Exchange Rate Pass-Through to Manufactured Import Prices: The Case of Japan

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Abstract

This paper examines the exchange rate pass-through to yen based manufactured import prices of Japan using asymmetric unit root and cointegration tests and asymmetric models. Due to sticky prices, for example, there are reasons to believe that the degree of pass-through depends on whether the exchange rate appreciates or depreciates. The sample used in this study covers the period January 1975 to June 1997. Using two state regime switching models, the estimated pass-through coefficients corresponding to appreciation and depreciation of the currency are found to be 98 percent and 83 percent respectively; these coefficients are shown to be significantly different, particularly in the post recession period. Moreover, we have shown that the recession in Japan in the 1990s has significantly affected the exchange rate pass-through relationship particularly when the yen depreciates and that the proposition that exchange rate depreciation and appreciation have systematic asymmetric effects on exchange rate pass-through coefficient. Forcing appreciations and depreciations to have the same effects on the import prices does not appear to uncover the true underlying exchange rate pass through relationship.

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Key words: Exchange rate pass-through, Japan, threshold autoregression.

1. Introduction

Modelling and estimating the relationship between the exchange rates and import/ export prices have been given considerable attention during the last two decades[1]. One of the major reasons which led to this increased attention is the adoption of a flexible exchange rate system by many countries since the collapse of the Bretton Woods system in 1973. Countries adopt a flexible exchange rate system to remedy the balance of payment crises they encounter. However, outcomes of the actions taken by the countries to depreciate the currencies were disappointing since they were unable to attain the desired results, refuting the theoretical proposition that depreciation of a currency would boost exports and discourage imports, leading to an equilibrium in the balance of trade. In the case of Japan, although the value of its currency was appreciating, it had a significant trade surplus against the US for a long period of time indicating that the changes in exchange rates were not fully passed through to import prices.

The objectives of this study are to (i) test whether the exchange rate pass-through to import prices is complete, (ii) understand the factors influencing the import prices, (iii) examine whether the import price responds aymmetrically to positive and negative changes in exchange rates and other variables, (iv) calculate a long-run measure of exchange rate pass-through for positive and negative changes in exchange rates and (v) ascertain whether exchange rate policy can be used to achieve balance of trade equilibrium of Japan.

According to the theory, a flexible exchange rate policy is expected to remedy the balance of payments crises, although empirical evidence is against it. For example,

[1] See for example, Spitaeller, 1980; Khosla and Teranishi,1989; Dwyer, Kent and Pease, 1994; Parsley, 1995, Menon,1996; Kenny and McGetton, 1998; Webber, 2000; Gil-Pareja, 2000.

Japan, enjoyed a significant trade surplus for a long time while the Japanese yen was appreciating against the US dollar.

Earlier approaches to exchange rate pass-through relationship emanated from estimating the import/export demand and supply elasticities (Branson ,1972). Several studies found that incomplete pass-through is very common in the short-run and it does not carry through to the long-run (Blejer and Hillman, 1982). Accordingly, complete exchange rate pass-through is a long-run phenomenon. In theory, incomplete pass-through is mainly due to market structure and product differentiation (Menon, 1996). In a perfectly competitive market where imported and domestically produced goods are perfect substitutes, the measurement of pass-through is similar to that of elasticities approach. When imperfect competition exists, the firms can charge a mark-up on their costs to earn above normal profits even in the long-run. This mark-up can vary depending on the degree of substitutability between the domestically produced goods and imported goods as determined by the product differentiation and the degree of market integration and separation.

In order to examine the theoretical propositions on the degree exchange rate pass-through, there has been a plethora of empirical studies using various methodologies and data series across different countries during the past three decades. We discuss some of these studies and the results. Menon (1993) studied the exchange rate pass-through to import prices of motor vehicles using Engle-Granger's two-step cointegration tests and error correction models and found that exchange rate pass-through to import prices of motor vehicles is incomplete even in the long-run. He has put forward two possible explanations for the incomplete pass-through: the presence of quantitative restrictions and the pricing practices on intra-firm sales by multinational firms. Dwyer, Kent and Peace (1994), on the other hand, found that in the long-run exchange pass-through over the docks is complete for both Australian imports and manufactured exports. However, they found that responses to currency movements differ significantly in the short-run. Pass-through to import prices is found to be more rapid than that to manufactured export prices. See also Dwyer and Lam (1994).

Parsley (1995) studied the US imports of ball bearings, bolts, nuts, screws and bicycles and showed that in general expectations of future exchange rates should be a determinant of pass-through and concluded that the pass-through may change when there are changes in expected future exchange rates. Lutz and Reilly (1997) used data for twelve European countries and found that the exchange rate pass-through is below 50 percent for all these countries. Their study further revealed that relative market share of domestic firms or imposition of non-tariff barriers has no relationship with the low degree of pass-through, while the study by Adolfson (1997) indicated the pricing to market behaviour in the majority of industries is the reason for limited pass-through[2].

