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2 June 2006

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Abstract: This paper investigates price discriminating 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 price adjustments are not (are) observed. Conversely, the pass-through 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 two Canadian provinces to the U.S. and Japan. Hansen's (2000) threshold estimation procedure is used to jointly test for currency invoicing and incomplete pass-through in the presence of menu costs. Inference is conducted using the bootstrap with pre-pivoting methods to deal with nuisance parameters. The existence of menu cost is supported by the data in three of the four cases. It also appears that Quebec pork exporters price discriminate and invoice in Japanese yen their exports to Japan. Manitoba exporters also seem to follow the same invoicing strategy, but their ability to increase their profit margin in response to large enough own-currency devaluations is questionable. Our currency invoicing results for sales to the U.S. are consistent with subsets of Canadian firms using either the Canadian or U.S. currency.

Keywords: Price discrimination, Currency invoicing, Menu costs, Threshold estimation, Bootstrap, Pork exports.

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

A JOINT TEST OF PRICE DISCRIMINATION, MENU COST AND CURRENCY INVOICING

1. Introduction

Since the pioneering work of Krugman (1986) on the concept of pricing-to-market (PTM), a considerable literature documenting evidence of price discrimination and incomplete exchange rate pass-through in international markets has emerged. Knetter (1989) was the first to document evidence of price discrimination in his analysis of the pricing strategies of German and American firms in response to exchange rate changes. Since then, the literature has followed two different paths. One strand of the literature has focused on macroeconomic implications. For example, Bergin and Feenstra (2001) blame price discriminating behaviour for the high degree of volatility in exchange rates. The other, and more popular, strand of the literature has concentrated on microeconomic implications. At the manufacturing level, Uctum (2003) and Sasaki (2002) have analyzed the price discriminating 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. PTM behaviour has also been observed in agri-food sectors. For example, Carew and Florkowski (2003) found evidence of price discrimination by Canadian and U.S. exporters of agri-food products. Brown (2001) found PTM effects in the 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.

Even though trading firms are confronted to the currency invoicing issue every day as they negotiate prices in their own currency, the importing country's currency or in a third country's currency, the literature on this subject is surprisingly thin. As indicated in the survey found in Bowen, Hollander and Viaene (1998), most studies attempt to rationalize the apparent stylized fact that most 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). Price discrimination is intrinsically linked to currency invoicing, yet few authors have formally tied the two concepts. One notable exception is Sato (2003). He developed an empirical model that distinguishes short-run from long-run pricing strategies used by Japanese exporters. Exporters can stabilize their export prices by adjusting their profit margin and invoicing in the importer's currency. Standard estimation techniques are adequate to analyze the pure/long-run price discrimination effect conditioned by the curvature of the importers' demand function, but cointegration techniques are required to disentangle short-run effects. The latter are of interest because short run volatility is influenced by the currency used for invoicing purposes.

Menu costs were first used in macro models to account for price rigidities (*e.g.,* Akerlof and Yellen, 1985). Menu costs are incurred by firms whenever they make changes to their pricing strategies. While the costs of reprinting restaurant menus and mail-order catalogues are obvious menu costs examples, one could wrongly infer that menu costs affect only a narrow segment of businesses. In fact, a wide range of costs could be construed as menu costs like, for instance, legal advice and translation services in contract negotiations, costs of communicating price changes to intermediaries, etc. Accordingly, it seems most appropriate to introduce menu costs in a price discrimination analysis. Intuitively, one could think that large enough menu costs could discourage price adjustments that would otherwise be observed.

The objective of this paper is to assess the plausibility of individual and joint hypotheses regarding the significance of menu costs, price discrimination and the choice of currency for invoicing purposes. To shed some light on these issues, we develop a simple theoretical model that provides the necessary insights regarding the parametric restrictions needed to conduct our empirical analysis which focuses on Quebec and Manitoba pork exports to the United States and Japan.[1] Pork processing is a highly concentrated business that has been the object of many market power studies.[2] Accordingly, the extent by which large pork processors are able to price discriminate on foreign markets is a pertinent empirical question. The choice of two destinations is motivated by the conjecture that menu costs are likely to be less important for transactions involving firms based in countries that are closely integrated. The United States and Japan are also the largest importers of Canadian pork. The null hypothesis of no menu costs is soundly rejected by the data in three of the four pass-through equations and the same can be said about the null hypothesis of domestic currency invoicing. The empirical model also fails to reject the joint null hypothesis of foreign currency invoicing and no price discriminating behaviour in two out of the four pass-through equations.

