of the exchange rate on import prices. The coefficients for cars are significantly less than unity, while in contrast, the coefficients for heavy cycles are close to one. For trucks and cycles the pass-through estimates reach a maximum with $k = 4$ quarterly lags. While there is no generally accepted technique for choosing the length for a polynomial lag (see Judge et al., 1980, Chap. 5), our approach in this paper is to make a pass-through of less than unity "prove itself" by including too many lags rather than too few. We experimented with a lag length of 3 and 4 quarters, but since the results were quite similar, we report below only the results for a second-order polynomial lag with $k = 4$. Note that using the polynomial lag leads to β_i coefficents which are not significantly different from the unconstrained estimates.

In Table 2 we report the estimates of (13) without imposing the symmetry or homogeneity constraints (14). Quarterly dummies were used in the motorcycle regressions, but were not needed for cars or trucks since their t-statistics in those cases were less than unity. The time trend in (13) was used in each of the regressions, but its coefficients and those of the quarterly dummies are not reported.[14] For heavy cycles the "pooled" regression in the last column simply stacks the "consumption" and "shipments" regressions, which differ in their measurement of the import price (see section 3).

Considering the results in Table 2, the pass-through coefficients range from 0.627 for trucks to about unity for cycles. These results differ slightly from Table 1 since the polynomial lag was not used there. It can be seen that the pass-through is quite rapid since the β_i estimates are nearly zero for the third lag, and also for the second lag except in cars. The coefficents on the final lag move away from zero in some cases, especially for cycles (shipments). When a fifth lag is added to the cycle regressions (without using the polynomial constraint) then the estimates of β_i for the fourth lag are still positive and high, but are zero for the fifth lag. Thus, there is some curious annual effect in the cycle data, whereby the import price depends on the exchange rate one year ago.

Turning to other coefficients in Table 2, the elasticity of the import price with reprect to the tariff is 0.570 for trucks and between 0.949 and 1.388 for cycles. These estimates are within two standard errors of the pass-through coefficients for the exchange rate, which suggests that the symmetry hypothesis will be accepted. The coefficients of the U.S. price differ substantially in magnitude over the products, but are insignificant. The coefficients of the German price, and income, are all highly insignificant. Note that the Durbin-Watson statistics are surprisingly good.

We also ran the regressions while separating w_t^* from the exchange rate terms. The coefficents we obtain on w_t^* vary substantially in magnitude (ranging from -1 for cars to 12 for cycles), but in most cases are insignificantly different from those in Table 2. If we omit w_t^* entirely, the estimated pass-through of exchange rates are similar to those in Table 2.

4.1 Test of Symmetry and Homogeneity

In Table 3 we report estimates of (13) when the symmetry and homogeneity constraints (14) are imposed. For cars and trucks the exchange rate pass-through changes slightly, while the coefficients of U.S. and German prices improve considerably. In contrast, for heavy cycles the coefficients of U.S. and German prices are erratic and insignificant. The pass-through of exchange rates and the tariff is between 0.971 and 1.272.

In Table 4 we report the calculated F-statistics testing whether the symmetry and homogeneity constraints (14) can be accepted. Note that because the regressions are estimated with instrumental variables, the F-statistics need not be positive, as observed for the homogeneity constraint in cars.[15] They are asymptotically distributed as $\chi^2(R)/R$ when the constraints hold, where R is the number of restrictions. Looking down Table 4, we see that the symmetry and homogeneity constraints are easily accepted for the individual products. In the second-last test, we report the F-statistic testing whether the cycle (consumption) and cycle (shipments) data can be pooled. We accept this hypothesis and, conditional on it, also accept symmetry and homogeneity in the pooled cycle regression.

Acceptance of the homogeneity constraint supports the overall specification of our regressions. Acceptance of the symmetry constraint is consistent with the model presented in section 2, and supports our prior beliefs. Applying this results to other markets, it means that the response of import prices to exchange rates can be used to predict the effect of changes in tariffs. This result could be quite useful for the analysis of trade policy. However, our results also show that the pass-through differs substantially across industries, and in particular, is not always less than unity.