Taking studies on exchange rate pass-through in the literature and econometric methodologies developed recently together, we believe that a part of the reason for the lack of empirical support for the exchange rate pass through is that previous empirical studies did not incorporate some well-known stylised facts into modelling and estimating the relationship between the import prices and the exchange rates. Our study attempts to (i) address issues such as asymmetric adjustment of the individual variables in testing for unit roots and cointegration, (ii) examine whether the response

[2] See Lee (1997) for a study on the pass-through of exchange rates to various Korean industries, Kenny and McGettigan (1998) for the relationship between import price and exchange rate in Ireland, and Gross and Schmitt (1999) for the exchange rate pass-through to import prices of small and medium sized automobiles in Swiss data. Webber (2000) found significant long-run estimates of import pass-through in seven out of nine countries in the Asia-Pacific region.

of the import price indices to positive and negative changes in exchange rates is asymmetric; and (iii) use a new model for the long-run exchange rate pass-through to examine the asymmetric response of import price to changes in exchange rates. See Enders and Granger (1998) for methodologies for testing asymmetric unit roots, cointegration of variables and others. Since the Japanese economy was in recession since 1989, we will examine its effects on the exchange rate pass through relationship. Such a study will undoubtedly contribute to the available vast literature on exchange rate pass-through relationship, and more importantly, to the debate between the US and Japan with regard to Japanese trade surplus against the US even when its currency was appreciating.

The paper is organized as follows: Section 2 briefly explains the development of the exchange rate pass-through model. The methodology used in this study is discussed in Section 3. Measurement of the variables used in this study is given in Section 4. Section 5 discusses the data series used and their time series properties. The empirical analysis and the results are reported and discussed in Section 6 and the final Section gives the conclusion of the study.

2. The Basic Model of Exchange Rate Pass-Through

Exchange rate pass-through can be defined as the extent to which changes in nominal exchange rates are reflected in import prices. This depends on various factors such as product differentiation and the nature of competition. Speaking broadly, exporters set their price (PX^*) at a profit margin (λ) over the cost of production (C^*). Therefore, the import price (PM) can be defined as follows:

$$PM = PX * ER = (1 + \lambda)C * ER \qquad (1)$$

where (*ER*) is the exchange rate. Now, set $(1+\lambda) = \rho$, where ρ is the profit mark-up. It has been hypothesised that exporters base their pricing decisions on competitive pressures in the domestic market, which are proxied by the gap between the prices of domestic import competing goods (*PD*) and exporters' cost in domestic currency (Hooper and Mann, 1989). The profit mark-up can thus be modelled as:

$$\rho = \{PD/(C * ER)\}^{\alpha} \tag{2}$$

Substituting ρ in equation (2) for $(1+\lambda)$ in equation (1) and taking logarithms of variables (denoting them in lower case letters), we can derive the equation for the import price as follows:

$$pm = \alpha pd + (1-\alpha)c* + (1-\alpha)er \tag{3}$$

The pass-through coefficient (1-α) is expected to be between 0 and 1. If the foreign firm is a price taker in the Japanese competitive market, then α = 1 and therefore, the pass-through is zero. In this case, it can be seen that the Japanese import price set by foreign firm is equal to the Japanese domestic price and changes in exchange rates and foreign costs have no effect on import prices. If the foreign firms do not face any competition in the Japanese market, both the changes in the exchange rates and the foreign costs are fully passed through to import prices leaving the mark-up unchanged. In this case α = 0. We will test these restrictions in Section 6. A trend variable will be included in the above model to capture the changes in the variables due to improvement in the technology over time.

3. Methodology

Since growing evidence suggests that time series are inherently non-linear, we assume that adjustments of variables to their respective equilibriums are asymmetric, depending on whether the short-term fluctuations from the long-run equilibriums are positives or negatives. Using the methodology developed in Enders and Granger

(1998), we test for the possibility of asymmetric adjustments of time series used in this study. Further, we compare the time series properties derived from the asymmetric models with those obtained by forcing the variables to have symmetric adjustments, that is, symmetric models.

Dickey and Fuller (1979) (DF) assume linearity and symmetric adjustment. In other words, in such models the null hypothesis of unit roots is tested against the alternative hypothesis of a symmetric adjustment of the variable of interest. However, recent studies have found that important economic and financial variables display asymmetric adjustment paths (see, for example, Nefti, 1984; Falk, 1986 and Enders and Granger,1998). Therefore, there is a possibility that the well-known DF type unit root tests reject the null hypothesis of integration due to an incorrect alternative hypothesis. To overcome this problem, Enders and Granger (1998) generalized the DF test by allowing for an asymmetric adjustment of variables. There are two major alternative models, which were proposed by Enders and Granger to deal with asymmetric adjustment paths of variables. These are known as the threshold autoregressive (TAR) model and the momentum threshold autoregressive (MTAR) model, depending on either the lagged levels or the lagged differences of the variables of interest are used in defining the heaviside indicator function.