The next section lays out the theoretical model that rationalizes price rigidity when exporters are facing menu costs. It is shown that the implications of the menu costs on the pass-through vary depending on the choice of currency used for invoicing. The third section describes the two-regime threshold model, the data, and it also describes the strategy used to make statistical inference in the absence of an adequate asymptotic theory. The fourth section reports on the estimation of the pass-through equations as well as about the individual and joint tests regarding price discrimination, the absence of menu cost and currency invoicing options. The fifth and last section presents concluding remarks.

2. The Theoretical Model

In this section, we develop a simple PTM model to assess the impact of menu costs and currency invoicing. 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 linearly increasing in output. 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 $c_{iQ} \equiv \partial c_i(.)/\partial Q_i > 0$ where Q_i is the total quantity produced by firm *i*. It is also assumed that marginal cost is constant, *i.e.* $c_{iQQ} \equiv \partial^2 c_i(.)/\partial Q^2 = 0$.

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:

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

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

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: $p_{11,t}\left(1+1/\varepsilon_{11,t}\right) = c_{1Q}$, where $\varepsilon_{11,t}$ is the demand elasticity facing firm 1 at home. 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 in output (*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 PTM 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 2$^{\text{nd}}$ period, firm 1 must decide whether to change its period 1 price and incur the menu cost or to keep it constant, 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 $(\gamma_{12} = \gamma_{21} = 0.8)$, the boundaries are closer to one another than when differentiation is higher $(\gamma_{12} = \gamma_{21} = 0.5)$. 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:

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

The first order conditions are:

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

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 if it were costless to do so. It is worth restating that this probability is directly influenced by the menu cost *m* and by the positioning of the exchange rate in period 1 in relation to the range of possible exchange rates in period 2. If firm 1 knew with certainty that the exchange rate would fall outside the bounds (*i.e.,* $e_2 \notin \left[e_2^{\min}, e_2^{\max}\right]$), 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 an incomplete exchange rate pass-through 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 above theoretical model is that firm 1 gets paid in its own currency. If its price were denominated in *country 2*'s currency, then interruptions in PTM outcomes, like the ones described above, would not be possible. To see this, we write the profit of firm 1 when it fixes its export price in country 2's currency as:

$$\begin{aligned}\pi_{1,t} = {} & p_{11,t}q_{11,t}\left(p_{11,t}\right)+e_t p_{12,t}q_{12,t}\left(p_{12,t}, p_{22,t}\left(p_{12,t}\right)\right) \\ & -c_1\left(q_{11,t}\left(p_{11,t}\right)+q_{12,t}\left(p_{12,t}, p_{22,t}\left(p_{12,t}\right)\right), \omega_{1,t}\right)\end{aligned} \tag{9}$$

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

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

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

It can be shown that the effect of a depreciation of country 1's currency on firm 1's domestic price remains zero given our assumptions regarding the technology (i.e., constant marginal cost and domestically-produced inputs).[5] For the export price, expressed in country 2's currency, we find:

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

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 under the same cost/technology assumptions, we find a positive relationship (*i.e.,* $\partial ep_{12}/\partial e > 0$).[6] This implies that under these conditions, the ratio of prices (in country 1's currency), p_{11}/ep_{12}, falls with e which confirms that PTM 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* response 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, could trigger identical price adjustments. 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, price adjustments did not become stronger, but disappeared. This contrast in response suggests that the null of significant thresholds outside of which long-run price adjustments are 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. In this case, 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 and the Estimation Strategy

The theoretical two-regime export price response induced by significant thresholds outside (inside) of which price adjustment are (not) observed can be construed as a joint hypothesis of price discrimination, menu cost and own-currency invoicing. Accordingly, we rely on Hansen's

