Exchange Rate Pass-through into Canadian Import Prices

by

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1. Introduction

Since the collapse of the Bretton Woods system in 1973, many countries have adopted a flexible exchange rate regime to remedy the balance of payments crises they encounter, which, during the last two decades, has motivated a number of empirical studies to step back and examine more closely the underlying relationship between exchange rates and the prices of internationally traded goods. This is well-known as "exchange rate pass-through."[1]

Exchange rate pass-through is the phenomenon whereby changes in the value of the foreign exchange rate are reflected in the domestic currency prices of traded goods, i.e., import and export prices (Menon 1996a, 434). This has been a hot issue because exchange rate pass-through can impact domestic market structure in ways that influence optimal monetary policy. Previous empirical studies in this area have mainly focused on the industrialized countries and can be broadly divided into three categories.[2] The first category focuses on modelling and estimating exchange rate pass-through into the import prices of specific industries, e.g., Kardasz and Stollery (2001); Parsley (2001). The second category examines exchange rate pass-through into aggregate import prices, e.g., Hooper and Mann (1989); Campa and Goldberg (2002); Wickremasinghe and Silvapulle (2001). The last category examines exchange rate pass-through into the Consumer Price Index (CPI) or the Wholesale Price Index (WPI), e.g., Papell (1994); McCarthy (2000). The growing volume of research on exchange rate pass-through at the industry-specific and aggregate level is partly motivated by the rise in industrial organization and strategic trade theory, while those studies of the exchange rate pass-through into the CPI or the WPI have accompanied the development of open economy macroeconomic models.

1. See for example, Spitaeller (1980); Kent and Dwyer (1994); Parsley (1995); Menon (1996b); Kenny and McGettigan (1998).
2. The detailed literature review will be given in the next section.

This paper has two main objectives. The first is to examine the degree of pass-through into aggregate import prices in Canada covering the period from 1980:1 to 2002:3. As Canada is often considered a small open economy, I compare my results with those of Menon (1996b), who studied the case of Australia, another typical small open economy, to determine whether his conclusion of incomplete pass-though for a small open economy holds or not in Canadian data. In this paper, I adopt the same theoretical model as Menon (1995b), which is derived from the Law of One Price.[3] This model has also been widely used in other empirical studies of exchange rate pass-through, for example, Hooper and Mann (1989), Kent and Dwyer (1994), Kenny and McGettigan (1996), and Campa and Goldberg (2002). However, in this paper, I also argue that there are some drawbacks associated with this approach.

The second objective of this paper is to compare the results for Canada with those for other industrialized countries, especially the U.S., to check whether the degree of pass-through depends on the country's size and the degree of openness.

This paper applies an econometric procedure which overcomes the pit-falls in previous studies that may lead to an underestimation of the degree of pass-through. After I applying the GLS-detrended Augmented-Dickey Fuller test of Elliott, Rothenberg and Stock (1996) to check the stationarity of the variables, due to the presence of structural break, I provide estimates of pass-through based on Bai and Perron's (2003) multiple structural change model of cointegration vectors, and also correct the small sample bias of the estimates. As a result, the pass-through estimate for Canada indicates a relatively faster import price adjustment in the long-run than is typically

3. The Law of One Price (LOP) states that all transactions in a competitive market occur at a single price if the market is in equilibrium. Under LOP, the nominal exchange rate should adjust to the point where the good costs the same amount when purchased in either country, as long as the price is measured in a common currency.

found in studies of other industrialized countries. The results in this study are found to be robust to a number of sensitivity tests.

The remainder of the paper is organized as follows. In section 2, I give a brief overview of previous studies of this issue. Section 3 provides a theoretical model of exchange rate pass-through. Section 4 describes the data used in the estimation. Econometric approaches are presented in section 5. Estimation results and sensitivity tests are reported in section 6. Finally, section 7 provides some concluding remarks.

2. Brief Literature Review

Most of the previous international analyses do recognize that changes in the exchange rate may affect not only import prices but also the prices of goods produced domestically, because imported goods are often used as inputs in domestic firms. But they have different views on the degree of pass-through. Some of the studies have concluded that incomplete pass-through is a common and pervasive phenomenon across a broad range of countries.[4] Several studies, however, found that incomplete pass-through is very common in the short-run, but it does not carry through to the long-run. Accordingly, complete exchange rate pass-through is a long-run phenomenon, especially in a small open economy.

From a theoretical perspective, incomplete pass-through is mainly the outcome of three factors: product differentiation and market structure, macroeconomic conditions, and individual price-setting behaviour. First, in a perfectly competitive market where imported and domestically produced goods are perfect substitutes, firms will charge a price equal to marginal cost. Under the conditions of imperfect competition, however, pricing will no longer be at marginal cost, and

4. Menon (1995a) summarized the results of the previous studies and found that only 6 out of the 46 studies report complete or close to complete pass-through. These are Spitaeller (1980), Garnaut and Baxter (1984), Helkie and Hooper (1988), Citrin (1989), Lawrence (1990) and Leith (1990).

firms would be in a position to charge a mark-up over their marginal cost to earn above normal profits even in the long-run.[5] This mark-up may vary in response to an exchange rate change depending on the degree of substitutability between domestically produced goods and imported goods as determined by the degree of product differentiation and the degree of market integration or separation. The lower the degree of substitutability between those goods and the lower the degree of market integration, the greater will be the market power of sellers and the greater will be the effects on prices in response to exchange rate changes.[6]

Second, according to the sticky price theory, foreign suppliers will face a menu cost in altering their usual supply patterns, so that they will only adjust their prices in a discrete interval. Prices are, for example, likely to be lowered only gradually in response to an exchange rate appreciation, as it takes time for firms to expand their supply capacity in response to price reductions. Thus, it is obvious that pass-through will be incomplete in the short-run.

Finally, at the level of individual industries, price-setting strategies may also affect the degree of pass-through. For example, if the agents have the power to set prices and seek to maximize profit, exchange rate pass-through is likely to be high regardless of other factors (Phillips 1988). Alternatively, if agents seek to maximize market share rather than profit, pass-through may be incomplete, as they will lower their mark-ups to deter the entry of potential competitors (Hooper and Mann 1989; Ohno 1990).

During the past two decades, a number of empirical studies have used various methodologies and data series for different countries to test these theoretical propositions on the degree of

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5. In an imperfectly competitive market, one pricing behaviour is called mark-up pricing. Mark-up pricing involves setting the market price equal to a mark-up over marginal cost, where the amount of the mark-up depends on the elasticity of demand.
6. Kent and Dwyer (1994) indicate that in an imperfectly competitive market, foreign suppliers may act to offset the effects of depreciation by lowering their margins so that exchange rate pass-through is incomplete.

exchange rate pass-through. In the remainder of this section, I will discuss some of these studies and their results. First, Menon (1995b) estimated the exchange rate pass-through relationship for Australian imports of manufactures covering the period 1981:3 to 1992:2. He applied an econometric procedure which avoids the pit-falls of previous studies to a carefully assembled data set,[7] and for the first time in the literature employed Johansen's Maximum Likelihood estimator of cointegrating vectors.[8] He found that exchange rate pass-through to import prices for a small open economy was incomplete, i.e., 66.27%, even in the long-run. Finally, he also put forward some implications of incomplete pass-through for policy and the macroeconomy. For example, incomplete pass-through implied that exchange rate policy may be a blunt instrument when used to restore external balance since relative price adjustments would be limited.[9] Kent and Dwyer (1994), on the other hand, tested and estimated the cointegrating relationship between the exchange rate, domestic prices and the world prices, using the Phillips & Hansen (1990) fully modified OLS estimator. They found that in the long-run exchange rate pass-through was complete for both Australian imports and manufactured exports.[10] However, in the short-run, pass-through would be incomplete. They also found that pass-through to import prices was more rapid than that to manufactured export prices. Finally, in order to confirm the robustness of their results, the Bewley (1979) transformation was applied to an unrestricted error correction model.[11]

7. He indicated that previous studies suffer from a number of short-comings. First, most researchers ignored the time series properties of the data, i.e., non-stationarity and time trends. Second, they use the unit value index as a proxy for the actual price index which is inadequate data. So, in his paper, he used actual import prices and overcame the problems associated with the 'world price index' by employing an import-weighted foreign cost of production index.
8. In Menon (1993), Engle-Granger's two step cointegration test was carried out. In Menon (1995b), however, he argued that this approach would be invalid in the case of more than one cointegrating vector and that it is unable to accommodate dynamics in the cointegrating regression.
9. See also Menon (1995a; 1996b).
10. In the absence of other shocks, the relative price elasticities of demand and supply are the principal determinants of exchange rate pass-through. For imports, the degree of pass-through will increase the lower is the elasticity of demand and the greater is the elasticity of supply. From this it follows that pass-through will be complete in the case of a small open economy.