Now, we briefly discuss the adjustment process of a time series variable, y_t. Consider the model:

$$\Delta y_t = I_t \rho_1 y_{t-1} + (1 - I_t) \rho_2 y_{t-1} + \varepsilon_t \tag{5}$$

where I_t is the heaviside indicator function, which can take one of the following four forms:

$$I_t = \begin{cases} 1 \text{ if } y_{t-1} \geq 0 \\ 0 \text{ if } y_{t-1} < 0 \end{cases} \tag{6}$$

or

$$I_t = \begin{cases} 1 \text{ if } y_{t-1} \geq a_0 \\ 0 \text{ if } y_{t-1} < a_0 \end{cases} \tag{7}$$

or

$$I_t = \begin{cases} 1 \text{ if } y_{t-1} \geq a_0 + a_1(t-1) \\ 0 \text{ if } y_{t-1} < a_0 + a_1(t-1) \end{cases} \tag{8}$$

or

$$I_t = \begin{cases} 1 \text{ if } \Delta y_{t-1} \geq 0 \\ 0 \text{ if } \Delta y_{t-1} < 0 \end{cases} \tag{9}$$

The models with indicator functions (7) and (8) are known as TAR models. When equation (7) is used to set the heaviside indicator function, it is assumed that the long-run equilibrium occurs at point $y_t = a_0$, provided $-2 <(\rho_1 , \rho_2)<0$. In this case, if $\rho_1 = \rho_2 = 0$, then the series is a pure random walk. When the heaviside indicator function is defined according to (8) and if $-2 <(\rho_1 , \rho_2)<0$, the trend line $y_t = a_0 + a_1 t$ is an attractor such that the (y_t) time series is trend stationary. When y_{t-1} is above the trend line, the time series tends to decay at the rate of ρ_1 and when it is below y_{t-1}, the time series tends to decay at the rate of ρ_2. If either ρ_1 or ρ_2 lies outside the interval (-2,0), then the sequence may not be trend stationary.

In the heaviside indicator function (6), it is assumed that the $y_t = 0$ is the long-run equilibrium value of the time series. Therefore, if $y_{t-1} > y_t$, then the adjustment is $\rho_1 y_{t-1}$, and if $y_{t-1} < y_t$, then the adjustment is $\rho_2 y_{t-1}$,. This type of TAR model can capture aspects of "deep" movements in a series. The models using the heaviside indicator functions in equations (6) and (9) are known as MTAR models. Replacing heaviside indicator function (6) with (9) is useful when the adjustment is asymmetric in such a way that the series exhibits more momentum in one direction than the other (Enders and Granger, 1998).

Equation (5) can be augmented with the lagged changes in the y_t to overcome any problem due to serial correlation in the error term, ε_t. In order to select the optimal lag length, the diagnostic checks such as Ljung-Box test and various model selection criteria such as AIC and/or BIC can be used.

If the unit root tests indicate that the variables are integrated of the same order, there may exist a cointegrating relationship between the variables of interest. If the variables are cointegrated, the long run relationship arising from the following equation should be stationary or *I(0)* (Engle and Granger, 1987):

$$pm_t = \beta_0 + \beta_1 er_t + \beta_2 c_t + \beta_3 c_t^* \tag{10}$$

In order to examine the long-run asymmetry of import prices to exchange rate changes, a new variable is constructed as in Webber (2000). Being in line with Webber, the exchange rate at any point in time *t* can usually be expressed as follows:

$$er_t = er_0 + er_t^A + er_t^D$$

where er_t is the initial exchange rate, $er_t^A \equiv \sum_{i=1}^{t} \theta_i (er_i - er_{i-1})$, $\theta = 1$ for $\mathrm{er_i} < er_{i-1}$ and $\theta_i = 0$ for $\mathrm{er_i} > er_{i-1}$, and $er_t^D \equiv \sum_{i=1}^{t} \theta_i^* (er_i - er_{i-1}), \theta_i^* = 1$ for $\mathrm{er_i} > er_{i-1}$. Accordingly, the variable er_t^A represents the accumulated sum of appreciation episodes and the variable er_t^D represents the accumulated sum of depreciation episodes. To test for long-run asymmetry, only the data series relating to one of the episodes has to be included in the exchange rate pass-through equation. Hence only the series for depreciation episodes is included in the tests for long-run asymmetry. Therefore, our new long-run exchange rate pass-through equation is:

$$pm_t = \beta_0 + \beta_1 er_t + \beta_2 er_t^D + \beta_3 c_t + \beta_4 c_t^* + u_t \qquad (11)$$

If pm_t er_t , c_t c_t^* and er_t^D are cointegrated, then there exists a long-run relationship among the variables. In the above equation, the long-run exchange rate pass-through coefficients are $(\beta_1 + \beta_2)$ and β_1, corresponding to depreciation and appreciation respectively. Therefore, the restriction that the long-run depreciation pass-through is equal to zero ($\beta_2 = 0$) is the test against the long-run asymmetry of the import price to exchange rate changes, which is tested within both the Engle and Granger and Johansen (1991,1995) cointegration frameworks.

If pm_t er_t , c_t and c_t^* are co-integrated, then the short-run asymmetry is tested by examining the following asymmetric error-correction model (ECM):

$$\begin{aligned}\Delta pm_t = \phi + \sum_{i=1}^{k-1}\omega_i \Delta pm_{t-i} + \sum_{i=1}^{k-1}\theta_i\, \Delta^+ er_{t-1} + \sum_{i=1}^{k-1}\theta_i\, \Delta^- er_{t-1} + \sum_{i=1}^{k-1}\gamma_i \Delta c_{t-i} \\ + \sum_{i=1}^{k-1}\lambda_i \Delta c_{t-1}^* + \varphi\varepsilon_{t-1} + \upsilon_{\mathrm{t}}\end{aligned} \qquad (12)$$

where Δ is the difference operator, ε_{t-1} is a one period lagged error term from the cointegrating equation and υ_t is a white-noise error term.