(2000) methodology to implement a two-regime pass-through model featuring a threshold variable. The pass-through equations are based on Knetter's popular specification:

$$p_t = \theta_0 + \theta_{1,1} e_t + \theta_{2,1} c_t + u_t \text{ if } |\Delta e_t| \leq \gamma \tag{13}$$

$$p_t = \theta_0 + \theta_{1,2} e_t + \theta_{2,2} c_t + u_t \text{ if } |\Delta e_t| > \gamma \tag{14}$$

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[7] and Δe_t is the threshold variable that is used to split the sample into two 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. This specification of the threshold implies that revising the export price is (not) profitable in regime 2 (regime 1).[8] The parameters θ_0, $\theta_1' = \begin{bmatrix} \theta_{1,1} & \theta_{2,1} \end{bmatrix}$, $\theta_2' = \begin{bmatrix} \theta_{1,2} & \theta_{2,2} \end{bmatrix}$ and γ need to be estimated. The sample length is denoted by T.

The estimation of the model depicted by (13) and (14) is done by sequential least squares. First, the model is rewritten as a single equation by creating a dummy variable $d_t(\gamma) = \{|\Delta e_t| \leq \gamma\}$ 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 pass-through model reduces to:

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

Let us define $\hat{\theta}$ and $\hat{\delta}_T$ as OLS estimators conditioned on γ. The parameter γ is restricted to a bounded set $\Gamma \equiv [\underline{\gamma}, \overline{\gamma}]$ that is approximated by a grid of observed exchange rate changes 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 $\left\{q_{(1)},\ldots,q_{(T)}\right\}$ 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 $\hat{\gamma}=\arg\min_{\gamma\in\Gamma} S(\gamma)$ among the N candidates.

We consider two distinct scenarios with respect to the choice of currency invoicing. The first scenario presumes that Canadian pork exporters invoice U.S. and Japanese importers 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 induces a variation in the export price such that $-1<\theta_{1,2}<0$ under the joint hypothesis of price discrimination and own-currency invoicing. When menu costs prevent the adjustment of the export price (*i.e.* $|\Delta e_t|<\gamma$), we must have $\theta_{1,1}=0$ under the joint hypothesis in the 1st regime. When pork exports are invoiced in the currency of the importer, the theoretical model shows that the export price in foreign currency does not change in response to a small exchange rate variation. This in turn implies that the export price denominated in Canadian dollars varies proportionally with the exchange rate. Thus we posit that $\theta_{1,1}=-1$ in the first regime under the joint hypothesis of price discrimination and foreign currency invoicing. However, large enough changes in the exchange rate trigger adjustments in the export price and it must be that $-1<\theta_{1,2}<0$ in the second regime.

The null hypothesis of no menu costs is a test of $\theta_1'=\theta_2'$. The asymptotic theory of threshold estimators is complex, but Hansen (2000) derives the asymptotic distribution of the threshold parameter and the slope coefficients under certain conditions. Under the assumption 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 rejecting 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 $\delta_T \to 0$ 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 $\delta_T \to 0$ must be regarded as asymptotically conservative if the error terms are normally distributed. While Hansen (2000) made significant improvements to the asymptotic theory of threshold models, there are still gaps that make statistical inference in our setting particularly challenging. We elected to pursue bootstrap methods to estimate the distributions of the estimators and test statistics. However, the test statistics are not asymptotically pivotal in the sense that their distributions depend on nuisance 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 through pre-pivoting methods introduced by Beran (1988) and summarized in Horowitz (2001).

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 reject this hypothesis about the absence of significant thresholds in the pass-through equation, and keeping in mind that *e* in the empirical section is defined in terms of units of foreign currency per Canadian dollar, we can then test: a) the null hypothesis of no price discrimination $H_0 : \theta_{1,2} = 0$; 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 price discrimination 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

Pass-through 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. 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 price-weighted exchange rates (units of foreign currency per Can$). There is a steady depreciation in the value of the Canadian currency with respect to the U.S. currency over the entire sample. Finally, there are wilder variations in the value of the Canadian dollar with respect to the Japanese yen (weighted by the Japanese food price index).

