A JOINT TEST OF PRICING-TO-MARKET, MENU COST AND CURRENCY INVOICING

Jean-Philippe Gervais

Canada Research Chair in Agri-industries and International Trade
CRÉA and Department of Agricultural
Economics and Consumer Studies, Laval University,
Quebec, QC, G1K 7P4. Email: Jean-Philippe.Gervais@eac.ulaval.ca.

Bruno Larue

Canada Research Chair in International Agri-Food Trade
CRÉA and Department of Agricultural
Economics and Consumer Studies, Laval University.

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Abstract: This paper investigates PTM behaviour and currency invoicing decisions of Canadian pork exporters in the presence of menu costs. It is shown that when export prices are negotiated in the exporter's currency, menu costs cause threshold effects in the sense that there are bounds within (outside of) which PTM is not (is) observed. Conversely, PTM is not interrupted by menu costs when export prices are denominated in the importer's currency. The empirical model focuses on pork meat exports from Canada to the U.S. and Japan. Hansen's (2000) threshold estimation procedure is used to jointly test for currency invoicing and PTM in the presence of menu costs. Inference is conducted using bootstrap methods. PTM effects are smaller when accounting for currency invoicing decisions and menu costs than under standard linear models. The data does not reject the null hypothesis that Quebec pork exporters exercise PTM behaviour in the Japanese market and invoice their sales in Japanese currency. Evidence of PTM behaviour and foreign currency invoicing is weak for the U.S. market. Ontario pork exporters do not exercise PTM behaviour in any market.

Keywords: Pricing to market, Currency invoicing, Threshold estimation, Pork exports.

J.E.L. Classification: F12, F14, C22

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1. Introduction

Krugman (1986) defined the concept of pricing to market (PTM) in international trade as the case in which exporting firms charge different prices in foreign and domestic markets. There exists a considerable literature documenting evidence of PTM in international markets. Knetter (1989) was the first to document PTM effects when analyzing the pricing strategies of German and American firms. Since then, the literature has followed two different paths. One strand of the literature has focused on the macroeconomic implications of PTM. For example, Bergin and Feenstra (2001) blame PTM behaviour for the high degree of volatility in exchange rates.

The other, and more popular, strand of the literature concentrates on PTM's microeconomic implications. At the manufacturing level, Uctum (2003) and Sasaki (2002) analyzed the PTM behaviour of Japanese exporting firms. Gil-Pareja (2002) found that the degree of mark-up adjustment in response to exchange rate changes is similar across export markets. Similarly, agri-food sectors are known to exhibit PTM effects. For example, Carew and Florkowski (2003) found price discriminatory behaviour by Canadian and U.S. exporters of agri-food products. Brown (2001) found PTM effects in pricing of Canadian canola exports. Other studies include Griffith and Mullen's (2001) analysis of Australia's rice exports and Pick and Carter's (1994) wheat study.

On the other hand, the literature related to currency invoicing is a little thinner. Bowen, Hollander and Viaene (1998) surveyed the literature on currency invoicing and trade. Most studies deal with the apparent stylized fact that international transactions are invoiced in the exporters' currency. Others have analyzed the role of currency invoicing as an exchange rate risk hedging strategy (Donnenfeld and Haug, 2003; Johnson and Pick, 1997).

The concept of pricing-to-market is intrinsically linked to currency invoicing; yet few authors have formally tied the two concepts. One notable exception is Sato (2003). He uses an empirical model that distinguishes short-run and long-run pricing strategies of Japanese exporters. Exporters can stabilize their export prices by adjusting their profit margin and invoicing in the importer's currency. Standard estimation techniques usually capture the long-run pure PTM effect of export pricing decisions (explained by the curvature of the importers' demand function), but cointegration techniques are required for the estimation of short-run effects as well (stabilization effect of currency invoicing).

Larue, Gervais and Rancourt (2003) found a case for which currency invoicing has important implications when measuring PTM effects. They investigated PTM behaviour in the presence of menu costs. Menu costs are incurred by firms whenever they make changes to their pricing strategies.[1] Their theoretical model demonstrates that when export prices are negotiated in the exporter's currency, menu costs trigger threshold effects in the sense that there are bounds within (outside of) which pricing-to-market is not (is) observed. Interestingly, they also show that PTM is not interrupted by menu costs when export prices are denominated in the importer's currency.[2] Their empirical application focuses on pork meat exports from Canadian provinces to the U.S. and Japan. Canadian pork exporters were found to exercise market power on the U.S. market. The threshold model did not reject the null hypothesis of menu cost *given* the assumption that exports were invoiced in Canadian currency. Evidence of PTM behaviour in the Japanese market was weaker.

The objective of this paper is to jointly test the null hypothesis of menu cost, pricing to market and currency invoicing decisions. In contrast with previous studies, we find that the evidence of market power is weaker when not restricting a priori the currency invoicing decisions such that they are carried out in the exporting country's currency. The null hypothesis of no menu costs in the PTM relation is strongly rejected by the data. However, the data rejects the null hypothesis of domestic currency invoicing. The empirical model also fails to reject the joint null hypothesis of foreign currency invoicing and no PTM behaviour in two out of four PTM equations.

The next section lays out the theoretical model that explains price rigidity when exporters are facing menu costs related to changing their prices. The third section presents the empirical model and jointly tests the PTM assumption under menu cost and currency invoicing options using pork exports from two Canadian provinces (Quebec and Ontario) to Japan and the U.S. The last section presents concluding remarks.

2. The Theoretical model

In this section, the model of Larue *et al* (2003) is used to illustrate the impact of menu costs and currency invoicing on pricing-to-market. In the tradition of Klemperer's (1987) switching costs models, we assume that firms have a two-period planning horizon. For simplicity, it is assumed that there are only two firms selling differentiated products. Firm 1, based in country 1, enjoys a monopoly position in its domestic market, but it competes with firm 2 in country 2. Ignoring menu costs for the time being, and assuming that firm 1 sets its export price in its local currency, the profit of firm 1 at time *t* is defined as:

$$\pi_{1,t} = p_{11,t}q_{11,t}\left(p_{11,t}\right) + p_{12,t}q_{12,t}\left(p_{12,t}/e_t\,,\,p_{22,t}\right) - c_1\left(q_{11,t}\left(p_{11,t}\right) + q_{12,t}\left(p_{12,t}/e_t\,,\,p_{22,t}\right)\right) \tag{1}$$

where $p_{ij,t}$ and $q_{ij,t}$ are the price and quantity chosen by firm *i* to be sold in country *j* at time *t*, e_t is the exchange rate expressed in terms of country 1's currency per unit of country 2's currency and $c_i(.)$ is a cost function. Prices $p_{11,t}$ and $p_{12,t}$ are denominated in country 1's currency while $p_{22,t}$ is denominated in country 2's currency. Accordingly, the profit of firm 2 at time *t* is:

$$\pi_{2,t} = p_{22,t}q_{22,t}\left(p_{12,t}/e_t\,,\,p_{22,t}\right) - c_2\left(q_{22,t}\left(p_{12,t}/e_t\,,\,p_{22,t}\right),\omega_{2,t}\right) \tag{2}$$

It is assumed that $c_{iQ} \equiv \partial c_i\left(.\right)/\partial Q_i > 0$ where Q_i is the total quantity produced by firm *i*. It is also assumed that marginal cost is constant, *i.e.* $c_{iQQ} \equiv \partial^2 c_i\left(.\right)/\partial Q^2 = 0$.