To explain these results, consider first the tariff coefficient for trucks. In Feenstra (1988b) the price and characteristics of trucks sold by American and Japanese producers are studied in detail. Prior to the tariff increase, nearly all compact trucks were produced by Japanese firms, some of which were marketed through American automobile companies. After the tariff was raised to 25 percent in August 1980, U.S. producers quickly introduced their own compact truck models, but with very similar characteristics to existing Japanese models. In this environment we would expect price competition to be intense, and Japanese firms would be very reluctant to pass through the full amount of the tariff. This is consistent with our estimated tariff coefficient of 0.582.

In contrast, for heavy cycles we estimate tariff coefficients of unity or even higher. We offer two explanations for this result. First, the tariff increase in April 1983 was applied to both imports from Japan <u>and</u> heavy cycles produced in the U.S. by Honda and Kawasaki. The latter producers operate plants in foreign trade zones (FTZ) in the mid-west. When a good is sold from a FTZ into the U.S., the producer can normally pay the U.S. tariff on either the final product or imported parts, whichever is less. However, for the case of heavy cycles, Honda and Kawasaki had to pay the final tariff on their U.S. sales from the FTZ (see U.S. International Trade Commission, 1987, p. A-6 and Appendix E). With this rise in U.S. prices (except for Harley-Davidson), it is not too surprising that the Japanese exporters (Honda, Suzuki and Yamaha) would pass-through much of the tariff.

A second important feature of this industry is the dramatic drop in production following the tariff: total production in Japan was cut in half from 1982 to 1984, 1985 or 1986 (U.S. International Trade Commission, 1987, Table 22). The reason is that U.S. exports accounted for about 60% of Japanese production prior to the tariff, but these exports dropped to about 15% of their former quantity after the tariff. Note that U.S. consumption of Japanese heavy cycles did not drop by this much, because many U.S. sales were made from previously accumulated inventories.

In order to cover fixed costs with this drop in production, it is quite possible that Japanese firms would have to raise their prices, which would cause consumer prices to increase by more than the tariff. From (4), this is a profit maximizing response when $\eta_p = 0$, implying that producer prices are a constant markup over marginal costs, and $\phi'' < 0$ so that marginal costs rise with the fall in output. This scenario is consistent with the tariff coefficient exceeding unity in our cycles (shipments) regression. However, we cannot make a strong case that the tariff was more than fully passed through, since the estimated coefficient in any of the cycle regressions is insignificantly different from one.

4.2 Pass-Through in Various Time Periods

The large swings in the value of the dollar during the 1980's has led to the hypothesis that the pass-through of exchange rates to U.S. import prices may have fallen. For example, if foreign firms believe that the recent depreciation of the dollar is temporary, they could resist raising their dollar prices too much, hoping to retain market share. Several models in which this phenomenon can occur are analysed by Krugman (1987), and empirical evidence supporting the hypothesis for aggregate U.S. import prices is presented in Mann (1986) and Baldwin (1987). Without making any strong prior case that this phenomenon might apply to our sample of industries, we can certainly check whether the pass-through has fallen.

In Table 5 we report the total pass-through of exchange rates (i.e., sum of β_i coefficients) for trucks and heavy cycles during various sub-periods. We did not impose symmetry or homogeneity in these regressions, so the pass-through coefficients can be compared with Table 2. We also show in Table 5 the F-statistics and critical values for testing whether the observations used there are drawn from the same population as Table 2 (i.e., whether the pass-through coefficients in the two tables are insignificantly different).[16]

For trucks we obtain a pass-through of 0.576 for the 1977-84 period, before the dollar began to depreciate. This is slightly lower, but insignificantly different, than the coefficent of 0.627 for the entire 1977-87 period. However, if we instead check for a structural break in the truck regression between 1977-80 and 1981-87, when the dollar began its appreciation, we obtain coefficents of 0.679 and 0.434, respectively. Thus, there is weak evidence that the pass-through relationship fell, though the change is not significant by the F-test. For cycles (shipments) we estimate a pass-through which is insignificantly higher during 1978-84, and lower during 1981-87, than for the entire 1978-87 period. In cycles (pooled) we obtain a coefficient of 1.051 over 1978-84, which compares with 0.886 over 1978-87, again giving weak evidence that the pass-through of exchange rates has fallen.