Whether the response of the import price changes to positive and negative changes in the exchange rate is asymmetric is tested using the Wald-type F-test. For comparison purpose, the above asymmetric error correction model and a symmetric or a standard error correction models are also estimated.

4. Measurement of Variables

In this section, we discuss how the variables used in the models were constructed. The variables constructed in this section are domestic import price, exchange rate index, foreign cost index and domestic production cost index.

Domestic Import Price

The price index for the manufactured imports is used as a proxy for the import price variable. Several studies used various proxies for the import price variable, including unit value of imports and the wholesale price index of trading partner countries. These proxies suffer from various limitations: unit values are suitable only when they are applied to a single product rather than to a group of heterogeneous products, thus when the products are not similar this is not a suitable proxy for the import price as the unit values reflect the changes of prices of dissimilar commodities. Further, the changes in commodity composition of imports make the unit values an inappropriate proxy for import price**.** Wholesale price indices themselves suffer from certain limitations. In their construction, non-traded goods are also taken into account and they even include the changes in costs of goods, which are not produced for exporting. Further, these price indices are constructed using domestic weights rather than international weights. Japan is among the few countries (along with Australia, Germany and the US) for which import price indices are available for an adequate length of time. Since these indices are based on transaction prices, they are free from limitations faced by many previous studies that used aforementioned proxies for the import price index.

Being consistent with the Standard International Trade Classification (SITC) of the United Nations for manufactured commodities, the import price index for manufactured imports of Japan is constructed by combining the import indices for textiles, metals and related products, wood, lumber and related products, chemicals, machinery and equipment using the respective weights that were utilised in constructing the overall import price index of Japan. Therefore, the import price index for the manufactured goods (*pm*) is constructed using the following formula:

$$pm = \sum_{i}^{k} w_i \, pm_s$$

where, w_s is the weight assigned to sector *s*, pm_s is the import price index of sector *s*. Two import price indices are constructed as above: one on the basis of yen and the other on the basis of the contractual currency.

Exchange Rate Index

In this study, we use a new Exchange rate index based on the contractual currency composition of imports to Japan. Athukorala and Menon (1994) have been the first to use a similar index of exchange rates in their study on pricing to market behaviour and exchange rate pass-through to export prices of Japan. Their index has been constructed by deflating the export price index in yen by an index representing the contractual currency price of exports. A similar exchange rate index reflecting the actual currency composition of imports is used in this paper which is constructed by deflating the domestic currency import price index by the contractual currency based import price index. An exchange rate index based on contractual currency prices better reflects real currency composition of imports/exports than the nominal effective exchange rate index, which was extensively used in many previous studies on exchange rate pass-through to import/export prices. Moreover, the nominal effective exchange rate may distort the reality by assigning higher weights to certain currencies on the basis of import shares although such a proportion of imports may not have been invoiced in these currencies. A single contractual currency based exchange rate index for the total manufactured imports was constructed by using the following formula:

$$er_t = \frac{pm}{\sum_i^k w_i pm_i^C}$$

where er_t is the contractual currency based exchange rate index, w_i is the weight assigned to contractual currency import price index for category i, pm is the yen based import price index for total manufactured imports explained above, and pm_i^C is the contractual currency import price index for category i.

Foreign cost index

The Foreign cost index was computed as:

$$c^* = \sum_{i=1}^k w_i c^*{}_i$$

where c^* is the foreign cost index, w_i is the share of manufactured imports to Japan from country i, and c^*_i is a proxy for cost of production of country i. To proxy the foreign production cost, producer price index for manufacturing was used. When producer price index for manufacturing was not available for the respective country, wholesale price index was used as a proxy for the production cost. Details of proxies used to represent foreign production cost are given in Appendix C.

Index of Domestic production cost

Producer price index for the manufacturing sector of Japan was considered as a proxy for the domestic cost of production.

5. Data series and the time series properties

The monthly data series used in this study spans the period 1975:01 to 1997:06. In order to examine how the recession in Japan – experienced since 1989 - has affected the exchange rate pass through relationship, the data series was divided into two sub-sample periods: 1975:01 to 1989:12 and 1990:01 to 1997:06. We believe that this relationship was affected during the recession, particularly when the yen based import price index is used in the empirical analysis. Figure 1 in Appendix B shows the time series behaviour of the variables used in this study. Figure 1 depicts that exchange rate and the yen based import price index follow each other closely. Overall there was an appreciation of the exchange rate till March 1995, then there is a dip in the exchange rate around April 1995, indicating a sudden increase in the value of yen. It is clear from the graph that the yen based import price index has responded quickly to the change in the exchange rate with a delay of two months. Subsequently, the exchange rate was depreciating till April 1997. This observed pattern is also reflected in the import price index. Foreign cost index has been on the increase during the whole sample period. However, the import price has not followed the changes in foreign costs indicating that foreign exporters absorb the changes in cost of production into profit margins. The Japanese cost of production shows it has been stable, indicating that over the long-run fluctuations in cost of production are stationary. Therefore, the exchange rate is expected to emerge as the main determinant of the yen based import prices.