With or without menu costs, it is assumed that play in country 2 is sequential with firm 1, the leader, announcing its price first. The home firm, Firm 2, enjoys the second-mover advantage on its own turf by announcing its price last. It also seems natural to have retailers in country 2 inquire about firm 2's price after getting firm 1's price quote, especially if it is costly for firm 1 to communicate with buyers in country 2. Conducting business in a foreign tongue with partners who have a distinct business culture can put an exporting firm at a disadvantage *vis-à-vis* home firms.

In the standard price leadership game, firm 1 picks prices $p_{11,t}$ and $p_{12,t}$ for each new realization of e_t, taking into account that firm 2 will be able to undercut its price. Defining firm 2's reaction function as $p_{22,t}\left(p_{12,t}/e_t\right) \equiv Arg\max \pi_{2,t}$, then firm 1's profit can be expressed as: $\pi_{1,t}\left(p_{11,t}, p_{12,t}, p_{22,t}\left(p_{12,t}/e_t\right)\right)$. The first order conditions for firm 1's profit maximization are:

$$\frac{\partial \pi_{1,t}}{\partial p_{11,t}} = q_{11,t} + \left(p_{11,t} - c_{1Q}\right)\left(\partial q_{11,t}/\partial p_{11,t}\right) = 0 \qquad (3)$$

$$\frac{\partial \pi_{1,t}}{\partial p_{12,t}} = q_{12,t} + \left(\frac{p_{12,t} - c_{1,Q}}{e_t}\right)\left(\frac{\partial q_{12,t}}{\partial\left(p_{12,t}/e_t\right)} + \frac{\partial q_{12,t}}{\partial p_{22,t}}\frac{\partial p_{22,t}}{\partial\left(p_{12,t}/e_t\right)}\right) = 0 \qquad (4)$$

Equations (3) and (4) indicate that the disadvantaged leader must equate its marginal revenues from domestic and export sales to its marginal costs. The domestic price equation in (3) can be manipulated to yield the more familiar monopoly rule: $p_{11,t}\left(1+1/\varepsilon_{11,t}\right) = c_{1Q}$. Equation (4) shows the direct and indirect effects of a change in $p_{12,t}$ on firm 1's profit. The former is simply the usual incentive of a firm to exploit the export demand for its product.

The indirect effect originates from firm 1's knowledge that firm 2 enjoys a strategic advantage in observing $p_{12,t}$ prior to choosing $p_{22,t}$.

The effect of the exchange rate on the equilibrium prices can be obtained by total differentiation of the first order conditions and the application of Cramer's rule. It can be shown that $dp_{11,t}/de_t = 0$ because the cost function is linear (*i.e.,* constant marginal and average costs) and no inputs are imported. These are the necessary conditions to analyze country 2's market in isolation from country 1's market, as is commonly assumed in the empirical literature.

Defining $\delta \equiv \frac{\partial q_{12,t}}{\partial\left(p_{12,t}/e_t\right)} + \frac{\partial q_{12,t}}{\partial p_{22,t}} \frac{\partial p_{22,t}}{\partial\left(p_{12,t}/e_t\right)} < 0$, a fluctuation of the exchange rate has the following impact on firm 1's export price expressed in its own currency:

$$\frac{dp_{12,t}}{de_t} = \frac{1}{|H|}\left(2\frac{\partial q_{11,t}}{\partial p_{11,t}}\frac{\delta}{e_t^{\,2}}\right)\left(2p_{12,t} - c_{1Q}\right) > 0 \tag{5}$$

where $|H| > 0$ from the second order condition. Furthermore, given that $p_{12,t} - c_{1Q} > 0$, it follows that the expression in (5) is unambiguously positive. Under these conditions, the ratio $p_{11,t}/p_{12,t}$ falls with e_t. This is the standard "pricing-to-market" outcome described in Bowen, Viaene and Hollander (1998). It is also possible to show that firm 1's export price expressed in country 2's currency actually falls as country 1's currency depreciates (*i.e.,* $\partial\left(p_{12,t}/e_t\right)/\partial e_t < 0$), an outcome usually referred to as an incomplete pass-through.

Let us now assume that when firm 1 wants to change $p_{12,t}$, it must incur a fixed menu cost *m*.[3] In the 2nd period, firm 1 makes its decision about changing its period 1 price or keeping it, with knowledge of the exchange rate in period 2. Hence it would not change its period 1 price in period 2 if :

$$\pi_{1,2}\left(p_{12,1}; e_2\right) \geq \pi_{1,2}\left(p_{12,2}; e_2\right) - m. \tag{6}$$

Forcing this relation to hold with equality enables us to define boundaries for period 2's exchange rate within which the firm will not find it profitable to change its price. The existence of these boundaries follows from the concavity of profit with respect to price. Hence, define the boundaries $e_2^{\min}\left(p_{12,1}, m\right)$ and $e_2^{\max}\left(p_{12,1}, m\right)$ whose difference is increasing with the menu cost.

Figure 1 illustrates these bounds using a numerical simulation under the assumptions of linear demand (*i.e.*, $q_{ij,t} = a - p_{ij,t}/e_t + \gamma_{ij} p_{jj,t}$), constant marginal cost and $e_1 = 1$.[4] The parameters γ_{12} and γ_{21} indicate the degree of substitutability between domestic and foreign products in country 2; the higher these parameters are, the less differentiated are firm 1 and 2's products from the consumers' perspective. As argued earlier, the exchange rate boundaries are widening in the menu cost. If the two goods are close substitutes $\left(\gamma_{12} = \gamma_{21} = 0.8\right)$, the boundaries are closer to one another than when differentiation is higher $\left(\gamma_{12} = \gamma_{21} = 0.5\right)$. In the latter case, the two firms face less stringent competition in country 2's market. As such, the variation in the exchange rate between the two periods needs to be large to make it profitable for firm 1 to change its price for a given menu cost.