Yang (1995) studied exchange rate pass-through in U.S. manufacturing industries and its cross-sectional variation, by applying the Dixit-Stigliz (1977) model of product differentiation. Using import price data for three- and four-digit SIC industries, the paper found that pass-through was incomplete and varied across industries. In the cross-sectional study, the degree of pass-through was found to be positively correlated with different proxies for product differentiation, and negatively correlated with a proxy for the elasticity of marginal cost with respect to output.

Campa and Goldberg (2002) provided cross-country and time series evidence for the imports of twenty-five OECD countries on the issue of whether producer-currency-pricing (PCP) or local-currency-pricing (LCP) of imports was more prevalent. Under the model associated with the Law of One Price and using import unit value data, they showed that, for the OECD as a whole, incomplete pass-through was the best description of import prices in the short-run. But in the long-run, although complete pass-through was still rejected for many countries, the pass-through elasticity was close to one.[12] Their study further revealed that pass-through into import prices was lower for countries with low average inflation and low exchange rate variability. The Chow test and Hansen's (1997) test showed that the pattern of pass-through was unstable, i.e., declining over time in some countries, which was contradictory to the results of Athokorala and Menon (1994), who found that the pass-through relationship had remained stable through the sample period. However, this conclusion was the same as that of Wickremasinghe and Silvapulle (2001), who concluded that exchange rate depreciation and appreciation had systematic asymmetric effects on Japanese manufactured import prices using Enders and Granger's (1998) threshold autoregressive (TAR) model and the momentum threshold autoregressive (MTAR) model asymmetric unit root

11.See Kent and Dwyer (1994), pp.411.

12. As an unweighted cross-country average, average pass-through into import prices is 0.61 in the short-run, and 0.77 in the long-run.

test,[13] Johansen's cointegration test, and an asymmetric error-correction model. The paper also indicated that incomplete pass-through in the short-run for the whole sample period may due to sticky prices.

Sahminan (2002) used Perron's (1989) unit root test[14] and Johansen's cointegration test to estimate exchange rate pass-through into import prices in some Southeast Asian countries, i.e., Thailand, Singapore, and the Philippines and found that in the short-run, fluctuations in the exchange rate did not have a significant effect on import prices in Thailand and Singapore, but had a significant negative effect on the aggregate price level in the Philippines. Inversely, Parsley (2001) estimated a panel of very disaggregated import unit-values for Hong Kong, doing the basic pooled regressions under the constraint that the pass-through coefficients were equal across Hong Kong's 21 major exporters. He confirmed that the changes in the exchange rate had a relatively greater effect on import prices than for other East Asian countries, given that Hong Kong had more price flexibility.

Kenny and McGettigan (1996) assessed the extent to which exchange rate changes affected Irish import prices using the same model and approaches as Menon (1995b). More notably, they concentrated on the case of a small open economy, and theirs was the first study in the literature to uncover two close to full pass-through long-run equilibrium relationships among the data, i.e., one between import unit values(pm_{unit}), the exchange rate(er) and foreign costs(cp) and another between domestic competing prices(pd) and the same two variables.[15] Finally, they put forward

13. As the alternative hypothesis of the traditional ADF test is a symmetric adjustment of the variables of interest, there is a possibility that the ADF test rejects 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. TAR and MTAR are two major alternative models. See Enders and Granger (1998) for details.
14. Sahminan (2002) pointed out that in the presence of structural breaks, the traditional ADF test was biased.

three possible explanations for the incomplete short-run pass-through: a very slow speed of adjustment to the long-run relationships; the "menu costs" associated with altering prices; and the existence of non-tariff barriers and the entry of firms associated with the exchange rate movements.

Taking studies on exchange rate pass-through in the literature and econometric methodologies developed recently together, previous empirical studies did not incorporate some well-known facts into modelling and estimating. Firstly, many used inappropriate data. Import unit values are not a suitable series to be used as price proxies, as the discrepancy between them and actual prices is large enough to result in biased estimators (Alterman 1991). Secondly, many studies ignored the time series properties of the data, i.e., structural breaks, the number of cointegrating relationships, non-stationarity, small sample bias, etc. Correspondingly, incorrect econometric approaches were used.

Therefore, my study attempts to (1) check time series characteristics in my datasets employing more reasonable approaches in testing for unit roots and cointegration; (2) estimate the degree of exchange rate pass-through both in the short-run and in the long-run; (3) check whether the results are sensitive to the measure of foreign cost; and finally (4) examine the relationship between the degree of pass-through and the country size.

3. Model of Exchange Rate Pass-through

In this section, an overview of the simple theoretical relationship between the exchange rate and import prices is discussed. As mentioned earlier, exchange rate pass-through is broadly

15. Kenny and McGettigan (1996) pointed out that many of the previous results demonstrating substantially less than full pass-through may be due to a failure to make proper allowance for the strong simultaneity which exists between import and domestic competing prices. So, their estimation models were as follows: $pm_{unit} = \alpha_1 er + \alpha_1 cp$; $pd = \alpha_2 er + \alpha_2 cp$.

defined as the percentage change in importer currency dominated import prices as a result of a one percent change in the exchange rate between the importing and exporting countries. A one-to-one response of import prices to the exchange rate is widely known as "full" or "complete" exchange rate pass-through, while a less than one-to-one response is, therefore, called "partial" or "incomplete" exchange rate pass-through.

In this paper, the model used to test the degree of the exchange rate pass-through is derived from a simple mark-up model,[16] which is similar to the model used in many of the studies in this area. This model is popular because traded goods are typically differentiated and traded in markets that are characterized by imperfect competition. Under imperfect competition, firms would charge a mark-up over their marginal cost in order to earn above normal profits in the long-run. Thus, it is assumed that foreign producers set their export prices (PX) in terms of their foreign currency as a mark-up (π) over their cost of production (CP) in their own currency as well:

$$PX = \pi CP . \tag{3.1}$$

As the mark-up π is equal to $1+\lambda$, where λ is the profit margin, the above equation can be rewritten as:

$$PX = (1+\lambda)CP . \tag{3.2}$$

According to the Law of One Price in the absolute sense, the importer's domestic currency (i.e. Canadian dollar) import prices (PM) are, therefore, given by:

$$PM = PX \times ER = (1+\lambda)CP \times ER = \pi CP \times ER , \tag{3.3}$$

16. It is argued that a simple mark-up approach is more appropriate for the types of goods involved in international trade where imperfect competition prevails.

where ER is the exchange rate defined as the domestic currency price of one unit of foreign currency, i.e., Canadian dollars per U.S. dollar. It is assumed that the mark-up, π, depends on competitive pressures in the domestic market, and the exchange rate. This competitive pressure is proxied by the gap between the prices of domestic import competing goods (PD) and the exporters' production cost (CP) (Hooper and Mann, 1989). Accordingly, the profit mark-up can be modelled as:

$$\pi = \{(PD)/(CP \times ER)\}^{\alpha}, \qquad (3.4)$$

where $\alpha \in [0,1]$. Substituting equation (3.4) into equation (3.3) yields:

$$PM = \{(PD)/(CP \times ER)\}^{\alpha}\{CP \times ER\} . \qquad (3.5)$$

Making some simple rearrangements and taking natural logarithms of the variables (denoting them in lower case letters), the equation for the import prices could be reformulated as follows:

$$pm = \alpha pd + (1-\alpha)cp + (1-\alpha)er . \qquad (3.6)$$

The pass-through coefficient $(1-\alpha)$ is expected to be between 0 and 1. In the extreme case, when the foreign firm is a price taker in the Canadian competitive market, i.e., $\alpha=1$, then the corresponding coefficients for foreign cost and the exchange rate will be zero, which means that the pass-through will be zero.[17] In this case, the import price set by the foreign firm will be equal to the domestic price, and therefore the exchange rate has no impact on import prices. On the other hand, when $\alpha=0$, the coefficients on foreign cost and the exchange rate will be one, which implies that the exchange rate pass-through is complete, because in this case the foreign firm does not face

17. As Menon (1993) pointed out, most studies make use of a representative "world price" rather than a cost variable. But the world price represents the pricing decision of foreign producers to all markets and not just to the home market. The use of a foreign cost variable, however, does not suffer from such drawbacks, as this cost does not depend on the particular export market being targeted.

any competition in the Canadian market. Hence, incomplete pass-through is characterized by α between 0 and 1.