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:

The ADF test was implemented on the logarithmic transformation of the price-weighted 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 the dependent variable 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 a confidence level of at least 90%.

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 to reinforce our confidence in the stationarity hypothesis.[9] As is well-known, the ADF and KPSS tests often yield conflicting evidence if the critical values of the tests are not adjusted. The results presented in Table 1 provide a vivid example of the notorious inconsistency between ADF and KPSS results. Carrion-i-Silvestre *et al.* (2001) conducted a Confirmatory Data Analysis (CDA) by computing critical values for the joint confirmation hypothesis of a unit root. They argue that their critical values generate more accurate results than standard ADF and KPSS critical values when the data generating process is integrated of order one. The CDA results presented in Table 1 show that most of the variables are stationary. Our analysis proceeds under the assumption that all of the variables are stationary.

As mentioned previously, the distribution of our regression coefficients is non-standard and there is no formal theory about the asymptotic distribution of the coefficients that can be relied upon given that the assumptions outlined by Hansen (2000) are not likely to hold. In a classical regression model, bootstrap methods may achieve better finite sample convergence than asymptotic methods. However, because the statistic $t_j = \hat{\theta}_j \big/ \sigma\left(\hat{\theta}_j\right)$ is not asymptotically pivotal (i.e., its distribution depends on unknown parameters under the null), bootstrap methods may not improve the rate of convergence of the statistic when compared to asymptotic theory. This is why we apply pre-pivoting methods, which can be loosely interpreted as bootstrap iterations. These methods are described in Horowitz (2001) and they entail drawing bootstrap samples from bootstrap samples to create an asymptotically pivotal statistic.

The independent variables in (15) are treated as fixed as well as the threshold variable. The regression vector of residuals $\hat{\mathbf{e}}^*$ is obtained by applying the sequential estimation procedure to (15). It constitutes the empirical distribution that is used for the first bootstrap. A sample of T observations is drawn with replacement from the empirical distribution and a vector for the dependent variable is generated under the null hypothesis being considered. The model is estimated by sequential least squares and a test statistic is computed; *e.g.,* $t_j^* = \hat{\theta}^* \big/ \sigma\left(\hat{\theta}^*\right)$. This procedure is repeated J times. G^* is defined as the statistic about the proportion of times that t_j^* falls below t_j given J. For each of the j^{th} bootstrap regression, a new vector of residuals, $\hat{\mathbf{e}}^{**}$, is also generated from the initial bootstrap regression. It defines the empirical distribution of the second bootstrap and is used to generate a new sample under the null hypothesis. The model is re-estimated using sequential least squares and a new statistic $t_k^{**} = \hat{\theta}^{**} \big/ \sigma\left(\hat{\theta}^{**}\right)$ is computed. This procedure is repeated K times. Thus G^{**} is defined as the statistic that counts the number of times that t_j^{**} falls below t_j^{*} divided by the number of draws in the second bootstrap (*K*). Finally, the *p-value* of the test statistic is: $pvalue = 1 - \#\left\{G^{**} < G^{*}\right\}/J$, where # counts the number of times that the expression inside the parentheses is true. This value is the bootstrap estimate of the asymptotic *p-value* for the *t*-statistic under the null hypothesis. The idea of the double bootstrap procedure is that under the null hypothesis, the statistic G^{**} follows a uniform distribution and is thus exempt of any nuisance parameters.

McCullough and Vinod (1998) suggest that the product of *J* and *K* (i.e., $L \equiv JK$) should be of an order of magnitude at least slightly greater than T^3. Booth and Hall (1994) suggest that the values of *J* and *K* should be set to $\gamma L^{2/3}$ and $\gamma^{-1} L^{1/3}$ respectively; where $\gamma \equiv \left[0.5\left(0.95\right)^{-2} 0.05\left(5/4 - 0.05\right)\right]^{1/3}$ when constructing confidence intervals at the 95% confidence level. A number of practical considerations must also be applied. Due to the discrete nature of the empirical distribution, it is desirable that $\left(J+1\right)/K$ and $K/2$ be integers (McCullough and Vinod, 1998). Finally, our threshold model involves many recursive regressions. Setting *J* and *K* according to the above guidelines implies that 330 millions regressions[10] are required to test a single hypothesis. An obvious drawback of our procedure is that it is time-consuming. For the purpose at hand, we set $J = 2199$ and $K = 440$.