In period 1, firm 1 knows that it will keep its period 1 price in period 2 as long as $e_2 \in \left[e_2^{\min}, e_2^{\max}\right]$. We assume that the firms' period 1 expectation of the exchange rate in period 2 is $E_1\left[e_2\right]$. For simplicity, let us assume that the exchange rate is drawn from a uniform distribution with support $\left[\underline{e}, \overline{e}\right]$, a mean of $\left(\overline{e} - \underline{e}\right)/2$, and that the parameter values are such that $\underline{e} < e_2^{\min} < e_2^{\max} < \overline{e}$. Hence, there is a probability $prob\left(e_2^{\min} < e_2 < e_2^{\max}\right) = \left(e_2^{\max} - e_2^{\min}\right)/\left(\overline{e} - \underline{e}\right) \in (0,1)$ that firm 1 will keep its period 1 price in period 2. Therefore, firm 1's optimization in period 1, given discounting parameter $\varphi < 1$, is as follows:

$$\max\ \pi_{1,1}\left(p_{11,1},p_{12,1};e_1\right)+\varphi\ prob\left(e_2^{\min}\le e_2\le e_2^{\max};\underline{e},\overline{e}\right)E\left[\pi_{1,2}\left(p_{11,2},p_{12,1};m\right)\right]. \tag{7}$$

The first order conditions are:

$$\frac{\partial\pi_{1,1}}{\partial p_{11,1}}=0;\qquad \frac{\partial\pi_{1,1}}{\partial p_{12,1}}+\varphi\ prob(.)\frac{\partial E\left[\pi_{1,2}\right]}{\partial p_{12,1}}+\varphi E\left[\pi_{1,2}\right]\frac{\partial prob(.)}{\partial p_{12,1}}=0. \tag{8}$$

The first expression reflects firm 1's ability to adjust its domestic market price without having to incur a menu cost. Hence, unless $e_2=e_1$, firm 1's domestic price will be subject to another optimization in period 2 and will change. The second expression makes it plain that the choice of export price must weigh the conditions prevailing in the market in period 1 against the ones expected to prevail in period 2. The extent by which firm 1's profit in the 2nd period must be taken into account in its 1st period optimization depends on the probability that the menu cost will be larger than the marginal gain from a price change when it is costless to do so. It should be noted that this probability is directly influenced by the menu cost *m* and by the choice of $p_{12,1}$ as indicated by the implicit definition of the bounds in (6). If firm 1 knew with certainty that the exchange rate would fall outside the bounds (*i.e.*, $e_2\not\in\left[e_2^{\min},e_2^{\max}\right]$, it would simply set $\partial\pi_{1,1}/\partial p_{12,1}=0$ in choosing $p_{12,1}$.

The introduction of menu costs implies that there is a probability that the export price, expressed in country 1's currency, will remain constant (*i.e.*, $p_{12,1}=p_{12,2}$) or will rise or fall depending on the realization of the exchange rate in period 2. A "small" depreciation of the domestic currency will not trigger changes in p_{11} and p_{12}, but it will make firm 1's export sales cheaper for foreign buyers because the ratio $p_{12,2}/e_2$ falls. As a result, we should not observe a pricing-to-market outcome in spite of our uncompetitive market structure. The same applies to a "small" appreciation of country 1's currency. The domestic-export price ratio would not respond to changes in exchange rate if the new exchange rate fell within the critical bounds. Systematic movements are expected when the exchange rate deviation is large enough to bring the new exchange rate above (below) the upper (lower) threshold. This is why threshold econometric techniques are most suited to empirically ascertain the validity of the theoretical model.

A key assumption in the model is that firm 1 gets paid in its own currency. If its price were denominated in country 2's currency, then interruptions in pricing-to-market outcomes, like the ones described above, would not be possible. Based on the theoretical model of Larue *et al.* (2003), we write the profit of firm 1 when it fixes its export price in country 2's currency as:

$$\pi_{1,t} = p_{11,t}q_{11,t}\left(p_{11,t}\right)+e_t p_{12,t}q_{12,t}\left(p_{12,t},p_{22,t}\left(p_{12,t}\right)\right)-c_1\left(q_{11,t}\left(p_{11,t}\right)+q_{12,t}\left(p_{12,t},p_{22,t}\left(p_{12,t}\right)\right)\right) \quad (9)$$

The first order conditions are quite similar to the ones derived previously:

$$\frac{\partial \pi_{1,t}}{\partial p_{11,t}} = q_{11,t}+\left(p_{11,t}-c_{1Q}\right)\frac{\partial q_{11,t}}{\partial p_{11,t}} = 0 \quad (10)$$

$$\frac{\partial \pi_{1,t}}{\partial p_{12,t}} = eq_{12,t}+\left(ep_{12,t}-c_{1,Q}\right)\left(\frac{\partial q_{12,t}}{\partial p_{12,t}}+\frac{\partial q_{12,t}}{\partial p_{22,t}}\frac{\partial p_{22,t}}{\partial p_{12,t}}\right) = 0\,. \quad (11)$$

The effect of a depreciation of country 1's currency on firm 1's domestic and foreign price is:

$$\frac{dp_{11,t}}{de_t} = 0\,;\quad \frac{dp_{12,t}}{de_t} = \frac{-2}{\left|G\right|}\left[\frac{\partial q_{11,t}}{\partial p_{11,t}}\delta\right]\frac{c_{1Q}}{e_t} \quad (12)$$

where $\left|G\right| > 0$ from the second order condition.

As in the other case, a depreciation of country 1's currency induces a decrease in firm 1's price denominated in country 2's currency. However, when the price is converted in country 1's currency, there is a positive relationship (*i.e.*, $\partial ep_{12}/\partial e > 0$). This implies that under these conditions, the ratio of prices (in country 1's currency), p_{11}/ep_{12}, falls with *e*. As expected, pricing-to-market behavior is robust to the denomination of export prices.

The introduction of menu cost implies the existence of exchange rate bounds within which firm 1 finds it more profitable not to update its first period price after observing the realization of the exchange rate in period 2. The rigidity of $p_{12,t}$ implies a larger increase in $e_t p_{12,t}$ and hence a *stronger* pricing-to-market effect than in the absence of a menu cost! It can then be foreseen that two very different exchange rate changes, one that keeps the exchange rate within the bounds and one that brings it outside, can trigger identical pricing-to-market responses. The implication for empirical analysis is that standard tests for a long run linear PTM relation are likely to be misleading. The rejection of a linear relation is likely to be misinterpreted as evidence of no long run relation between the export price and the exchange rate while in reality there would be one for "small" fluctuations in the exchange rate and one for "large" ones. Recall that when the export price is quoted in country 1's currency and in the presence of a significant menu cost, evidence of PTM did not become stronger, but disappeared. This contrast in response suggests that the null of significant thresholds outside of which long-run PTM behavior is observed is a joint test of menu cost and invoicing in one's own currency.

It must also be noted that there is a possibility that prices are quoted in a third-country currency that is not an interested party to the transaction. Larue *et al.* (2003) considered that case and found that the adjustment processes triggered by large exchange rate shocks (or in the absence of menu costs) are similar regardless of the choice of currency. However, small shocks (or when menu costs matter) can unleash very different adjustment processes that are difficult to track theoretically.

3. The Empirical model

While Larue *et al.* (2003) identified significant threshold effects in the PTM relation, they did not formally test for the own currency invoicing decision. From a theoretical perspective, the contrasts in the export price responses following changes in the exchange rate suggest that the null of significant thresholds outside of which adjustment to export prices are observed is a joint test of menu cost and own-currency invoicing.