5. Conclusions

In this paper we have used data on U.S. imports of Japanese cars, trucks and motorcycles to estimate the pass-through of exchange rates and tariffs. For trucks we have found a pass-through of about 0.6. This means that the increase in the tariff from 4 to 25 percent in August 1980 raised consumer prices by an estimated 13 percent, and lowered Japanese producer prices by about 8 percent. In contrast, for heavy cycles we found a pass-through of about unity, so the tariff increase in April 1983 and subsequent decreases had little effect on Japanese producer prices. These results have very different implications for trade policy. In trucks, the drop in the producer price corresponds to a terms of trade gain. While there are several other factors which should be considered before evaluating welfare (see section 2.2), this is a first step toward establishing a gain for the U.S. For heavy cycles, the constant producer prices means that the tariff led to a conventional deadweight loss.

However, our purpose in this paper has not been to judge the efficacy of trade policy in specific industries. Instead, we have evaluated more generally a model of imperfect competition and trade, in which foreign firms respond to tariff and exchange rate changes by adjusting their producer prices. This model led to the estimating equation (3'), which performed quite well on the industry data. In particular, we were able to accept the hypothesis of symmetric pass-through of tariffs and exchange rates. This means that the response of import prices to exchange rates can be used to predict the effect of changes in tariffs. But the variation in our results across industries means that empirical evidence is needed in each case, and that we cannot make general statements about the extent of pass-through.

Footnotes

1. In contrast to this paper, much of the literature deals with a "profit shifting" motive for import protection. See Brander and Spencer (1981), Dixit (1984) and Eaton and Grossman (1986). Horstmann and Markusen (1986) argue that the "profit shifting" motive disappears when entry is free, but a "terms of trade" motive persists, as analyzed here.
2. See Dunn (1970), Clark, Logue and Sweeney (1974), Isard (1977), Kravis and Lipsey (1977), Richardson (1978), Schembri and Robicheau (1986) and the papers mentioned above.
3. For expositional convenience, we consider only a single variety of each imported and domestic good. The case of many varieties of each is examined in Feenstra (1986), and the results are basically identical to those reported below.
4. Thus, we are assuming the foreign and domestic firms act as Betrand competitors, whereas Brander and Spencer (1984) consider Cournot competition. Our results are very similar, which demonstrates that they do not depend on the market structure assumed. In contrast, see Eaton and Grossman (1986) who analyze how the "profit shifting" motive for import protection depends on the market structure.
5. More general sources of uncertainty are examined in Giovannini (1987).
6. To see this, differentiate (1') with respect to p obtaining $H(p)x_p/e = 0$, where $H(p) = r(p,q,I) - \phi'(x)w$ as in (2). The second order condition is then $H'(p)x_p/e - H(p)x_{pp}/e < 0$. Since $x_p < 0$ and $H(p) = 0$ from (2), the second order condition reduces to $H'(p) > 0$, which can be rewritten as $(4) > 0$.

7. More generally, we can think of $\eta_p > 0$ as specifying the class of demand curves which are less convex (more linear) than constant-elasticity curves, while $\eta_p < 0$ specifies the class which are more convex.
8. With quarterly data an alternative procedure would have been to include the spot rate at the end of the last quarter in (12), but this rate suffers from daily random error. We could also have included the forward rate in (12), but from the results of Meese and Rogoff (1983) this is not a better predictor of the exchange rate than the lagged spot rate.
9. The tariff protection was meant to expire in April 1988, but ended early at the request of Harley-Davidson. The reasons for protection in each of these industries and background data are provided by Hufbauer, Berliner and Elliot (1986).
10. The instrumental variables were the U.S. price of steel (from U.S. Bureau of Labor Statistics, Producer Price Indexes), wages in car or truck parts and body manufacturing (from U.S. Bureau of Labor Statistics, Employment and Earnings), U.S. consumer prices, U.S. private consumption, U.S. treasury bill rate, German aggregate wages, German consumer prices, and the current and two lags of the quarterly-average mark/dollar exchange rate (all from International Monetary Fund, International Financial Statistics).
11. More disaggregated data were not reported to preserve business confidentiality. Before 1983 German imports are less than five percent of Japanese imports (by value or quantity of quarterly shipments), but during and after 1983 this figure ranges between 20 and 35 percent.

12. These data were kindly provied by Juanita Kavalauskas of the U.S. International Trade Commission, who also helped with my many questions. All data used in this study are available on request.