However, I argue that there exist two drawbacks of this model. One is that it imposes the restriction that the pass-through coefficients of the foreign cost and the exchange rate into import prices are the same, which I don't think is necessary for the empirical analysis, because in the short-run, the exchange rate is more volatile than costs. A reasonable conjecture is that exporters will be more willing to absorb into their mark-up changes in exchange rates than changes in costs, which are likely to be permanent (Bache 2002). Moreover, apart from purely economic reasons, the coefficient restrictions may not hold due to incompatibility of the price proxies, which may result from differences in aggregation levels and methods of data collection. Therefore, I do not impose such a restriction on my estimating equation.

The second argument is that the Law of One Price may not hold in the absolute sense, given that all the price series are weighted indices and not single-good prices. Therefore, a non-zero constant term is added to the regression equation so as to permit the importers to choose a steady state value for pm that differs from that predicted by equation (3.6). The economic interpretation of the constant term is that the Law of One Price holds in the relative sense in the case of α=0.

Hence, after taking the above two points into account, equation (3.6) is modified as follows:[18]

$$pm_t = \beta_0 + \beta_1 \times er_t + \beta_2 \times cp_t + (1-\beta_1) \times pd_t + u_t\,, \qquad (3.7)$$

where the residual u_t is assumed to be white-noise.

Therefore, equation (3.7) is interpreted as a long-run cointegrating relationship. The import price determination in the short-run will be explained by a dynamic model, i.e., an Error-Correc-

18. The mark-up in this case is defined as follows: $\pi = \{(PD)/CP^{\beta} \times ER\}^{\alpha}$.

tion representation, where changes in import prices depend on the deviation from the long-run relationship and on current and lagged changes in explanatory variables.

4. Data

In this section, I will discuss how the variables used in the model are constructed. The variables constructed for this paper are the domestic import price index (pm), the exchange rate index (er), the foreign cost index (cp), and the domestic production cost index (pd).

In estimating the model, I use quarterly data covering the period from 1980:1-2002:3. The proxies for the variables used in the model are as follows. The domestic import price index (pm) is measured by the index of average import prices in Canadian dollars, which is calculated as the ratio of nominal import value (mn) and real import value (m).[19]

For the exchange rate (er), I specially construct an index of nominal import-weighted effective exchange rates using the following formula:

$$er = \sum_i w_i \times pfx_i \tag{4.1}$$

where w_i is the assigned weight, and pfx_i is the corresponding nominal bilateral exchange rate, i.e., Canadian dollar price of foreign currency. The import shares are based on bilateral trades from foreign countries to Canada during 1980-2002. The Appendix shows the weights in detail.

I use the foreign GDP deflator as the proxy for foreign cost (cp),[20] because the GDP deflator is a broad measure of the general price level of the economy, given that the import price is a general aggregate over all sectors. The foreign cost index is constructed by taking the weighted aver-

19. As the real import series in CANSIM II are constructed such that 1997=1.00, all the time series used in equation (3.7) are rebased to 1997=1.00.
20. Unit Labour Cost (ULC) should be the most accurate proxy for measuring foreign cost. The problem is that, however, some countries of the Euroland area do not published such data.

age of the top four exporters to Canada, i.e., the U.K., the U.S., Japan and Euroland, using the following formula:

$$cp = \sum_i w_i \times gdpd_i , \tag{4.2}$$

where w_i is the same import weight used in the calculation of er; and $gdpd_i$ is the corresponding GDP deflator for those countries. Accordingly, the proxy for domestic production cost (pd) is the GDP deflator for Canada.

Detailed information about the data sources is provided in the Appendix. All the variables used in the estimation are in logarithmic form. Figure 1 shows the time series behaviour of each individual variable. As shown, all the data have an obvious upward time trend and non-zero mean. Apparent structural breaks are found in both the import price and the exchange rate.

5. Econometric Methodology

Because there is a possibility of long-run equilibrium among the variables used in this paper, I intend to use a Vector Error Correction Model (VECM) to estimate their relationship. Application of the VECM requires that each series in the model be non-stationary but they have a linear combination that is stationary. Therefore, at first, unit root tests are carried out. Meanwhile, according to Figure 1, there might be a structural break for both the exchange rate and import prices; hence, a multiple structural change model is provided to determine the break point. Then, Johansen's cointegration test is applied to both the entire sample and several segments. Finally, a VECM is constructed based on Johansen's cointegration test. In the remainder of this section, I describe briefly the econometric procedures employed for the unit root test, the structural break test, the cointegration test, and the VECM estimation.

5.1 Unit Root Test

The poor power of the standard ADF test us associated with the fact that the true autoregressive coefficient may be very close to unity. Elliott, Rothenberg and Stock (1996) constructed an efficient univariate unit root test based on local-to-unity asymptotic theory, denoted as ADF^{GLS}, given that all the series have an obviously upward linear deterministic trend. Running the ADF test using GLS-detrended data leads to a better overall performance in terms of small sample size and power than the standard ADF test.[21]

Suppose the series y_t is given by:

$$y_t = d_t + u_t \quad , \tag{5.1}$$

$$d_t = \beta' x_t \quad , \tag{5.2}$$

where x_t contains optional exogenous regressors which may consist of constant, or a constant and trend. In the following explanation, suppose $x_t = (1, t)'$, i.e., y_t has a linear deterministic trend. The GLS-detrended series y_t is given by $y_t^d = y_t - \tilde{\beta}' x_t$. where $\tilde{\beta}$ is the least squares estimate of the regression of $\tilde{y}_t$ on $\tilde{x}_t$, i.e., $\tilde{\beta} = (\sum \tilde{x}_t \tilde{x}_t')^{-1} \sum \tilde{x}_t \tilde{y}_t$, where $\tilde{y}_t$ and $\tilde{x}_t$ are the quasi-differences of y_t and x_t, respectively, i.e., $\tilde{y}_t = (y_1, (y_2 - a y_1), \ldots, (y_T - a y_{T-1}))'$, and $\tilde{x}_t = (x_1, (x_2 - a x_1), \ldots, (x - a x_{T-1}))'$. $a = 1 + \bar{c}/T$ represents the local alternative.[22]

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21. ADF^{GLS} does not allow for the presence of structural breaks. Perron and Rodríguez (2002), however, modified this test with unknown break points. They considered two specific models: one with a break in the slope of the trend function and one with a break in both the intercept and slope. Nevertheless, their approach is not implemented in any econometric software, and it is really very difficult for me to apply it empirically at this stage.
22. The selection of the non-centrality parameter $\bar{c}$ is based on a feasible optimal point test, denoted by P_T^{GLS} in ERS (1996). Specifically, $\bar{c} = -7$ corresponds to the tangency between the asymptotic local power function of the test and the power envelope at 50% power in the case with a constant, in which $x_t = 1$. For the case with constant and trend, $\bar{c} = -13.5$.

Then, estimate the regression below:

$$\Delta y_t^d = \alpha y_{t-1}^d + \beta_1 \Delta y_{t-1}^d + \ldots + \beta_p \Delta y_{t-p}^d + v_t \ . \qquad (5.3)$$

Since the y_t^d are detrended, x_t is not included in the test equation (5.3).

The issue of the lag selection in ADF^GLS regressions has received much attention recently. In subsequent applications, I employ a new lag selection procedure proposed by Ng and Perron (2001), the Modified Akaike Information Criterion (MAIC), that provides the best combination of size and power in finite samples.[23] Under this criterion, the lag length k is defined as follows:

$$k_{maic} = \arg min_{k \in [0, k_{max}]} \left\{ \log\left(s_v^2\right) + \frac{2(\hat{\Gamma}_T(k) + k)}{T^*} \right\} , \qquad (5.4)$$

where $\hat{\Gamma}_T(k) = (s_v^2)^{-1} \hat{\alpha}^2 \sum_{t = k_{max}+1}^{T} (y_{t-1}^d)^2$, $\hat{\alpha}$ is obtained from equation (5.3) and $s_v^2 = (T-k)^{-1} \sum_{t=k+1}^{T} \hat{v}_t^2$; $T^* = T - k_{max}$. Note that once the lag length k is selected, say at k^*, equation (5.3) will be re-estimated from $t = k^* + 1$ to T.

As with the standard ADF test, we consider the t-ratio for $\hat{\alpha}$ from equation (5.3). While the ADF^GLS t-ratio follows an ADF distribution in the constant case only, the asymptotic distribution differs when including both a constant and trend. ERS (1996, Table 1, 825) simulated the critical values of the test statistic in the latter case for $T = \{50, 100, 200, \infty\}$. The null hypothesis of a unit root ($\alpha = 0$) will be rejected if the t-statistic falls below those critical values.

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23. The advanage of the MAIC over other criteria is that it takes into account the estimate of the autoregressive parameter $\hat{\alpha}$, which allows a better measure of the cost of each lag-length choice.