Hansen's (2000) methodology is used to implement a two-regime PTM equation that is conditional on a threshold variable. The PTM equations are:

$$p_t = \theta_{0,1} + \theta_{1,1} e_t + \theta_{2,1} c_t + u_t \text{ if } \left|\Delta e_t\right| \leq \gamma \tag{13}$$

$$p_t = \theta_{0,2} + \theta_{1,2} e_t + \theta_{2,2} c_t + u_t \text{ if } \left|\Delta e_t\right| > \gamma \tag{14}$$

where *p* is the export price denominated in Canadian dollars, *e* is the exchange rate defined as units of foreign currency per Canadian dollar weighted by the destination consumer price index for food products, *c* is a marginal cost proxy and Δe_t is the threshold variable that is used to split the sample into two groups which are called regimes. The threshold is defined as the absolute value of the change in the exchange rate because the presence of menu costs defines boundaries for the exchange rate within which the firm will not find it profitable to change its price. The specification of the threshold implies that revising the export price in regime 1 is not profitable while additional profits from revising the export price in regime two is greater than the menu costs. The parameters $\theta_1' = \left[\theta_{0,1} \quad \theta_{1,1} \quad \theta_{2,1}\right]$, $\theta_2' = \left[\theta_{0,2} \quad \theta_{1,2} \quad \theta_{2,2}\right]$ and γ need to be estimated. The sample length is denoted by T.

Estimation of the model in (13) and (14) is done by sequential least squares. The model can be written in a single equation using a dummy variable $d_t(\gamma) = \left\{\left|\Delta e_t\right| \leq \gamma\right\}$ such that $\mathbf{X}_t(\gamma) = \mathbf{X}_t d_t(\gamma)$; where $\mathbf{X}_t$ is the vector of independent variables in (13) and (14). The PTM equation is:

$$p_t = \theta' X_t + \delta_T X_t(\gamma) + e_t \tag{15}$$

Define the OLS estimators $\hat{\theta}$ and $\hat{\delta}_T$ conditional on γ. The parameter γ is assumed to be restricted to a bounded set $\Gamma \equiv \left[\underline{\gamma}, \overline{\gamma}\right]$ that is approximated by a grid defined by:

$\Gamma \cap \{q_{(1)}, \ldots, q_{(T)}\}$. The estimation procedure requires $N < T$ evaluations of equation (15); where N is selected such that the 10% upper and lower percentiles of $\{q_{(1)}, \ldots, q_{(T)}\}$ are not included in Γ. A natural estimator for γ is to minimize the sum of squared errors $S(\gamma) = \sum_{t=1}^{T} \hat{e}_t^2(\gamma)$ such that $\hat{\gamma} = \arg\min_{\gamma \in \Gamma} S(\gamma)$.

We consider two distinct scenarios with respect to the choice of currency invoicing. The first scenario presumes that Canadian pork exporters invoice their exports in Canadian dollars. When the exchange rate variation is large enough such that it is profitable to revise the export price (*i.e.* $|\Delta e_t| \geq \gamma$), the variation in the exchange rate will produce a variation in the export price which will be linearly related to the exchange rate such that $-1 < \theta_{1,2} < 0$ under the PTM hypothesis. In the other case (*i.e.* $|\Delta e_t| < \gamma$), menu costs will prevent the adjustment of the export price and thus the 1st regime will hold with $\theta_{1,1} = 0$. In the 2nd scenario, pork exports are invoiced in the destination currency. The theoretical model shows that when exchange rate variations are small, no adjustment will be made to the export price in foreign currency, but the export price denominated in Canadian dollars will vary proportionally with the exchange rate and thus $\theta_{1,1} = -1$. When the exchange rate variation is large enough, the export price will adjust such that $-1 < \theta_{1,2} < 0$.

The null hypothesis of no menu costs is a test of $\theta_1' = \theta_2'$. The asymptotic theory of thresholds is complex; but Hansen (2000) derives the asymptotic distribution of the threshold parameter and the slope coefficients under certain conditions. If one assumes that the threshold parameter is known, the two-step least squares estimator of the regression coefficients converges to a normal distribution. However, this distribution is likely to under-represent the uncertainty in the parameters in finite sample or when the threshold effect is small. Hansen (2000) suggests working with conservative bounds to reduce the probability of wrongly rejected the null. Moreover, it is often the case that inference about the threshold effect is needed. If the threshold effect is represented by $\delta_T \equiv \theta_1' - \theta_2'$, Hansen shows that one strategy is to assume that $\delta_T \rightarrow 0$ as the sample size, *T*, tends to infinity. The null of $\gamma = \gamma_0$ can be tested with a likelihood ratio test whose non-standard distribution can be conveniently computed in closed-form. However, there is no reason to believe in our context that menu costs will disappear as the sample size increases. If δ_T is fixed as *T* increases, the asymptotic distribution of the likelihood ratio test under $\delta_T \rightarrow 0$ must be regarded as asymptotically conservative if the error terms are normally distributed.

While Hansen (2000) provides nice improvements in the asymptotic theory of threshold models, there are still important ambiguities that can have important implications in our setting. One alternative is to use bootstrap methods to estimate the distribution of the estimators and test statistics. However, the test statistics are not asymptotically pivotal in the sense that their distribution will depend on unknown population parameters. In that case, bootstrap estimates of the statistic's distribution converge at the same rate as conventional asymptotic approximations (Horowitz, 2001). Improvements in the rate of convergence of the bootstrap can be achieved using pre-pivoting methods introduced by Beran (1988) and summarized in McCullough and Vinod (1998).

To circumvent the aforementioned ambiguities in asymptotic theory, we proceed with bootstrap methods to test the various hypotheses of the theoretical model. Four sets of hypotheses will be tested. First, the null hypothesis of no menu costs will be tested $H_0 : \theta_1' = \theta_2'$. If we accept the alternative hypothesis of significant thresholds in the PTM equation and keeping in mind that *e* in the empirical section is defined in terms of Can\$, we can test: a) the null hypothesis of no PTM effect or $H_0 : \theta_{1,2} = 0$; b) the null hypothesis of Canadian (foreign) currency invoicing or $H_0 : \theta_{1,1} = 0$ $\left(H_0 : \theta_{1,1} = -1\right)$; and c) the joint null hypothesis of no PTM and Canadian (foreign) currency invoicing $H_0 : \theta_{1,1} = \theta_{1,2} = 0$ $\left(H_0 : \theta_{1,2} = 0; \theta_{1,1} = -1\right)$.

4. Data and estimation

PTM equations are specified for exports from two different Canadian provinces to two destinations over the period beginning in January 1992 and ending in December of 2003. Export data were obtained from Statistic Canada while the exchange rate and the consumer price index for food items were collected in publications from each country's central bank. The marginal cost proxy in (13) and (14) are the monthly hog prices in each province and were obtained from Agriculture and Agri-food Canada. The United States and Japan represent the most important market for Canadian pork exporters. Exports of each province are depicted in Figures 2a and 2b. Quebec is the largest pork meat exporting province but growth in exports is observed in Ontario as well over the period. Figures 3a and 3b present the unit export values by source for the Japanese and U.S. markets. Differences in export unit values are especially important at the beginning of the sample but they tend to shrink over time.

Figure 4 plots the hog price in each province from January 1992 to December 2003. Although prices in each province follow a similar trend, there are some differences in the three series that can be attributed to different hog marketing institutions (Larue *et al.*, 2000). Figures 5 presents the value of the real exchange rates (units of foreign currency per Can$). There is a steady depreciation in the real value of the Canadian currency with respect to the U.S. currency over the entire sample. Finally, there are wilder variations in the real value of the Canadian dollar with respect to the Japanese yen.