13. Note that the tariff rates only applied to Japanese imports of heavy cycles exceeding certain quotas (ranging from 6,000 to 10,000 units depending on the year). Thus, for shipments exceeding the quota, the unit-value of imports was multiplied by the appropriate tariff.

14. The time trends did not show any consistent pattern, and were often insignificant.

15. The SSR shown in Tables 2 and 3, and used in the F-tests, equals $(y-Xb)'(y-Xb)$ where b is the vector of estimated coefficients using instrumental variables, but X is the matrix of actual dependent variables before they are regressed on the instruments. This SSR will not necessarily rise when constraints are imposed on b.

16. The second test we perform for trucks is a conventional Chow test, where the entire sample period is split between two sub-periods. However, in other cases we did not have sufficient observations to estimate the coefficients in both of the sub-periods. In those cases the SSR in the shorter sub-period is zero, with the number of coefficients exactly equal to the number of observations. The F-statistic is then calculated as described in the notes to Table 5.

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Table 1. Pass-Through of Exchange Rates with Various Lag Lengths

		Number of Lags			
Sample	Period	2	3	4	5
Cars	1974:1-1981:1	0.797	0.712	0.707	0.723
		(0.081)	(0.082)	(0.087)	(0.108)
Trucks	1977:1-1987:1	0.521	0.581	0.625	0.592
		(0.062)	(0.072)	(0.082)	(0.092)
Cycles	1978:1-1984:4	0.740	0.835	0.907	0.827
(Consumption)		(0.444)	(0.327)	(0.385)	(0.426)
Cycles	1978:1-1987:1	0.687	0.780	1.061	0.832
(Shipments)		(0.309)	(0.362)	(0.614)	(0.772)

Notes:

Standard errors are in parentheses.

Table 2. Unconstrained Regressions, Dependent Variable: Import Price

	Cars	Trucks	Cycles (Consump.)	Cycles (Shipments)	Cycles (Pooled)
Period	74.1-81.1	77.1-87.1	78.1-84.4	78.1-87.1	78.1-87.1
N, K[a]	29,9	41,9	28,13	37,13	65,13
$\bar{R}^2$	0.988	0.989	0.907	0.769	0.833
SSR[b]	0.0159	0.0191	0.0635	0.4720	0.5890
w_t^*/s_t	0.444* (0.104)	0.282* (0.057)	0.288 (0.256)	0.798 (0.722)	0.447* (0.212)
w_t^*/s_{t-1}	0.316* (0.043)	0.139* (0.030)	0.172* (0.095)	-0.042 (0.282)	0.104 (0.093)
w_t^*/s_{t-2}	0.166* (0.080)	0.061 (0.050)	0.117 (0.146)	-0.335 (0.568)	-0.031 (0.156)
w_t^*/s_{t-3}	-0.0079 (0.053)	0.047 (0.028)	0.124 (0.084)	-0.083 (0.234)	0.042 (0.087)
w_t^*/s_{t-4}	-0.205* (0.102)	0.098 (0.079)	0.192 (0.232)	0.715 (0.823)	0.324 (0.224)
Exchange Rate[c]	0.713* (0.100)	0.627* (0.081)	0.893* (0.362)	1.053* (0.564)	0.886* (0.216)
Tariff		0.570* (0.138)	0.949* (0.219)	1.388* (0.296)	1.129* (0.155)
U.S. Price	1.002 (0.934)	0.029 (0.399)	0.682 (0.601)	1.143 (2.171)	0.572 (0.588)
German Price	0.084 (0.090)		0.056 (0.107)	0.124 (0.227)	0.063 (0.105)
Income	-0.026 (0.117)	-0.032 (0.059)	-0.227 (1.693)	-0.215 (0.648)	0.016 (0.010)
Durbin-Watson	2.43	1.75	2.73	1.69	--

Notes:

* Significant at 95% level. Standard errors in parentheses.

a N is the number of observations and K the number of independent variables. Coefficients for time trends and quarterly dummies are not reported.

b Sum of squared residuals.

c Sum of coefficients for w_t^*/s_{t-i}, $i=0,1,\ldots,4$. w_t^* is an aggregate of foreign factor prices, and s_{t-i} is the spot exchange rate (yen/$).