Table 1 report the results of the unit root tests against TAR/MTAR adjustments for levels of the variables. As shown in column seven in Table 1, none of the variables was found to be stationary in their levels. However, the first differences were found to be stationary. Further, none of the variables shows any asymmetric adjustment towards its equilibrium point (constant (μ) or trend attractor (T). Note that, in the absence of asymmetric adjustments of the variables, the ADF test was shown to be a more powerful test than the asymmetric unit root tests. The ADF test results reported in Table 2 are consistent with those reported in Table 1.

6. Empirical Analysis and the Results

Cointegration test results using the unit root tests against TAR and MTAR adjustment are given in Table 3. The test for a unit root in the residuals from the cointegrating equation (long-run exchange rate pass-through equation) shown in columns two and four in the table rejects the null hypothesis of no cointegration at the one percent level. However, the results reported in columns three and five do not reject the hypothesis of symmetric adjustment of the cointegrating residuals towards their long-run equilibrium point. It is evident that the conventional unit root tests fail to detect the long run relationship, while the asymmetric unit root tests uncover such relationship. If the long run relationship is undetected, there is a possibility of misspecification of the model in the subsequent analysis.

We carried out the analysis for the full and two sub-samples with and without imposing the asymmetry, and the results are reported in Tables 4 and 5 . The results of the symmetric long-run pass-through equation suggest that all the independent variables are significant at the one percent level except in one case. That is in the second sub-sample period only the exchange rate index is significant when the Johansen method is used. Estimation results for the asymmetric exchange rate pass-through equation reported in Table 5 indicate that the exchange rate variable for depreciation episodes is significant only in the second sub-sample period. The appreciation exchange rate pass-through coefficients vary between 62 percent and 114 percent for the EG method and the 98 percent and 157 percent for the Johansen method among the three sample periods used. When depreciation pass-through coefficients are considered they vary between 48 percent and 76 percent for the EG method and 18 percent and 124 percent for the Johansen method among the three sample periods used. The hypothesis that the coefficient of the asymmetry variable is zero is reported in Table 7 and is rejected at one percent level only during the second sub-sample period for the EG method. However, the Johansen method rejects the same in all the sample periods. The stability of the exchange rate pass-through coefficient was examined by estimating exchange rate pass-through equations recursively and the time varying pass-through coefficients are plotted in Figure 2 in Appendix B with their two standard error bands. The Figure shows that the pass-through coefficients were increasing since October 1980 till March 1981 and then decreasing till the end of June 1997. During the entire sample period, the exchange rate pass-through coefficients varied between 0.674 and 1.202.

To examine the significance of the short-run dynamics of the exchange rate in the pass-through relationship, two error correction models were estimated. The results for the error-correction models are reported in Table 6. According to the results of the asymmetric error-correction model, only positive changes in exchange rate (depreciation) are significant determinants of import prices during the total sample period while this variable and the negative changes were insignificant in the two sub-sample periods. When the symmetric error-correction model is considered, changes in the exchange rate variable were significant only during the total sub-sample period. Foreign cost variable and the error-correction term are significant only during the second sub-sample period in both symmetric and asymmetric models.

7. Conclusion

This paper investigates the exchange rate pass-through to yen based manufactured import prices of Japan using asymmetric unit root and cointegration tests and asymmetric models. Due to sticky prices, for example, there are reasons to believe that the degree of pass-through depends on whether the exchange rate appreciates or depreciates. The sample used in this study covers the period January 1975 to June 1997. Since Japan experienced a recession since 1989, the data series was divided into two sub-sample periods: 1975:01 to 1989:12 and 1990:01 to 1997:06 to find evidence of significantly different exchange rate pass-through relationships over the pre- and post-recession periods. We believe that this relationship was affected during the recession, particularly when the contractual currency based exchange rate index is used in the empirical analysis

Using two state regime switching models, the estimated pass-through coefficients corresponding to appreciation and depreciation of the currency are found to be 98 percent and 83 percent respectively for the Johansen method for the total sample period; these coefficients are shown to be significantly different, particularly in the post recession period. The corresponding exchange rate pass-through coefficients for the EG method are 70 percent and 72 percent respectively which are not significantly different from each other. Moreover, we have shown that the recession in Japan in the 1990s has adversely affected the exchange rate pass-through relationship; both EG and Johansen methods indicate a significant asymmetric log-run response of import prices to changes in exchange rates. Forcing appreciations and depreciations to have the same effects on the import prices, the model does not appear to uncover the true underlying relationship. The analysis carried out in this paper suggests that one can learn much about the underlying adjustment processes in unit root testing, and the possibility of different responses of import prices to exchange rate changes.