5.2 Cointegration Test and Error-Correction Model

5.2.1 Multiple Structural Change Model

Normally, if all the variables in the estimation equation are I(1), a cointegration test is applied to examine the possibility of a long-run relationship. However, according to Figure 1, there appears to be a structural break in the data, though the exact break point is unknown. Therefore, in order to test the cointegrating relationship pre- and post-break, Bai's and Perron's (2003) multiple structural change model is used first to test for the break point, under the assumption that a long-run cointegrating relationship holds among the variables.[24]

It is true that the Bai and Perron's tests are valid only for stationary data. However, they are applied anyway by some authors either to help pinpoint the date of a structural break or under the assumption that the data are stationary around a trend, as in Burdekin and Siklos (1999) and Ben-David and Papell (2000), etc. Therefore, at this point, the tests may be valid for my case, but the appropriate critical values are as yet unknown. So, in my following empirical studies, Bai and Perron's tests are still applied.

Suppose a multiple linear-regression system with m breaks ($m+1$ sub-samples) is expressed as follows:

$$y_t = x'_t\beta + z'_t\delta_j + u_t, \quad t = T_{j-1} + 1, \ldots, T_j \tag{5.5}$$

for $j = 1, \ldots, m+1$; where $x_t(p \times 1)$ and $z_t(q \times 1)$ are vectors of covariates, and β and δ_j are the corresponding vectors of coefficients.[25] Note that only the parameter vector δ differs across sub-sam-

24. The subsequent empirical results support my prior that there is a long-run cointegrating vector among the variables.

25. This is a partial structural change model because the coefficient vector β is not allowed to shift and is estimated using the entire data set.

ples. Here the break points, $(T_1, ..., T_m)$, are treated as unknown.[26] Equation (5.5) can also be written in matrix form:

$$Y = X\beta + \bar{Z}\delta + U, \tag{5.6}$$

where $Y = (y_1, ..., y_T)'$, $X = (x_1, ..., x_T)'$, $U = (u_1, ..., u_T)'$, $\delta = (\delta_1', ..., \delta_{m+1}')'$, $\bar{Z} = diag(Z_1, ..., Z_{m+1})$ with $Z_i = (z_{T_{i-1}+1}, ..., z_{T_i})'$.

Bai and Perron (2003) constructed a sup F type test with the null hypothesis of no structural break ($m = 0$) versus the alternative hypothesis of k breaks ($m = k$). Let the structural break point be $T_i = T\lambda_i$, $(i = 1, ..., k)$, where the λ_i are the break fractions. Let R be the matrix such that $(R\delta)' = (\delta_1' - \delta_2', ..., \delta_k' - \delta_{k+1}')$. Define

$$F_T(\lambda_1, ..., \lambda_k; q) = \frac{1}{T}\left(\frac{T-(k+1)q-p}{kq}\right)\hat{\delta}'R'(R\hat{V}(\hat{\delta})R')^{-1}R\hat{\delta}, \tag{5.7}$$

where $\hat{V}(\hat{\delta})$ is the estimate of the variance covariance matrix of $\hat{\delta}$, i.e., of

$$V(\hat{\delta}) = \sigma^2(T^{-1}\bar{Z}'M_X\bar{Z})^{-1}, \tag{5.8}$$

which can be estimated using $\hat{\sigma}^2 = T^{-1}\sum_{t=1}^{T}\hat{u}_t^2$. In this paper, $\hat{V}(\hat{\delta})$ is constructed under the assumption of serially uncorrelated errors and the same asymptotic distribution for those estimates obtained by global minimization.[27] The test statistic is $\text{Sup}F_T(k;q) = F_T(\hat{\lambda}_1, ..., \hat{\lambda}_k, q)$, where $\hat{\lambda}_1, ..., \hat{\lambda}_k$ minimize the sum of squared residuals of equation (5.6). Bai and Perron (2003) pointed out that

26. T_j, $j = 0, 1, ..., m+1$, is the number of observations up to j break point. For example, $T_0 = 0$ and $T_{m+1} = T$, where T is the number of total observations.
27. Bai and Perron (2003) considered three alternatives that differ in terms of the assumptions made about autocorrelation and heteroskedasticity.

this is asymptotically equivalent to maximizing the F-test of equation (5.7), because the estimated dates of the break points will be consistent even if there is autocorrelation between the error terms across segments.[28] Bai and Perron (2003) also present two additional tests of the null hypothesis that there are no breaks against the alternative of an unknown number m of breaks. Further details regarding these two tests are provided in their paper.

Once the number of break points has been determined, the dates of the breaks can be treated as endogenous, and estimated simultaneously with the corresponding coefficients of the segments by OLS. Let $\hat{\beta}$ and $\hat{\delta}$ be the estimate of β and δ, respectively. The method of estimation considered is that based on the least-squares principle. For each break point, the associated least-squares estimates of $\hat{\beta}$ and $\hat{\delta}$ are obtained by minimizing the sum of the squared residuals, which is defined as follows:

$$(Y - X\hat{\beta} - \bar{Z}\hat{\delta})'(Y - X\hat{\beta} - \bar{Z}\hat{\delta}) = \sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} (y_t - x'_t\hat{\beta} - z'_t\hat{\delta}_i)^2 . \qquad (5.9)$$

So, in this case, the regression parameter estimates are the estimates associated with breaks, i.e., $\hat{\beta} = \hat{\beta}(\{T_j\})$, $\hat{\delta} = \hat{\delta}(\{T_j\})$. Then the break points can be chosen to be those that minimize the sum of squared errors.

5.2.2 Johansen's cointegration Test

Given that each variable is I(1), it is possible that they are cointegrated, that is, they have a linear combination that is stationary. In this paper, Johansen's cointegration test, which is described as follows, is employed to test for the presence of cointegration.

28. There is no restriction that the distribution of the regressors z_t be the same across segments, because it may imply an inaccurate approximation especially if a small segment length is allowed, in which case the exact moment matrix of the regressors may deviate substantially from its full sample analog.

Consider a VAR representation:

$$Y_t = \Pi_0 + \Pi_1 Y_{t-1} + \ldots + \Pi_p Y_{t-p} + \varepsilon_t \, , \qquad (5.10)$$

where $Y_t = (y_{1t}, y_{2t}, \ldots, y_{kt})'$, $\Pi_0(k \times 1)$ is a vector of intercepts, $\Pi_i(k \times k)$ is a coefficient matrix, and ε_t is a $k \times 1$ vector that is identically independently normally distributed for all t with mean zero and variance Ω.

Equation (5.9) can be rewritten in terms of first differences to obtain

$$\Delta Y_t = \Pi_0 + \Pi Y_{t-p} + \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} + \varepsilon_t \, , \qquad (5.11)$$

where $\Pi = -(I - \Pi_1 - \ldots - \Pi_p)$, and $\Gamma_i = -(I - \Pi_1 - \ldots - \Pi_i)$. The optimal lag length is chosen by the sequential modified likelihood ratio criterion.

Now, let r be the rank of the matrix Π . If $r = 0$, then equation (5.11) is just a VAR representation of Y_t in the variables in differences. If $r = k$, then Y_t is a stationary process, that is, there is no cointegrating relationship among the variables in Y_t . If $0 < r < k$, then there must exist a matrix of cointegrating vectors, β, and a matrix of adjustment parameters, a, such that $\Pi = a\beta'$ and $\beta' Y_t$ is stationary.

The number of significant cointegrating vectors is determined by examining the trace statistic (Q_{trace}) or maximum characteristic root statistic (λ_{max}) given by:

$$Q_{trace} = -T \sum_{i = r-1}^{k} \ln(1 - \hat{\lambda}_i) \, , \qquad (5.12)$$

or

$$\lambda_{max} = -T\ln(1-\hat{\lambda}_{r+1}), \tag{5.13}$$

where $\hat{\lambda}$ is the estimated value of characteristic roots. The difference between these two tests is that for the maximum characteristic root test the null hypothesis of r cointegrating vectors is tested against the alternative of $r+1$, while for the trace test the alternative hypothesis is that the number of cointegrating vectors is k, where k is the number of endogenous variables. Moreover, concerning the sample size, a small sample correction has been recommended for Johansen's $\lambda_{\max}$ test by Reimers (1992). This modification is shown as follows.

$$\lambda'_{max} = \frac{(T-kp)}{T}\lambda_{max}\ , \tag{5.14}$$

where T is the number of observations used to estimate the model, k is the number of variables included in the VAR system, and p is the number of lags in the VAR system.