As it is usually the case with monthly time series, the degree of integration in each variable is an important preoccupation. The first step of the empirical strategy is thus to investigate the stochastic properties of the data. To this end, the Augmented Dickey-Fuller (ADF) test is implemented by regressing the first difference of a series on the lagged level of the series, a constant and, if needed, a time trend and *w* lagged first differences of the series to insure that the residuals are white noise:

$$\Delta y_t = \alpha + \beta t + \rho y_{t-1} + \sum\nolimits_{j=1}^{w} \upsilon_j \Delta y_{t-j} + \varepsilon_t \qquad (16)$$

The null of non-stationarity of the ADF test involves testing a zero restriction on ρ.

The ADF test was implemented on the logarithmic transformation of the real exchange rate, export unit values and hog prices in each province. The results are reported in the second column of Table 1. The first column indicates whether a time trend (T) or no time trend (NT) were used. Following Hall's (1994) recommendations, we used the SBC information criterion to select the lag length in (16) because it makes the ADF test more powerful in small samples than the AIC criterion. The null hypothesis of a unit root is rejected for all variables at the 90% or higher confidence level.

Even though the null hypothesis of a unit root is rejected in favour of stationarity, the stationarity test developed by Kwiatkowski *et al.* [hereafter referred to as KPSS, (1992)] was also carried out for each series. The KPSS testing procedure differs from standard unit root tests since the null hypothesis is that of stationarity in the level of a series. The KPSS test involves estimating the equation:

$$y_t = \delta t + \zeta_t + \varepsilon_t\,;\;\; \zeta_t = \zeta_{t-1} + u_t\,;\;\; u_t \sim iid\left(0, \sigma_u^2\right) \qquad (17)$$

The null hypothesis of trend stationarity is about the validity of a zero restriction on σ_u^2. Testing the null of level stationarity instead of trend stationarity involves regressing the series on a constant instead of a trend variable. The KPSS test is computed using the Bartlett kernel to account for the potential correlation of the residuals with a bandwidth selected using the procedure suggested by KPSS; *i.e.* $l = trunc\left\{4\left(0.01T\right)^{0.25}\right\}$. The third column of table 1 reveals that the null hypothesis of stationarity is rejected for all export unit values at the 95% confidence level. In contrast, hog prices in each province and the real exchange rate between the Canadian and Japanese currencies seem stationary. Overall, the ADF and KPSS tests yield conflicting evidence.

Carrion-i-Silvestre *et al.* (2001) showed that simultaneous testing of the null hypotheses of stationarity and unit root should not be conducted using standard marginal critical values for each test. They studied a Confirmatory Data Analysis (CDA) method by computing critical values for the joint confirmation hypothesis of a unit root. They argue that their set of critical values generates more accurate results than standard critical values for each test when the data generation process is integrated of order one. The CDA shows that only two of the eight variables are integrated of order one. Hence, the analysis proceeds as if the variables are stationary.

Table 2 presents the OLS estimates of the PTM equations for pork exports from Quebec and Ontario to each destination (U.S. and Japan). The two regimes are identified separately in Table 2. The coefficient estimate and its standard error between parentheses are in the first line of each cell. The number underneath is the *p-value* for the null hypothesis of zero coefficient. As mentioned previously, the distribution of the coefficients is non-standard and bootstrap methods were implemented for inference purposes. The bootstrap procedure is the following. The independent variables in (15) are treated as fixed as well as the threshold variable. The regression vector of residuals $\hat{\mathbf{e}}^*$ constitutes the empirical distribution that is used for the bootstrap. A sample of T observations is drawn with replacement from the empirical distribution under the null hypothesis considered. Using the bootstrapped sample, the model in (15) is estimated with and without the restriction implied by the hypothesis and a test statistic is computed: $t_j^* = \hat{\theta}^* \big/ \sigma\left(\hat{\theta}^*\right)$. This procedure is repeated 2000 times[5] and the percentage of draws for which the simulated statistic exceeds the actual test statistic is computed.[6] This value is the bootstrap estimate of the asymptotic *p-value* for the *t*-statistic under the null hypothesis.

The point estimate for the real exchange rate in each regime of the PTM equations always has the expected algebraic sign in Table 2. The coefficient of the hog price is always significant in the 2nd regime of the PTM equations but the point estimate in the first regime is not statistically different from zero in two PTM equations. Moreover, the two coefficients have a positive algebraic sign which runs counter to the intuition that an increase in processors' marginal cost will increase their export price and lower their sales.[7]

The null hypothesis of no threshold is tested using the likelihood ratio statistic proposed by Hansen (2000), $LR(\gamma=0)=T\left(S(\gamma=0)-S(\hat{\gamma})\right)/S(\hat{\gamma})$; where $S(\hat{\gamma})$ and $S(\gamma=0)$ are respectively the sum of squared residuals for models with and without threshold. The *p-value* is computed by simulating a sample of *T* observations under the null hypothesis and computing the number of times that the bootstrap statistic falls below the actual *LR* statistic. The *LR* statistic always rejects the null hypothesis of no threshold at the 5 percent significance level.[8] It should be noted that the threshold value in Table 2 does not provide a direct estimate of menu costs. The threshold variable is function of menu cost but also of the structural parameters of demand in the importing country. Hence, a large threshold estimate does not necessarily imply large menu costs. This may explain why the thresholds are larger in the U.S market than in the Japanese market for Quebec and Ontario pork exports. Larue, Gervais and Rancourt (2003) estimated larger menu costs in the Japanese market than in the U.S. market for Canadian exporters and justified their findings by pointing out that the U.S. and Canada share a common border, common language and similar institutions; all factors that suggest small menu costs. Our results do not invalidate their conclusions. Finally, the threshold estimates in the Japanese market for Quebec pork exports and Ontario pork exports are equal.

Given that we can reject the null hypothesis of no menu costs, we can test the currency invoicing and PTM hypotheses individually. The null hypothesis of no PTM is a test about the statistical significance of the coefficient of the real exchange rate in the 2nd regime of the PTM equation $\left(H_0 : \theta_{1,2} = 0\right)$. The null hypothesis of no-PTM behaviour is rejected for Quebec exports to Japan. There is no significant evidence of PTM behaviour for Ontario pork exports. It is interesting to contrast the current results to linear PTM equations that do not include menu costs. We already argued that the empirical model strongly support the hypothesis of menu costs; but of particular interest is the statistical significance of the PTM coefficient in the linear specifications. The literature cited at the beginning of the paper routinely finds significant PTM effects and most of the empirical models only account for a linear relationship between the exchange rate and the export price. Table 3 provides the coefficient estimates of the linear PTM equations with their standard error between parentheses. The *p-value* for the hypothesis of a zero coefficient for the real exchange variable is always lower than 0.01 using conventional asymptotic theory. Hence, a linear model finds significant PTM effects in all four equations.

The null hypothesis of domestic or own currency invoicing is a test about the significance of the coefficient of the real exchange rate $\left(H_0 : \theta_{1,1} = 0\right)$. This hypothesis is strongly rejected in all PTM equations. Finally, the foreign currency invoicing hypothesis is a test about the plausibility of: $H_0 : \theta_{1,1} = -1$. Although the *p-values* for this test are not reported in Table 2, the bootstrap simulations indicate that this hypothesis is rejected for both Quebec PTM equations. We found support for foreign currency invoicing only for Ontario pork exports. The *p-values* for the U.S. and Japanese markets are respectively 0.679 and 0.127.