APPENDIX A

Table 1. Results of the tests for asymmetric unit roots

Sample Period	Variable	Model	Levels		First Differences	
			Test for Unit roots	Test for asymmetry	Test for Unit roots	Test for asymmetry
1975:01 to 1997:06	*er*	TAR	3.454 (T)	0.517	40.804 (μ)**	0.574
		MTAR	3.240 (T)	0.941	23.845 (μ)**	0.834
	pm	TAR	2.877 (T)	0.485	43.687 (μ)**	0.321
		MTAR	2.664 (T)	0.792	43.011 (μ)**	0.000
	*c**	TAR	2.115 (T)	0.400	5.060 (T)	0.885
		MTAR	1.403 (T)	0.833	5.473 (T)	0.368
	c	TAR	2.905 (T)	0.604	15.986 (T) **	0.947
		MTAR	3.557 (T)	0.215	16.049 (T) **	0.733
1975:01 to 1989:12	*er*	TAR	2.278 (T)	0.615	13.995(μ)**	0.625
		MTAR	2.185 (T)	0.789	14.292(μ)**	0.387
	pm	TAR	1.461 (T)	0.235	26.994(μ)**	0.246
		MTAR	1.641 (T)	0.533	26.656(μ)**	0.324
	*c**	TAR	1.154 (T)	0.716	3.366 (T)	0.834
		MTAR	0.880 (T)	0.957	3.362 (T)	0.847
	c	TAR	1.060 (T)	0.753	15.460 T)**	0.422
		MTAR	1.237 (T)	0.503	14.977 T)**	0.915
1990:01 to 1997:06	*er*	TAR	2.278 (T)	0.615	13.995 (μ)**	0.625
		MTAR	2.185 (T)	0.789	14.292 (μ)**	0.387
	pm	TAR	1.460 (T)	0.235	26.994 (μ)**	0.246
		MTAR	1.641 (T)	0.532	26.656 (μ)**	0.325
	*c**	TAR	1.154 (T)	0.716	3.366 (T)	0.834
		MTAR	0.880 (T)	0.957	3.362 (T)	0.847
	c	TAR	1.060 (T)	0.753	15.460 (T)**	0.422
		MTAR	1.237 (T)	0.503	14.977 (T)**	0.915

Notes: *er, pm, c*, and c* are contractual currency exchange rate index, import price index in Japanese yen, foreign cost index and the domestic cost index. Columns four and five of the table report the results of the test for unit roots using equation (5) in the text augmented with lags of the dependent variable to produce whitenoise residuals. The number of lags of the dependent variable was determined by AIC and Ljung-Box-Q statistic. The null hypothesis for the unit roots using equation (5) is $\rho_1=\rho_2=0$ whereas that for asymmetry is $\rho_1=\rho_2$. 'μ' and 'T' within brackets with a unit root test statistic indicate whether the heaviside indicator function is a constant or a constant and a time trend respectively. Since the unit root test statistic in this case has a non-standard distribution, critical values generated by Enders and Granger which are given below were used;

Ender and Granger's Critical Values (1998, Table 1, p.306)

Model	Sample size	Estimated Constant Attractor			Estimated Trend Attractor		
		Probability of a smaller value			Probability of a smaller value		
		90%	95%	99%	90%	95%	99%
TAR	100	3.79	4.64	6.57	5.27	6.30	8.58
	250	3.74	4.56	6.47	5.18	6.12	8.23
MTAR	100	4.11	5.02	7.10	5.74	6.83	9.21
	250	4.05	4.95	6.99	5.64	6.65	8.85

To make inferences regarding the hypothesis for asymmetry standard F-statistic are used (Enders and Granger, 1998, p.307).

Table 2: Augmented Dickey-Fuller test results for Unit roots

Sample Period	Variable	Levels		First Differences	
		With Intercept	With intercept & Trend	With Intercept	With intercept & Trend
1975:01 to 1997:06	*er*	-1.065 (1)	-2.153 (2)	-6.901 (2)**	-6.897 (2)**
	er^D	0.189 (2)	-1.722 (2)	-10.463 (1)**	-10.458 (1)**
	pm	-1.144 (1)	-2.298 (1)	-6.766 (2)**	-6.787 (2)**
	c	-2.770 (3)	-2.429 (3)	-4.035 (2)**	-4.288 (2)**
	*c**	-2.867 (5)	-1.637 (5)	-3.633 (4)**	-4.371 (1)**
1975:01 to 1989:12	*er*	-0.837 (1)	-1.853 (1)	-5.264 (2)**	-5.248 (2)**
	er^D	-0.082 (1)	-1.414 (1)	-5.376 (2) **	-5.362 (2)**
	pm	-1.494 (1)	-1.693 (1)	-4.892 (2)**	-4.937 (2)**
	c	-2.265 (3)	-1.849 (3)	-3.186 (2)*	-3.440 (2)*
	*c**	-2.642 (1)	-0.532 (1)	-3.275 (2)*	-3.776 (2)*
1990:01 to 1997:06	*er*	-1.535 (1)	-1.569 (1)	-5.139 (1)**	-5.190 (1)**
	er^D	1.794 (3)	-0.671 (3)	-4.989 (2)**	-5.524 (2)**
	pm	-1.714 (1)	-1.994 (1)	-5.669 (1)**	-5.711 (1)**
	c	-1.221 (4)	-2.856 (4)	-4.992 (1)**	-5.017 (1)**
	*c**	-1.153 (4)	-3.0328 (3)	-5.047 (3)**	-5.077 (3)**

Note: See Table 1 for the definition of the variables except for er^D which is the exchange rate variable for depreciation episodes. '**' ('*') indicates significance at one (five) percent level. Figures within brackets associated with ADF t-statistics in columns 3 to 6 indicate the number of lags of the dependent variable in the ADF regression to eliminate the serial correlation from the residuals.