The critical values for Q_{trace} and $\lambda_{\max}$ are provided by Osterwald-Lenum (1992). All the test statistics should be greater than the critical value in order to reject the null hypothesis.[29]

5.2.3 Small Sample Bias Adjustment

Given that equation (3.7) is a single equation of the VECM, if all the variables in the estimating equation are cointegrated and weakly exogenous, then equation (3.7) could be estimated using ordinary least squares (OLS) for both the entire sample and pre- and post-break segments. However, another problem would arise, i.e, small sample bias, as the sample period used in my estimation is

29. The traditional Johansen's cointegration test does not allow for structural break. Maddala and Kim (1998) discussed the effects of structural break on cointegration tests (in chapter 13 of their book). There are some residual-based tests for cointegration that allow for structural break. But personally, I do not know how to extend Johansen's test to allow for structural break. That issue could be a good direction for my future research.

from 1980:1 to 2002:3, a total of 91 observations, which is relatively small and may cause a downward bias in the estimates of the cointegrated system. Thus, in this paper, I use an OLS estimation method suggested by Stock and Watson (1993) that consists of adding contemporaneous values, leads and lags of the first difference of the integrated regressors on the right hand side of the regression equation.[30] This will reduce small sample bias because these new added regressors capture part of the correlation between the dependent variable and the independent variables.[31]

In my case, I add the contemporaneous values and one lag of the first difference of all independent variables only, in order not to reduce the degrees of freedom too much. Therefore, equation (3.7) is modified as follows to estimate the long-run relationship:

$$pm_t = \beta_0 + \beta_1 \times er_t + \beta_2 \times cp_t + (1-\beta_1) \times pd_t + \sum_{i=0}^{1} \delta_i \Delta er_{t-i} + \sum_{i=0}^{1} \varphi_i \Delta cp_{t-i} + \sum_{i=0}^{1} \theta_i \Delta pd_{t-i} + u_t \quad (5.15)$$

Then, the short-run relationship is estimated using the following ECM:

$$\Delta pm_t = c + \varphi \hat{u}_{t-1} + \sum_{i=1}^{k-1} w_i \Delta pm_{t-i} + \sum_{i=1}^{k-1} \theta_i \Delta er_{t-i} + \sum_{i=1}^{k-1} \gamma_i \Delta cp_{t-i} + \sum_{i=1}^{k-1} \beta_i \Delta pd_i + \upsilon_t , \quad (5.16)$$

where c is a constant term which is related to the behavior of the deterministic components of the time series; $\hat{u}_{t-1}$ is a one period lagged error term from the cointegrating equation (5.15); υ_t is a white-noise error term; and φ is known as the speed of adjustment.

I choose the optimal lag value to be 2, given 91 total observations.

30. Stock, James H. and Watson, Mark W. (1993), "A Simple MLS of Cointegrating Vectors in Higher Order Integrated Systems," *Econometrica*, July 1993, quoted in Caballero (1994).

31. Caballero (1994) indicated that an operational problem of this procedure, however, is that for variables with strong dynamic behaviour, such as the stock of capital, running out of degrees of freedom is a serious consideration.

6. Empirical Analysis and the Results

6.1 Analysis of the Results

First, I apply the ADF^{GLS} unit root test to check the stationarity of all the variables used in equation (3.7). Tables 1 and 2 report the test results for those variables in levels and first differences respectively. The optimal lags used in the test are chosen to minimize the Modified Akaike Information Criterion (MAIC). As shown, I can not reject the null hypothesis of a unit root for all variables in levels at the 1% level of significance while I can reject the hypothesis that those variables have unit roots in first differences at any conventional level of significance.[32] Furthermore, as a sensitivity check, I also report the traditional ADF unit root test results in the two tables. The same conclusion holds for most cases except that we do not reject a unit root for the first difference of the exchange rate at the 1% level of significance.

As discussed in the previous section, a visual inspection of a weak relationship between the nominal effective exchange rate and the import price in the early sample and a much stronger correlation between the two in the later sample also suggests the likelihood of the presence of structural breaks among these variables. In order to identify the number and the time of the break points, I apply a statistical test originally proposed by Bai and Perron (2003) to equation (3.7) under the assumption that there is a long-run cointegrating relationship among the variables. Table 3 presents the results of the structural break test with the null hypothesis of no structural break point ($k = 0$) over the whole sample period against the alternative of one structural break ($k = 1$). In addition to the SupF test, I also report the results of two alternative tests, i.e., D_{max} and WD_{max}.[33] As shown, all tests reject the null hypothesis of no structural break point at the 1% level

32. We reject the null hypothesis of unit root for series *pm* at both 5% and 10% level of significance.
33. See Bai and Perron (2003) for details. Note that these three tests would be statistically identical when $k = 1$.

of significance and conclude that there is one break point. The high value if the test statistic suggests that the test result is more likely to be correct, even though the critical values are not appropriate for I(1) data. Furthermore, the SupF test also estimates the presence of a break point at 1988:2, which is consistent with our prior as indicated in Figure 2. This break was mainly due to the Free Trade Agreement between Canada and the U.S., which became effective from August 1988. Under this agreement, the economic structure of Canada thus changed, since the U.S. became the largest trade partner for Canada. Those structural changes include increased aggregate employment, lower tariffs, higher prices of most commodities, etc. Moreover, since the U.S. bears 86.51% of the weight in calculating both er and cp, it is not surprising that has a large impact on the estimation results.

Next, I apply Johansen's cointegration test to the entire sample as well as two sub-samples that are separated by the identified structural break point from the SupF test to examine the presence of the long-run relationship among variables. Those test results are presented in Table 4.[34] In terms of the choice of optimal lag length for the test, which is also listed in Table 4, I use the sequential modified likelihood ratio criterion. As shown, both the trace statistic and the modified Max-Eigen Statistic reject the null hypothesis of no cointegrating vector at the 5% level of significance except for one case when the Max-Eigen statistic is used for the first sub-sample. However, it rejects the null hypothesis at the 10% significance level. This result suggests that the evidence of a long-run relationship among the variables is stronger in the second sub-sample than in the first one. Also, I can see that both test statistics almost uniformly do not reject the hypothesis of one cointegrating vector at the 5% significance level, which implies that there exists one long-run relationship among those variables.

34. Reasonably, Max-Eigen statistic is more reliable than trace statistic, because the small sample correction is employed.

The presence of the cointegrating relationship allows us to apply OLS to equation (5.15) to estimate the degree of exchange rate pass-through in the long-run after the small sample adjustment. Table 5 reports the estimation results for the whole sample period as well as two sub-samples. The coefficient associated with the nominal effective exchange rate (er in the table) measures the degree of pass-through. Almost all the estimates of exchange rate pass-through are statistically significant at the 1% level of significance regardless of the sample range. The estimated average degree of pass-through for the entire sample is around 0.75 while this estimate differs substantially across the two sub-samples, with about 0.33 in the first sub-sample and 0.78 in the second one. These results suggest that the pass-through of the exchange rate to domestic import prices is much stronger in the second sub-sample than in the first one, which is an indication of the breakdown of this long-run relationship due to the presence of structural breaks in certain variables. Over the entire sample period, the estimate of 0.75, although not statistically significantly different from one, still reveals a close to full pass-through long-run equilibrium relationship, which is consistent with the hypothesis underlying a small open economy. In terms of the estimates of the coefficient associated with the foreign cost measure (cp in the table), they exhibit the expected signs and are statistically significant at the 1% significance level in both the entire sample and the second sub-sample, while they do not appear to have the right signs for the first sub-sample. The estimates of the coefficients of the prices of domestic competing goods (pd in the table) are all statistically significant at the 1% significance level and have the correct signs while the magnitude of those estimates differ significantly across the two sub-samples. The larger estimates associated with pd in the first sub-sample relative to those in the second sub-sample indicate that over the first sub-sample period, the impact of domestic competing goods prices on import prices dominates the exchange rate effect. Furthermore, I test the null hypothesis that the degree of pass-through for the exchange rate (er) and foreign cost (cp) is equal. Table 6 shows that we strongly reject the null hypothesis, which is consistent with the argument I raised in section 3 that the exchange rate is more volatile than costs.

To examine the short-run dynamics of imported prices, I estimate an error-correction model as presented in equation (5.16). The estimation results are reported in Table 7. As shown, changes in the exchange rate in one lagged period are significant determinants of import prices at the 10% significance level during the second sub-sample only, while insignificant in the whole sample period and the first sub-sample. Moreover, the estimated speed of adjustment for the entire sample is around -0.111, while this estimate differs substantially across the two sub-samples, with -0.029 and -0.334 for the first and second sub-sample respectively. The null htpothesis that the speed of adjustment is zero can be rejected only for the second sub-sample at the 10% level of significance. These results indicate that when import prices exceed their long-run equilibrium by 1%, they would be adjusted downwards at a rate of 0.11%, 0.03% and 0.33% a quarter for the corresponding sample range. These slow adjustment speeds maybe caused by the "menu costs" of altering prices and also indicate incomplete exchange rate pass-through in the short-run.