The previous two hypotheses (currency invoicing and no PTM) can also be tested jointly. Define the null hypothesis of domestic currency invoicing and no-PTM as:

$H_0:\theta_{1,1}=\theta_{1,2}=0$. The inference strategy is to write the restrictions on the parameter in (15) as $H_0:\mathbf{R}\Theta=\mathbf{r}$, where $\Theta=\begin{bmatrix}\boldsymbol{\theta} & \boldsymbol{\delta}_{\mathbf{T}}\end{bmatrix}'$ is a 6×1 vector and the matrix **R** selects the appropriate elements from the vector θ to be restricted according to **r**. Under the null hypothesis, we have that $\mathbf{R}=\begin{bmatrix}0 & 1 & 0 & 0 & 0 & 0\\ 0 & 0 & 0 & 0 & 1 & 0\end{bmatrix}$ and $\mathbf{r}=\begin{bmatrix}0 & 0\end{bmatrix}'$. The test statistic is $F=\left(\mathbf{R}\hat{\Theta}-\mathbf{r}\right)'\left[\mathbf{R}\operatorname{cov}\left(\hat{\Theta}\right)\mathbf{R}'\right]^{-1}\left(\mathbf{R}\hat{\Theta}-\mathbf{r}\right)\Big/2$. The inference is made possible by using bootstrap samples as described previously. The joint null hypothesis of domestic currency invoicing and no pricing to market behaviour is strongly rejected by the data in all four cases as indicated by the *p-value* of the *F* statistics in the next to last row in Table 2.

A similar testing procedure is carried out for the joint foreign currency invoicing and no PTM hypotheses. Interestingly, this latter set of hypotheses is not rejected when analyzing pork exports from Ontario to either the U.S. or Japanese market. Hence, the data reveals that Ontario pork exporting firms invoice in the currency of their customers and that they do not possess market power. The evidence about the joint hypothesis of no PTM and foreign currency invoicing for Quebec pork exports is not clear given the *p-value* of 0.053. The joint hypothesis of no PTM and foreign currency invoicing is rejected for Quebec pork exports to Japan. Given we rejected the single null of no PTM for Japan, the results suggest that pork exporters could invoice their exports to Japan in a third country currency. Perhaps, the evidence is consistent for Quebec pork exports is consistent with US$ invoicing transactions.

5. Conclusion

This paper applied the Pricing-to-Market (PTM) theoretical framework developed in Larue, Gervais and Rancourt (2003) to investigate whether Canadian pork exporters exercise market power in export markets. The theoretical model introduces menu costs that make it costly for exporters to revise their prices in response to changes in exchange rates. This introduces a non-linearity between the exchange rate and the export price. This non-linearity motivates the specification of a two-regime PTM model to analyze the pricing decisions of pork exporters from two Canadian provinces to the U.S. and Japan.

The empirical model finds that the null hypothesis of no menu costs is rejected for all four PTM equations. Evidence of PTM behaviour is weak for pork exports from Ontario. to Japan and the U.S. Moreover, the null hypothesis of foreign currency invoicing and no PTM effect could not be rejected in the two Ontario PTM equations. Evidence of PTM behaviour is stronger for Quebec exports. Significant PTM effects are found in the Japanese market. The data rejects the null hypothesis involving either currency invoicing alternative (in yen and Can$). It thus suggests that third country invoicing procedure should be investigated further. The results for the U.S. market are mixed. We do not reject the null of no PTM and US$ currency invoicing with a *p-value* of 0.053. When currency invoicing and PTM effects are tested separately, the evidence overwhelmingly rejects PTM behaviour in the U.S. market and also rejects both currency invoicing hypotheses.

These results contrast with the ones in Larue *et al.* (2003). They found significant market power exercised by Canadian pork exporters under the assumption that exports are invoiced in Canadian currency and using a different time period than in the current study. Moreover, they did not formally test the statistical significance of the PTM coefficients in their empirical models perhaps because coefficients of threshold cointegration models have non-standard distributions. Another implication of the current framework is that linear models can yield significant estimates of PTM effects while in reality exporters do not possess any market power.

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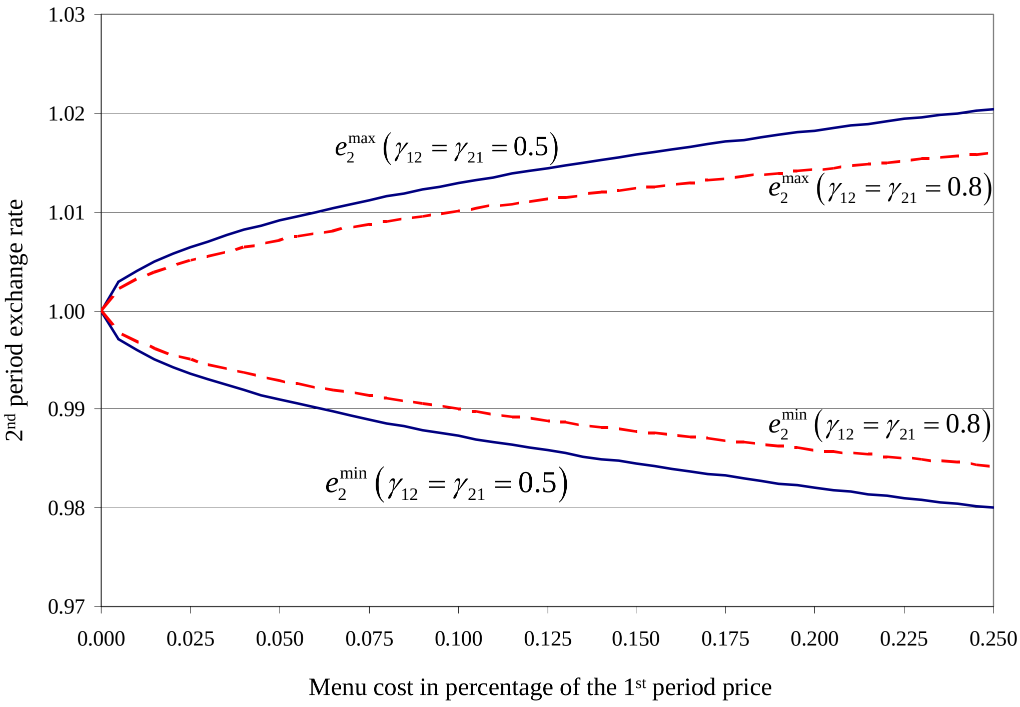

Figure 1. Simulated exchange rate band as a function of a fixed menu cost *m*.