Table 3: Cointegration test results using asymmetric unit roots

Sample Period	Model			
	TAR		MTAR	
	Test for unit roots	Test for asymmetry	Test for unit roots	Test for asymmetry
1975:01 to 1997:06	8.002 (μ)***	0.578	7.223 (μ)***	0.634
1975:01 to 1989:12	4.141 (μ)*	0.587	4.833 (μ)*	0.205
1990:01 to 1997:06	5.187 (μ)**	0.680	5.697 (μ)**	0.301

Note: '***','**' and '*' denote significance at one, five and ten percent level respectively. The lags of TAR and MTAR models were selected using AIC. Variables included in cointegration regression are *pm*, *er, c*, and *c**. See note for Table 1 and 2 for the definitions of the above variables and critical values.

Table 4: Estimation results for the symmetric long-run exchange rate pass-through equation, Dependent variable: Yen based import price index

Variable	Engle-Granger Method			Johansen Method		
	1975:01 to 1997:06	1975:01 to 1989:12	1990:01 to 1997:06	1975:01 to 1997:06	1975:01 to 1989:12	1990:01 to 1997:06
Constant	0.153	-0.320	5.447**	-	-	-
Trend	-0.002**	-0.001**	-0.003**	-0.002**	-0.005**	-0.002
er	0.713**	0.677**	0.769**	0.958**	1.016**	0.633**
*c**	0.274**	0.571**	-1.016**	1.641**	-2.144**	0.811
c	0.773**	0.555**	0.969**	-0.990**	2.529**	-0.308

** significance at 1% level. In the Johansen cointegration equation, the dependent variable that is the yen based import price index was normalised to one.

Table 5: Estimation results for the asymmetric long-run exchange rate pass-through equation, Dependent variable: yen based import price index

Variable	Engle-Granger Method			Johansen Method		
	1975:01 to 1997:06	1975:01 to 1989:12	1990:01 to 1997:06	1975:01 to 1997:06	1975:01 to 1989:12	1990:01 to 1997:06
Constant	0.164	-0.394	1.501	-	-	-
Trend	-0.002**	-0.002**	0.001	-0.002	0.004**	0.002
er	0.704**	0.620**	1.143	0.977**	1.574**	1.144**
er^D	0.023	0.137	-0.667	-0.152	-0.334	-0.965**
c	0.230**	0.685**	-0.841	-1.019**	-2.590**	-0.344
*c**	0.760**	0.422**	1.552	-1.730**	2.995**	2.130**

Notes: ** significance at 1% level. In the Johansen cointegration equation, the dependent variable that is the yen based import price index was normalised to one.

Table 6: Results of the estimation of error correction models, Dependent variable: first differences of the yen based import price index

Variable	Sample period					
	1975:01 to 1997:06		1975:01 to 1989:12		1990:01 to 1997:06	
	Symmetric	Asymmetric	Symmetric	Asymmetric	Symmetric	Asymmetric
ε_{t-1}	-0.001	-0.002	0.038	0.036	-0.241**	-0.248**
Δpm (-1)	0.751***	0.745***	0.769***	-0.760***	0.659	0.689**
Δer (-1)	-0.262*	-	-0.284	-	-0.272	-
Δper (-1)	-	-0.287*	-	-0.325	-	-0.243
Δner (-1)	-	-0.242	-	-0.257	-	-0.349
Δc (-1)	-0.080	-0.087	-0.223	-0.226	-0.545	-0.491
Δc* (-1)	-0.075	-0.035	0.092	0.152	-1.025**	-1.127**

Notes: '***' , '**' and '*' indicate significance at one, five and ten percent levels respectively. The figures within brackets associated with a variable in column one denote the first lag of the respective variable. ε_{t-1} denotes the error-correction term which is the one period lagged residual from the symmetric cointegrating equation. Δ denotes the first difference of a variable. See note for Table 1 for the definitions of the variables.

Table 7: Test results for long-run asymmetry

Sample Period	Chi-square test statistic	
	Engle-Granger Method	Johansen Method
1975:01 to 1997:06	0.553	4.156*
1975:01 to 1989:12	1.980	2.681*
1990:01 to 1997:06	15.128**	2.844*

Notes: '**, and ' *' indicate significance at one and five percent levels respectively. Chi-square test is relevant for testing null hypothesis that the coefficient of the asymmetry variable (er^D) is significantly different from zero.

APPENDIX B

Figure 1. Time series behaviour of the log of the variables used in the exchange rate pass-through equation.