6.2 Robustness Checking and Comparison with Other Studies

Given that numerous studies of exchange rate pass-through use different possible measures of foreign cost variables, I employ one alternative measure, i.e., the foreign producer price index (cp_{ppi}), to check the robustness of my previous results.[35] As shown in Table 8, both the trace statistic and the modified Max-Eigen statistic still reject the null hypothesis of no cointegrating vector at the 5% level of significance except for one case which is the same as discussed before. Also, for all cases, the hypothesis of one cointegrating vector is uniformly not rejected at the 5% signif-

35. Unit root has been checked. See Tables 1, 2 for details.

icance level. These results are fully consistent with those reported in Table 4, so my conclusion on the number of cointegrating vectors is not sensitive to different measures of foreign cost.

Table 9 reports estimation results for the long-run relationship equation after the small sample adjustment using equation (5.15).[36] As shown, the estimates of the degree of exchange rate pass-through are very similar to those reported in Table 5 even though an alternative measure of one explanatory variable is used, and they are statistically significant at the 1% level of significance. The degree of exchange rate pass-through for the first sub-sample is still much smaller than that for the second sub-sample as well as the whole sample. The estimate of the coefficient associated with the foreign cost variables is also statistically significant at the 1% level and exhibits correct signs in both the second sub-sample and the whole sample, while it still does not have the right sign for the first sub-sample. Again, my estimation results on the long-run relationship are quite robust to this alternative measure of one of the explanatory variables. Regarding the estimation of short-run dynamic equation, i.e., the error-correction model, the estimates of the speed of adjustment to the long-run relationship, i.e. the coefficient associated with u_{t-1} in Table 10, are still very slow, and insignificant at 10% level of significance for two sub-samples, implying that incomplete exchange rate pass-through holds in the short-run. To summarize, my results presented in this section are consistent with those reported earlier and very robust to an alternative measure of the foreign cost variable.

Comparing my results with those of previous studies. I argue that Menon's (1995b) conclusion might be incorrect in the sense of a small open economy, while I agree with Kenny and McGettigan (1996), and Kent and Dwyer (1994). Moreover, judging by the point estimates, the degree of exchange rate pass-through varies from country to country. In particular, Australia and the U.S.

36. The sample starts from 1981:1 due to the availability of PPI series.

have relatively lower incomplete pass-through in the long-run, i.e., 0.596 and 0.292, respectively, during 1975-1999 (Campa and Goldberg, 2002). This wide range of point estimates indicates that the degree of aggregate pass-through is affected by the changes in the country composition of trade, while country size and openness are irrelevant.

7. Conclusion

This paper investigates the exchange rate pass-through to domestic currency based aggregate import prices of Canada using a simplified mark-up model. The sample used in this study covers the period from 1980:1 to 2002:3. Due to the Free Trade Agreement between the U.S. and Canada, there is reason to believe that a structural break exists. The SupF test indicates that 1988:2 is a break point. Therefore, the entire data series is divided into two sub-periods: 1980:1 to 1988:2 and 1988:3 to 2002:3, to examine the exchange rate pass-through relationships. I believe that the relationship is significantly different over the two sub-samples, particularly because the U.S. takes such a large trade weight in Canadian total imports.

Adjusted for small sample bias, the estimated degree of exchange rate pass-through for the entire sample and two sub-samples are 74.71 percent, 33.31 percent and 78.42 percent respectively, which uncovers a close to full pass-through long-run relationship especially in the second sub-sample. The larger estimates associated with the price index of domestic competing goods and the smaller one associated with the exchange rate indicate that the impact of the exchange rate is dominated by that of domestic competing goods in the first sub-sample. Furthermore, the degrees of pass-through for the exchange rate and foreign costs are found to be statistically different, due to the volatility of the exchange rate, compared with cost variables. Moreover, on the issue of the short-run relationship, my results suggest that any deviation from long-run equilibrium will take some time to be restored for import prices due to the sticky price theory, and hence support the notion of incomplete pass-through in the short-run. Finally, my results are very robust to the alternative measure of one explanatory variable.

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Figure 1

Time series behavour of the log of the variables used in the exchange rate pass-through equation. (individual)

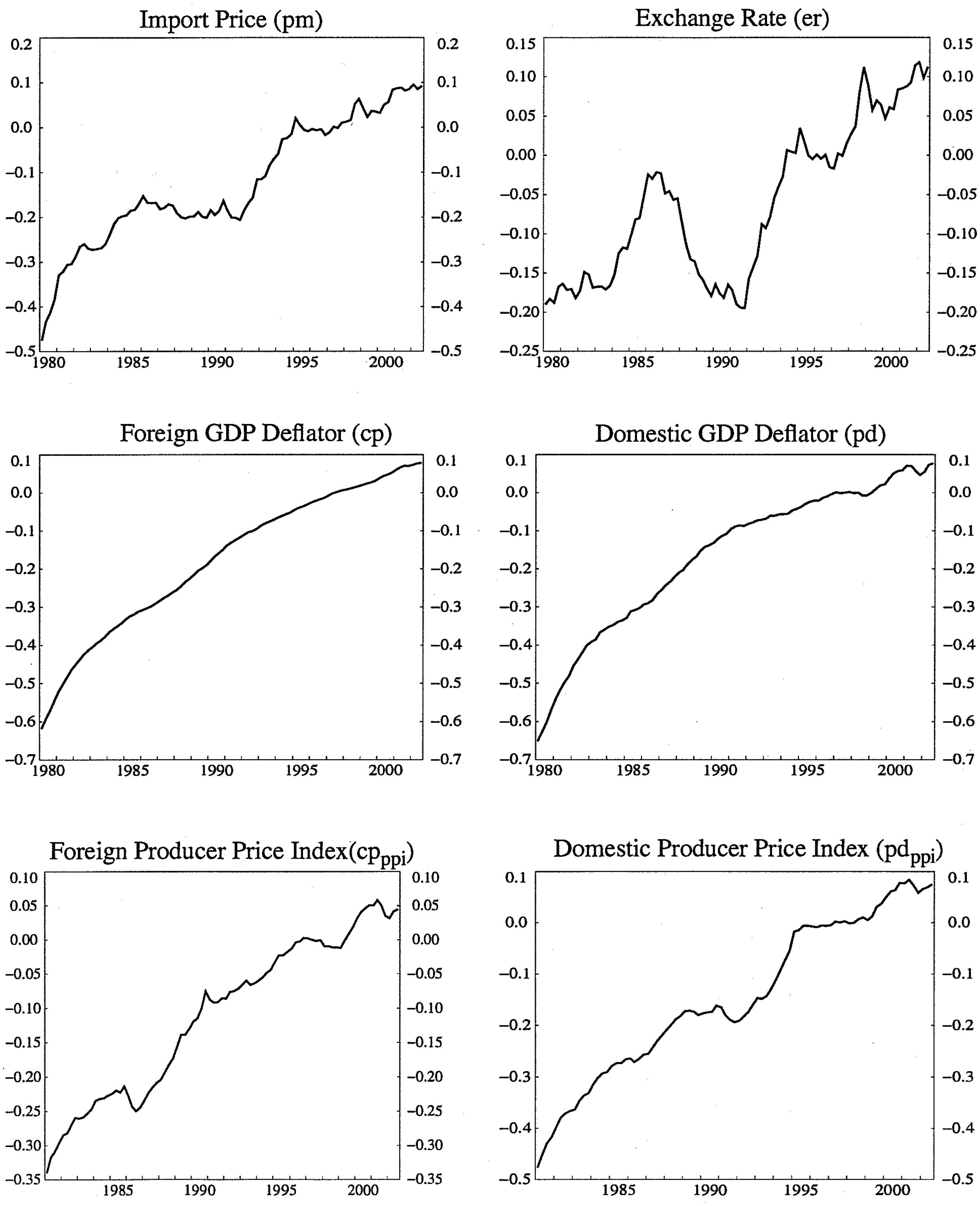

Figure 2

Relationship among all the variables.