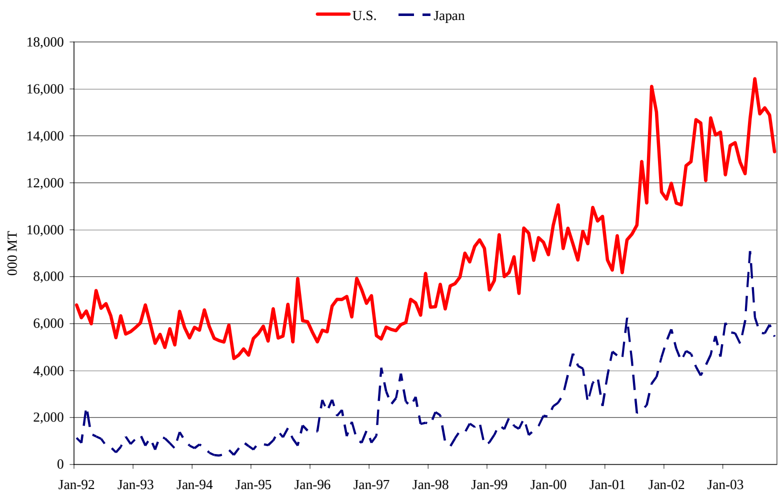

Figure 2a. Pork meat exports from Quebec to the U.S. and Japan from January 1992 to December 2003

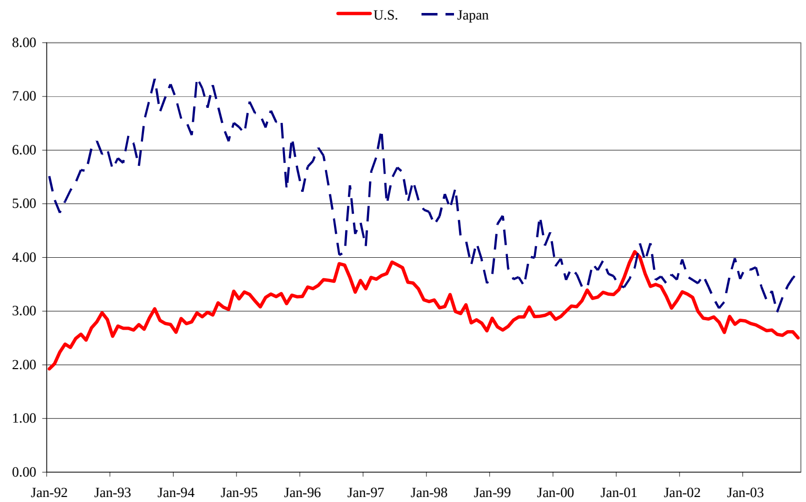

Figure 3a. Quebec export unit values to the U.S. and Japan from January 1992 to December 2003

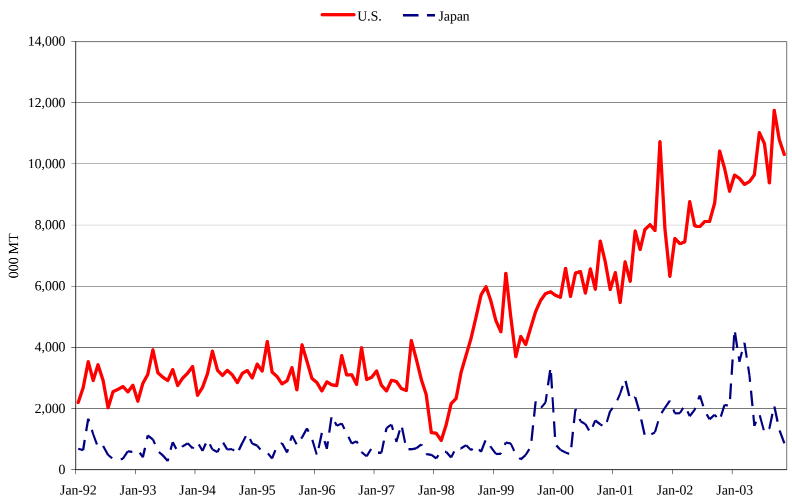

Figure 2b. Pork meat exports from Ontario to the U.S. and Japan from January 1992 to December 2003

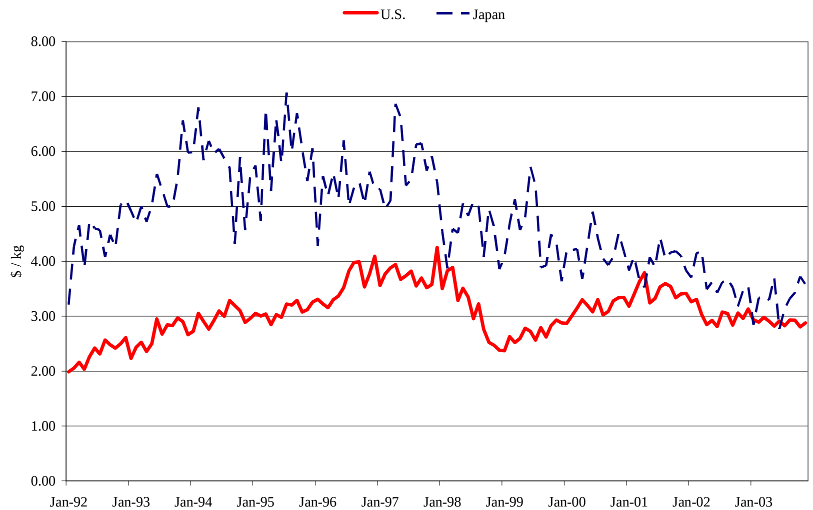

Figure 3b. Ontario export unit values to the U.S. and Japan from January 1992 to December 2003

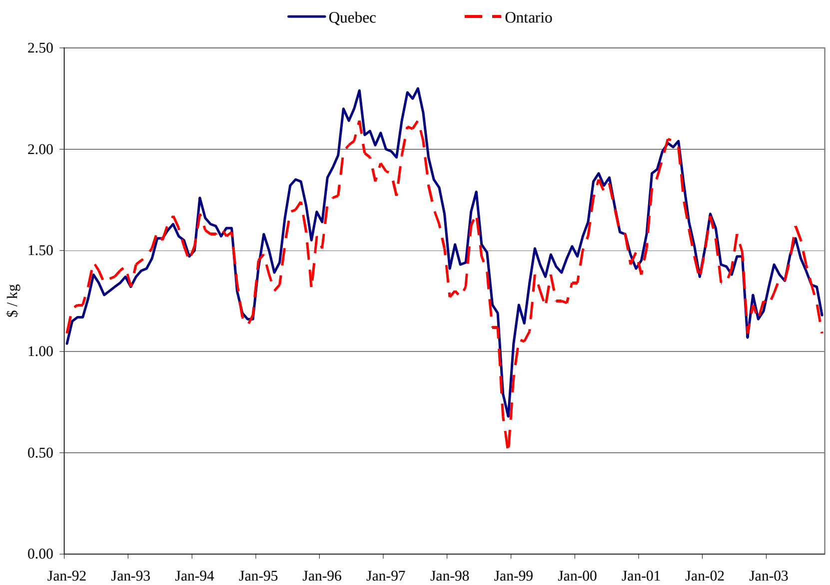

Figure 4. Hog prices in Quebec, Ontario and Manitoba from January 1992 to December 2003