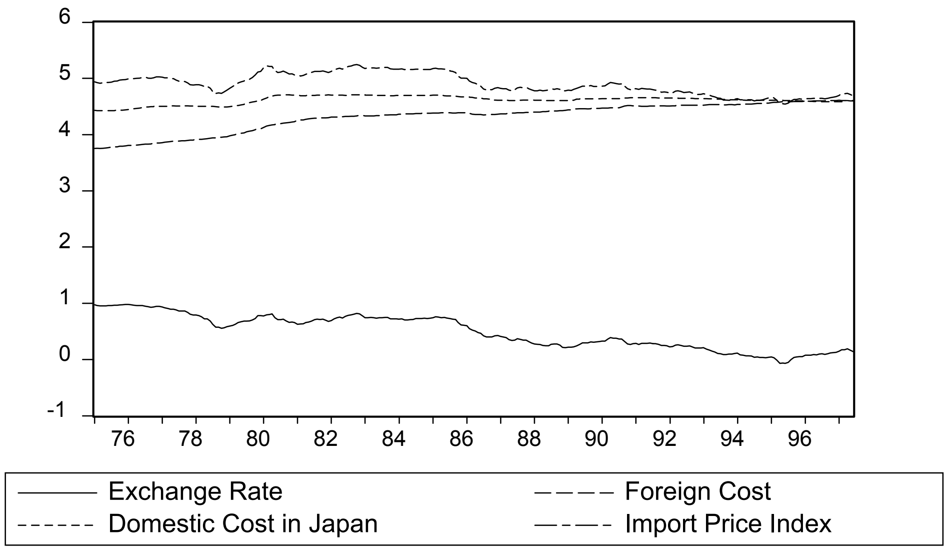
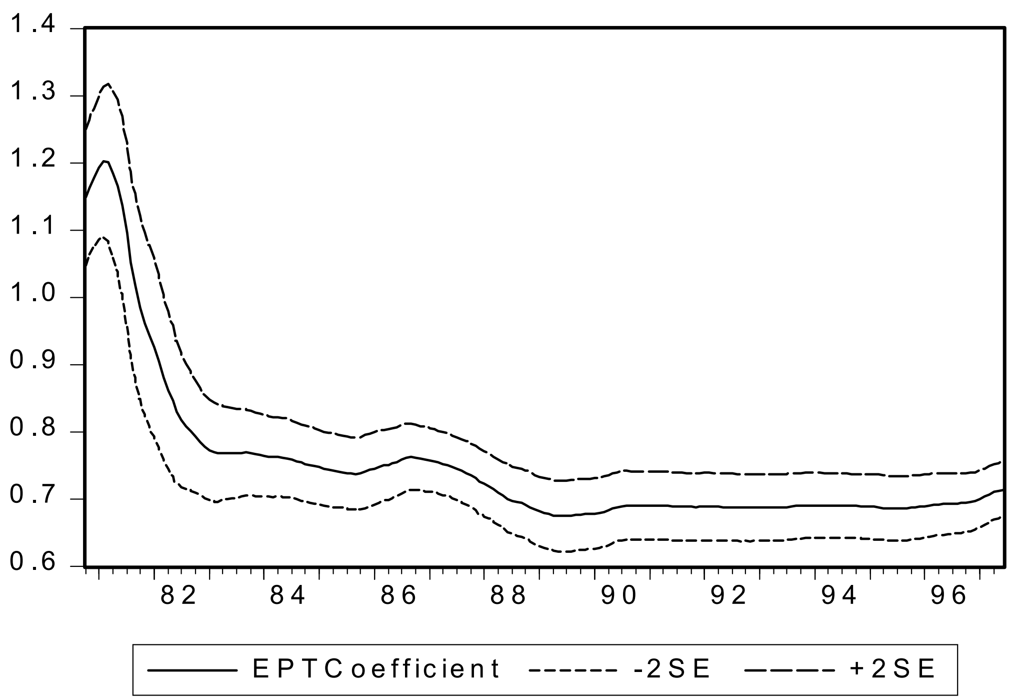


Figure 2

Recursive Estimate of Exchange Rate Pass-through Coefficient (er) and its Two Standard Error (SE) Bands for Total Manufactured Imports from October 1980 to June 1997 using an initial sample of 70 observations.



APPENDIX C

Japan's manufactured Import shares

Country	Weight
USA	0.4207
Korea	0.1138
Germany	0.1063
UK	0.0481
Thailand	0.0472
Singapore	0.0441
France	0.0427
Indonesia	0.0333
Switzerland	0.0271
Sweden	0.0198
Canada	0.0192
Philipines	0.0166
Australia	0.0146
Ireland	0.0141
India	0.0123
Netherlands	0.0113
Spain	0.0091

Data Sources

Exchange rates for all the countries were obtained from International Financial Statistics (IFS) CD-ROM-2000 of International Monetary Fund. Producer Price Index for manufacturing for Canada, Japan, Korea, France, Germany, Ireland, Switzerland, the USA and the UK were obtained from OECD main Economic Indicators and those for the other countries were obtained from the IFS CD-ROM, 2000. Price indices for manufactured imports and their respective weights were obtained from the Bank of Japan website. Value of manufactured imports according to SITC classification used to calculate import share were obtained from Foreign Trade by Commodities for 1992-97 published by OECD.

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