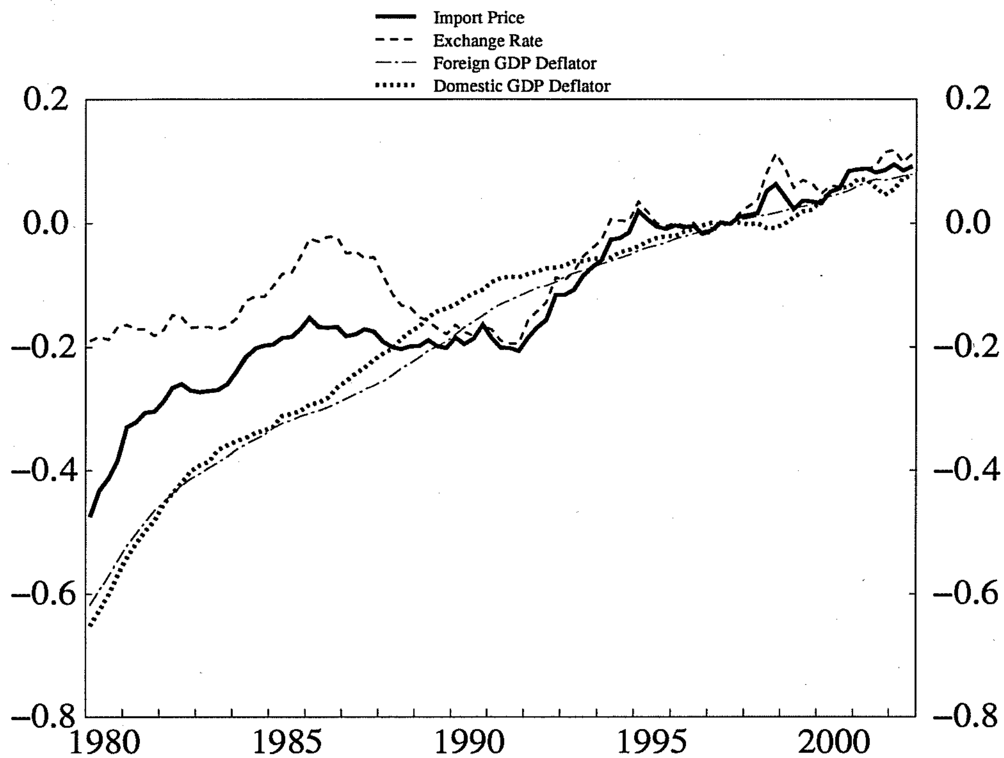

Figure 2 shows a visual inspection of a weak relationship among those variables in the early sample and a much stronger correlation in the later sample, which suggests that there might be a structural break among the variables.

Table 1: Results of Unit Root Tests for Level

Variables	ADF	ADF^{GLS}	Optimal Lag
cp	-0.0833	-1.0243	3
er	-2.3610	-1.7297	1
pd	-2.4686	-0.8728	3
pm	-2.4451	-1.5974	4
cp_{ppi}	-2.8418	-1.6273	1
pd_{ppi}	-2.5391	-1.4163	1

Note:

1. ADF/ADF^{GLS} = Augmented Dickey Fuller Test, GSL-Detrended Augmented Dickey Fuller Test.
2. A constant and linear time trend are included in all cases.
3. Optimal lag length is the one which minimizes the Modified Akaike Information Criterion (MAIC).
4. The last two are used in section 6.
5. Critical Value: Critical values for ADF^{GLS} are obtained from Elliott, Rothenberg and Stock (1996), Table 1, pp. 825.

Significance Level	ADF	ADF^{GLS}
1%	-4.0742	-3.77
5%	-3.4652	-3.19
10%	-3.1589	-2.89

Table 2: Results of Unit Root Tests for the First Difference

Variables	ADF	ADF^{GLS}	Optimal Lag
Δcp	-3.8359	-3.2927	2
Δer	-3.4918	-3.4350	2
Δpd	-7.7117	-4.0632	3
Δpm	-4.4913	-3.5656	5
Δcp_{ppi}	-3.7620	-3.5853	2
Δpd_{ppi}	-4.5457	-3.2473	2

Note:

1. ADF/ADF^{GLS} = Augmented Dickey Fuller Test, GSL Detrended Augmented Dickey Fuller Test.
2. A constant is included in all cases.
3. Optimal lag length is the one which minimizes the Modified Akaike Information Criterion (MAIC).
4. The last two are used in section 6.
5. Critical Value: Note that the critical values for ADF^{GLS} test are the same as those for the standard ADF test without constant and no trend.

	ADF	ADF^{GLS}
1%	-3.5132	-2.6062
5%	-2.8976	-1.9467
10%	-2.5858	-1.6131

Table 3: Results of Multiple Structural Break Test

Hypothesis	Tests	Test Statistics	Critical Values 1%	5%	10%
H_0: k=0	SupF	65.56	19.22	15.19	13.22
H_a: k=1	D_{max}	65.56	19.22	15.28	13.36
	WD_{max}	65.56	20.34	16.13	14.16

Table 4: Results of Johansen's Cointegration Test

Hypothesized No. of CE(s)	Trace Statistic 1980:1-2002:3	Trace Statistic 1980:1-1988:2	Trace Statistic 1988:3 2002:3	Max-Eigen Statistic 1980:1-2002:3	Max-Eigen Statistic 1980:1-1988:2	Max-Eigen Statistic 1988:3-2002:3
None	67.11**	58.02**	67.10**	30.11**	26.01*	29.28**
At most 1	25.61	29.44	25.61	17.02	13.18	16.67
At most 2	17.70	14.38	17.70	14.66	7.78	14.35
At most 3	2.26	5.49	2.26	2.15	4.80	2.11
Optimal Lag	1	1	2	1	1	2

Note:
1. (**), (*) denotes rejection of the hypothesis at the 5%, 10% levels, respectively.
2. Max-Eigen Statistic is modified by the small sample correction.
3. Optimal lag length is chosen by the sequential modified likelihood ratio criterion.
4. Critical value:

	Q_{trace} 5%	Q_{trace} 10%	λ_{max} 5%	λ_{max} 10%
None	47.21	43.84	27.07	17.15
At most 1	29.68	26.70	20.97	13.39
At most 2	15.41	13.31	14.07	10.60
At most 3	3.76	2.71	3.76	2.71

Table 5: Estimation Results for the Long-run Exchange Rate Pass-through Equation

Variables	1980:1 to 2002:3 after adjust	1980:1 to 1988:2 after adjust	1988:3 to 2002:3 after adjust
constant	**-0.0169*** **	**-0.0491** **	**-0.0183*** **
	(0.0000)	(0.0460)	(0.0024)
er	**0.7471*** **	**0.3331*** **	**0.7842*** **
	(0.0000)	(0.0020)	(0.0000)
cp	**0.1351*** **	**-0.2698** **	**0.1746*** **
	(0.0000)	(0.0154)	(0.0000)
pd	**0.2529*** **	**0.6669*** **	**0.2158*** **
	(0.0000)	(0.0002)	(0.0000)
Δ er	**0.1020**	**0.5229** **	**-0.0742**
	(0.5713)	(0.0384)	(0.4665)
Δ er(-1)	**1.2027**	**0.4750* **	**-0.0716**
	(0.6440)	(0.0671)	(0.4593)
Δ cp	**0.0703**	**-1.7111**	**2.1233** **
	(0.9191)	(0.3053)	(0.0332)
Δ cp(-1)	**0.0568**	**2.4554**	**-0.1033**
	(0.2151)	(0.1869)	(0.9168)
Δ pd	**-0.5954**	**-1.2740* **	**-0.5778* **
	(0.1395)	(0.0869)	(0.0951)
Δ pd(-1)	**0.0895**	**-0.2326**	**0.1276**
	(0.8245)	(0.7328)	(0.7039)
Observations	88	31	57
R-squared	0.9847	0.9614	0.9910
DW Statistic	0.2592	0.6280	0.3652
Likelihood	241.9183	90.1964	182.2902

Note:
1. The figures in parentheses are p-values.
2. (***), (**), and (*) indicate statistical significance at the 1, 5, and 10 percent levels, respectively.
3. All the statistically significant results in this table are robust to changes in the adjustment method, i.e., to adding lead variables to the adjustment process.