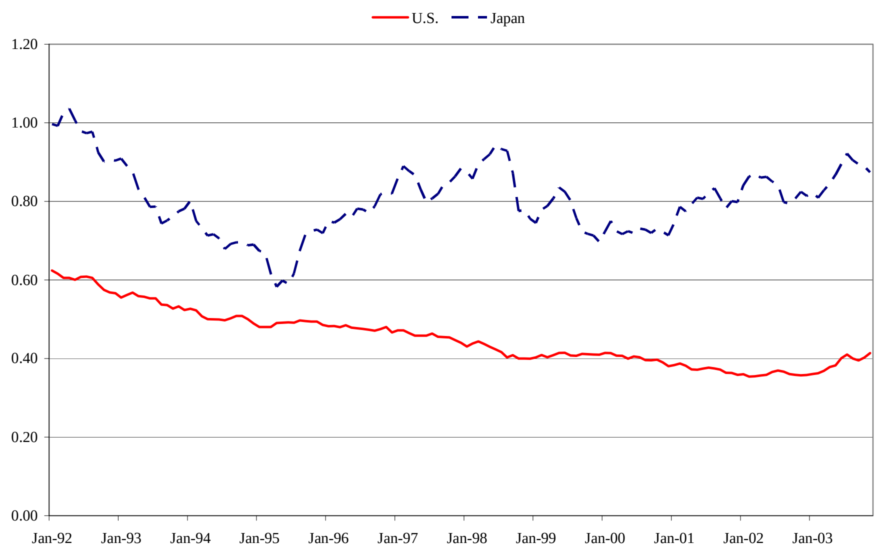

Figure 5. Value of the foreign currency per Can$ weighted by the consumer food price index

Table 1. Unit root testing

Variables	ADF test		KPSS test	Joint confirmation of a unit root
	Lag	Statistic		
Quebec				
Export price to the U.S. (NT)	0	-3.37*	0.46*	No
Export price to Japan (T)	0	-4.61*	0.22*	No
Hog price (NT)	0	-3.06*	0.22	No
Ontario				
Export price to the U.S. (NT)	2	-2.64**	0.59*	Yes
Export price to Japan (T)	1	-4.33**	0.47*	No
Hog price (NT)	1	-3.67*	0.14	No
U.S. real exchange rate (NT)	0	-2.70**	2.78*	Yes
Japan real exchange rate (NT)	1	-2.63**	0.23	No

The symbols * and ** denote rejection of the null hypothesis at the 95% and 90% confidence levels respectively. Critical values for the ADF test were obtained from Davidson and Mackinnon (1993) and the KPSS critical values were obtained from Kwiatkowski *et al.* (1992). The critical values for the Joint hypothesis of a unit root were taken in Carrion-i-Silvestre *et al.* (2001).

Table 2. Estimates of the PTM equation and inference

	Quebec		Ontario	
Parameters	U.S.	Japan	U.S.	Japan
1st regime				
constant	-1.83 (0.35)	1.24 (0.10)	-2.79 (0.53)	1.09 (0.14)
	0.000	0.000	0.001	0.000
real x-rate	-0.73 (0.09)	-0.56 (0.13)	-0.95 (0.13)	-0.77 (0.17)
	0.000	0.000	0.000	0.000
farm price	-0.44 (0.07)	-0.10 (0.09)	-0.57 (0.08)	0.00 (0.11)
	0.000	0.415	0.000	0.980
2nd regime				
constant	1.08 (0.30)	-0.14 (0.09)	0.99 (0.31)	0.43 (0.13)
	0.004	0.101	0.012	0.006
real x-rate	-0.02 (0.07)	-0.15 (0.07)	-0.02 (0.08)	-0.01 (0.09)
	0.781	0.000	0.852	0.899
farm price	-0.37 (0.08)	0.75 (0.08)	-0.37 (0.08)	0.36 (0.09)
	0.000	0.000	0.000	0.002
Threshold	1.563	0.973	1.038	0.973
Likelihood ratio test of no menu costs	34.36	65.45	46.12	17.23
	0.000	0.000	0.000	0.017
Joint no PTM and domestic currency invoicing	34.35	11.91	25.55	10.62
	0.000	0.000	0.000	0.001
Joint no PTM and foreign currency invoicing	4.55	7.85	0.101	0.99
	0.053	0.007	0.955	0.635

Table 3. Linear PTM equations

	Quebec		Ontario	
Parameters	U.S.	Japan	U.S.	Japan
constant	-0.53	0.88	0.23	0.73
	(0.29)	(0.09)	(0.30)	(0.11)
real x-rate	-0.40	-0.43	-0.20	-0.21
	(0.07)	(0.07)	(0.07)	(0.08)
farm price	-0.41	0.20	-0.41	0.17
	(0.07)	(0.07)	(0.07)	(0.08)

Endnotes

[1] Menu costs were first used in macro models to account for price rigidities (*e.g.*, Akerlof and Yellen, 1985).

[2] Knetter (1994) also argues that asymmetric adjustments can be observed when firms face capacity constraints in distribution networks or quantitative restrictions. He conjectured that in such cases PTM should be greater after a depreciation of the exporter's currency. PTM should also be greater after an appreciation of the exporter's currency when an exporting firm is trying to increase its market share subject to the threat of a trade restriction.

[3] This would be the case for instance when translating and legal services must be contracted to implement a price change. Alternatively, a price change may cause an interruption in deliveries which may differ in length depending whether the price is increased or decreased. In this instance, the menu costs would be asymmetric.

[4] Detailed numerical simulations are available from the authors upon request.

[5] Andrews and Buchinsky (2000) propose a methodology to choose the number of bootstrap replications to achieve a desired level of accuracy. However, it relies upon the asymptotic distribution of the quantities of interest and is unfortunately inapplicable in the present context.

[6] As alluded to earlier, the pre-pivoting method was also considered in order to conduct statistical inference. The statistic t_j^* is not asymptotically pivotal because its distribution depends on unknown parameters; but it still is a valid approximation of the 1st order asymptotic distribution. Pre-pivoting methods (or bootstrap iterations) are described in Horowitz (2001). Pre-pivoting entails drawing bootstrap samples from bootstrap samples to create an asymptotically pivotal statistic. McCullough and Vinod (1998) present a convincing case for using pre-pivoting methods. However, our estimation procedure complicates the application given that it involves many recursive regressions. For example, denote the number of bootstrap samples from the empirical distribution by J and the number of bootstrap sample from the bootstrap distribution by K. McCullough and Vinod (1998) suggest that the product of J and K should be of an order of magnitude slightly greater than T^3. In our case, a single hypothesis test would involve over 330 millions regressions given the sequential least square procedure. This was deemed too demanding and this is why the inference was carried out using the single bootstrap method. However, the double bootstrap with J = 999 and K = 200 was carried out as an experiment. It produced *p-values* significantly different in some instances and is worth investigating in future research.

[7] As Larue *et al.* (2004) and Gervais *et al.* (2004) explained, there may exist significant capacity constraints in the hog/pork industry due to lags between hog production and marketing decisions. In particular, the current hog price may not be strongly correlated with producers' supply of live hogs. Marketing lags are likely to influence PTM outcomes and it deserves to be investigated in future research.

[8] It is also interesting to compare the bootstrap critical values with the asymptotic critical values reported in Hansen (2000). The empirical critical values are always larger than the corresponding critical values. For example, the empirical critical value of the likelihood ratio test of $\gamma = 0$ at the 95% confidence level is never inferior to 14.45 while the asymptotic critical value is 8.75 (Hansen, 2000, p. 582).