Table 3. Constrained Regressions, Dependent Variable: Import Price

	Cars	Trucks	Cycles (Consump.)	Cycles (Shipments)	Cycles (Pooled)
Period	74.1-81.1	77.1-87.1	78.1-84.4	78.1-87.1	78.1-87.1
N, K[a]	29,8	41,7	28,11	37,11	65,11
$\bar{R}^2$	0.989	0.990	0.916	0.785	0.834
SSR[b]	0.0155	0.0196	0.0642	0.4754	0.6083
w^*_t/s_t	0.419* (0.096)	0.285* (0.055)	0.329* (0.175)	0.753* (0.321)	0.488* (0.171)
w^*_t/s_{t-1}	0.318* (0.041)	0.124* (0.026)	0.184 (0.081)	0.061 (0.176)	0.154* (0.087)
w^*_t/s_{t-2}	0.181* (0.075)	0.040 (0.044)	0.116 (0.136)	-0.189 (0.284)	0.017 (0.139)
w^*_t/s_{t-3}	0.0081 (0.047)	0.032 (0.023)	0.127 (0.080)	0.0052 (0.139)	0.078 (0.080)
w^*_t/s_{t-4}	-0.202* (0.098)	0.101 (0.074)	0.215 (0.188)	0.640 (0.422)	0.337 (0.201)
Exchange Rate[c]	0.725* (0.095)	0.582* (0.062)	0.971* (0.152)	1.272* (0.276)	1.075* (0.153)
Tariff		0.582* (0.062)	0.971* (0.152)	1.272* (0.276)	1.075* (0.153)
U.S. Price	0.259* (0.149)	0.397* (0.071)	0.606 (0.508)	-0.161 (0.383)	-0.083 (0.180)
German Price	0.115 (0.079)		0.064 (0.096)	-0.025 (0.171)	-0.0092 (0.097)
Income	-0.099 (0.071)	0.021 (0.026)	-0.641 (0.575)	-0.086 (0.330)	0.018* (0.010)
Durbin-Watson	2.46	1.70	2.75	1.51	--

Notes: See Table 2.

Table 4. Tests of Symmetry and Homogeneity

Sample	Restrictions (Number)	Calculated F[a]	0.90 Critical Value[b]
Cars	Homogeneity (1)	-0.503	2.71
Trucks	Symmetry and Homogeneity (2)	0.419	2.31
Cycles (Consumption)	Symmetry and Homogeneity (2)	0.083	2.31
Cycles (Shipments)	Symmetry and Homogeneity (2)	0.086	2.31
Cycles (Pooled)	Cross-equation (13)	0.565[c]	1.52
Cycles (Pooled)	Symmetry and Homogeneity (2)	0.852	2.31

Notes:

[a] $F = [(SSR_r - SSR_u)/R]/[SSR_u/(N_u - K_u)]$, where SSR_u, N_u, K_u are from Table 2, SSR_r is from Table 3, and R is the number of restrictions.

[b] $\chi^2_{0.9}(R)/R$.

[c] SSR_u, N_u, K_u are from the fifth column of Table 2, while SSR_r is from the sum of the third and fourth columns.

Table 5. Pass-Through of Exchange Rates in Various Time Periods

Sample	Period	Pass-Through[a]	SSR (N,K)[b]	Calculated F[c]	0.90 Critical Value[d]
Trucks	77.1-84.4	0.576 (0.187)	0.0172 (32,9)	0.282	1.63
	77.1-80.4	0.679 (0.363)	0.0138 (41,17)	1.152[e]	1.68
	81.1-87.1	0.434 (0.091)			
Cycles (Shipments)	78.1-84.4	1.475 (0.706)	0.2713 (28,13)	1.233	1.63
	81.1-87.1	0.544 (1.776)	0.8989 (25,13)	-0.475	1.54
Cycles (Pooled)	78.1-84.4	1.051 (0.326)	0.4227 (56,13)	1.88	1.63

Notes:

a Standard errors in parentheses.

b SSR is the sum of squared residuals, N is the number of observations and K the number of independent variables.

c $F = [(SSR_r - SSR_u)/R)]/[SSR_u/(N_u - K_u)]$, where SSR_u, N_u, K_u are from above (column 3), SSR_r, N_r are from Table 2, and $R = (N_r - N_u)$.

d $\chi^2_{0.9}(R)/R$.

e In this case $R = (K_u - K_r)$.