Table 6: Hypothesis Test Results

Hypothesis	1980:1 to 2002:3 Test Results	Conclusion	1980:1 to 1988:2 Test Results	Conclusion	1988:3 2002:3 Test Results	Conclusion
H_0: $\beta_1 = \beta_2$	$\chi^2_{(1)} = 910.20$	reject	$\chi^2_{(1)} = 26.59$	reject	$\chi^2_{(1)} = 164.54$	reject
	p_{value}=0.000		p_{value}=0.000		p_{value}=0.000	
H_a: $\beta_1 \neq \beta_2$						

Table 7: Estimation Results for Error-Correction Model

Variables	1980:1 to 2002:3	1980:1 to 1988:2	1988:3 to 2002:3
Constant	**0.0010**	**-0.0099**	**0.0089**
	(0.7760)	(0.1917)	(0.1252)
u_{t-1}	**-0.1108**	**-0.2548**	**-0.3340***
	(0.2391)	(0.1354)	(0.1007)
Δpm_{t-1}	**0.2395**	**-0.1854**	**0.8041**
	(0.2844)	(0.5689)	(0.0423)
Δpm_{t-2}	**0.0516**	**-0.0825**	**-0.0597**
	(0.8105)	(0.7936)	(0.8701)
Δer_{t-1}	**-0.0723**	**0.3592**	**-0.6551***
	(0.7109)	(0.1517)	(0.0690)
Δer_{t-2}	**-0.1043**	**-0.1668**	**0.0744**
	(0.5747)	(0.5309)	(0.8163)
Δcp_{t-1}	**-0.2336**	**2.2247**	**-2.5544****
	(0.7854)	(0.0618)	(0.0403)
Δcp_{t-2}	**1.0632**	**0.5021**	**1.9854***
	(0.1869)	(0.7194)	(0.1034)
Δpd_{t-1}	**-0.1886**	**-0.6708**	**-0.2494**
	(0.5570)	(0.2281)	(0.6018)
Δpd_{t-2}	**-0.1773**	**-0.1718**	**-0.1869**
	(0.6119)	(0.7303)	(0.7188)
Observations	88	31	57
R-squared	0.1201	0.5587	0.1702

Note:
1. The figures in parentheses are p-values.
2. (***), (**), and (*) indicate statistical significance at the 1, 5, and 10 percent levels, respectively.

Table 8: Results of Johansen's Cointegration Test for (pm, er, cp_{ppi}, pd_{ppi})

Hypothesized No. of CE(s)	Trace Statistic			Max-Eigen Statistic		
	1981:1-2002:3	1981:1-1988:2	1988:3-2002:3	1981:1-2002:3	1981:1-1988:2	1988:3-2002:3
None	58.93**	43.96**	62.97**	34.02**	20.29*	31.58**
At most 1	24.90	20.29	26.23	12.94	12.55	11.93
At most 2	11.96	5.65	12.36	10.55	4.05	9.21
At most 3	1.41	0.92	1.64	1.41	0.789	1.41
Optimal Lag	2	1	2	2	1	2

Note:
1. (**), (*) denotes rejection of the hypothesis at the 5%, 10% levels, respectively.
2. Max-Eigen Statistic is modified by the small sample correction.
3. Optimal lag length is chosen by the sequential modified likelihood ratio criterion.
4. Critical value:

	Q_{trace}		λ_{max}	
	5%	10%	5%	10%
None	47.21	43.84	27.07	17.15
At most 1	29.68	26.70	20.97	13.39
At most 2	15.41	13.31	14.07	10.60
At most 3	3.76	2.71	3.76	2.71

Table 9: Estimation Results for the Long-run Relationship Using PPI

Variables	1981:1 to 2002:3 after adjust	1980:1 to 1988:2 after adjust	1988:3 to 2002:3 after adjust
constant	**-0.0064*** **	**-0.0944*** **	**-0.0034***
	(0.0006)	(0.0000)	(0.0655)
er	**0.7017*** **	**0.2528*** **	**0.7373*** **
	(0.0000)	(0.0004)	(0.0000)
cp_{ppi}	**0.1779*** **	**-0.5642*** **	**0.1817*** **
	(0.0007)	(0.0000)	(0.0002)
pd_{ppi}	**0.2983*** **	**0.7472*** **	**0.2627*** **
	(0.0000)	(0.0000)	(0.0000)
Δ er	**0.3098*** **	**0.5777*** **	**0.2283** **
	(0.0002)	(0.0000)	(0.0048)
Δ er(-1)	**0.2151** **	**0.5638*** **	**0.1481***
	(0.0114)	(0.0004)	(0.0600)
$\Delta\ cp_{ppi}$	**0.0709**	**0.7075** **	**-0.1533**
	(0.7384)	(0.0112)	(0.4617)
$\Delta\ cp_{ppi}(-1)$	**-0.1216**	**0.5123** **	**-0.4332** **
	(0.5487)	(0.0456)	(0.0398)
$\Delta\ pd_{ppi}$	**0.0376**	**-1.0870*** **	**0.2259**
	(0.8429)	(0.0021)	(0.2117)
$\Delta\ pd_{ppi}(-1)$	**0.2288**	**-0.6780***	**0.2278**
	(0.2441)	(0.0515)	(0.2047)
Observations	85	28	57
R-squared	0.9939	0.9876	0.9951
Likelihood	276.5546	107.1240	199.6196

Note:

1. The figures in parentheses are p-values.
2. (***), (**), and (*) indicate statistical significance at the 1, 5, and 10 percent levels, respectively.
3. All the statistically significant results in this table are robust to changes in the adjustment method, i.e., to adding lead variables to the adjustment process.

Table 10: Estimation Results for Error-Correction Model

Variables	1981:1 to 2002:3	1980:1 to 1988:2	1988:3 to 2002:3
Constant	**0.0047****	**0.0087**	**0.0060****
	(0.0302)	(0.1170)	(0.0358)
u_{t-1}	**-0.2940***	**-0.1179**	**-0.0772**
	(0.0673)	(0.7903)	(0.7772)
Δpm_{t-1}	**0.2356**	**0.9533**	**0.4624**
	(0.4744)	(0.1129)	(0.3625)
Δpm_{t-2}	**-0.0353**	**0.5151**	**-0.2330**
	(0.9172)	(0.3129)	(0.6360)
Δer_{t-1}	**-0.0666**	**-0.2782**	**-0.3245**
	(0.7911)	(0.4539)	(0.4078)
Δer_{t-2}	**-0.0506**	**-0.5340**	**0.1333**
	(0.8501)	(0.1744)	(0.7396)
Δcp_{t-1}	**0.0104**	**0.2839**	**-0.3164**
	(0.9735)	(0.5441)	(0.4729)
Δcp_{t-2}	**-0.2201**	**-0.2730**	**-0.2309**
	(0.4501)	(0.5477)	(0.5853)
Δpd_{t-1}	**-0.0491**	**-0.4313**	**-0.0320**
	(0.8513)	(0.4511)	(0.9271)
Δpd_{t-2}	**0.1031**	**-0.5557**	**0.2286**
	(0.6918)	(0.3671	(0.4802)
Observations	84	27	57
R-squared	0.0856	0.3748	0.0638

Note:
1. The figures in parentheses are p-values.
2. (***), (**), and (*) indicate statistical significance at the 1, 5, and 10 percent levels, respectively.
3. All the statistically significant results in this table are robust to changes in the adjustment method, i.e., to adding lead variables to the adjustment process.

Appendix

The export weights used in this paper are calculated by choosing the top four Canadian trading partners, which covers above 90% of the annual average of Canadian total imports (from 1995 to 2002). Source: Direction of Trade Statistics Yearbook, International Monetary Fund, various years.

Canadian Import Shares

Country	Weight
United States	0.8651
Euroland Area*	0.0599
Japan	0.0531
United Kingdom	0.0219

* Euroland Area, in this case, includes 11 countries, i.e., Germany, France, Italy, Belgium, Luxembourg, Austria, Finland, Ireland, Netherlands, Portugal and Spain.

Data Source: The sample period for all the series is from 1980:1 to 2002:3.

	Description	Source	Mnemonics
mn	Import of goods and services (nominal)	Statistics Canada, CANSIM II	v498106- 380-0002
m	Import of goods and services (real)	Statistics Canada, CANSIM II	v1992063- 380-0002
pfx_i	Nominal bilateral exchange rate	Bank of Canada Database	U.S.: iexe0101; Euro: ecucae01 (before 1998:4), eurocae01 (after 1998:4); Japan: iexe0701; U.K.: iexe1201.
$gdpd_i$	Foreign GDP deflator	Bank of Canada Database	U.S.: us_p; Euro: euro_p; Japan: jap_p; U.K.: gbr_p.

	Description	Source	Mnemonics
pd	GDP deflator for Canada	Statistics Canada, CANSIM II	Nominal GDP: v498086- 380-0002; Real GDP: v1992067- 380-0002.
cp_{ppi}	Producer price index for foreign countries	OECD, Bank of Canada Database	U.S.: q.usa.ppiamp01.ixob; Euro: euro_ppi (defore 1989:4), q.emu.ppiamp01.ixob (after 1989:4); Japan: q.jpn.ppiamp01.ixob; U.K.: q.gbr.ppiamp01.ixob.
pd_{ppi}	Producer price index for Canada	OECD	q.can.ppiamp01.ixob