Table 2 presents the OLS estimates of the pass-through equations for pork exports from Quebec and Manitoba to each destination. The coefficient estimate and its standard error, between parentheses, are in the first line of each cell for both regimes. The number underneath is the *p-value* for the null hypothesis of a zero coefficient. The point estimate for the exchange rate in each regime of the pass-through equations has the expected algebraic sign. The first period pass-through coefficients being larger in absolute value than their second-period counterparts are consistent with the menu cost and foreign currency hypotheses. At first glance, foreign currency invoicing seems an especially likely hypothesis for exports to Japan because the first period pass-through coefficients are close to -1. The second period pass-through coefficients being between 0 and -1 are consistent with price discriminating behaviour. The coefficient of the hog price is significant except in the 1st regime of the Quebec-Japan and Manitoba-Japan pass-through equations. Statistically significant coefficients have the expected algebraic sign as an increase in the processors' marginal cost is expected to induce an increase the export price. Moreover, the first regime marginal cost coefficients in the U.S. export equations are lower than their second regime counterparts. As argued earlier, this is consistent with the exchange rate being partly correlated with processors' marginal cost.

For comparison purposes, Table 3 presents the coefficient estimates and standard errors for linear pass-through equations. The reported exchange rate coefficients are at least twice as large as their standard error which is consistent with the notion that pork exporters price discriminate. The pass-through coefficients in Table 3 ought to be compared to the second regime coefficients in Table 2 and we find that they are quite similar in magnitude, except for the pass-through coefficient for the Manitoba-Japan equation since the coefficient reported in Table 3 better reflect the coefficient for regime 1 in Table 2.

It should be noted that the threshold values reported in Table 2 do not provide a direct estimate of menu costs because the threshold variable is actually conditioned by menu costs *and* by the structural parameters of demand in the importing country. Hence, a large threshold estimate does not necessarily imply large menu costs, but large menu costs make the threshold parameter larger. As anticipated, the threshold estimates are always larger for the Japanese market than for the U.S. market. The United States and Canada share a common border, common language and similar institutions and as such one would expect that menu costs should be less important in transactions involving the United States. Our results support this argument.

Our overview of the regression and threshold coefficients provide some insights, but it is no substitute for rigorous hypothesis testing. We begin with the null of no threshold which 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. As mentioned earlier, the *p-value* is computed by simulating a sample of *T* observations under the null hypothesis and by computing the proportion of occurrences for which the bootstrap statistic falls below the actual *LR* statistic. This procedure is repeated using the bootstrap sample as the basis for the empirical distribution of another bootstrap simulation. Once again, the proportion of times that the 2nd bootstrap statistic falls below the initial bootstrap statistic is computed. The final *p-value* is obtained by comparing the two statistics defined over the $[0,1]$ interval.

The *LR* statistic rejects the null hypothesis of no threshold in three of the four equations.[11] The non-rejection of the null of no menu costs for the Quebec-Japan equation can be attributed to the relatively large standard error associated with the first regime pass-through coefficient. We conjecture that the non-rejection finding has more to do with the relatively short length of our sample than with the actual size of menu costs faced by exporters doing business with a country as geographically remote as Japan. This is why we proceeded with tests about price discrimination and currency invoicing, taking as valid the specification of our two-regime pass-through equations.

The null hypothesis of no price discrimination is a test about the statistical significance of the coefficient of the exchange rate in the 2nd regime of the pass-through equation $\left(H_0:\theta_{1,2}=0\right)$.

The null hypothesis of no-price discriminating behaviour is rejected at conventional significance levels in all cases except for Manitoba exports to Japan. As such, Quebec and Manitoba pork exporters appear to exercise some market power in the U.S. market. Historically, Quebec has been the dominant exporter of pork products among Canadian provinces while Manitoba has been known to produce and export large volumes of live hogs. Furthermore, Quebec processors have been catering to the wants and needs of Japanese buyers long before the Canada-U.S. hog/pork disputes in the mid 1980s.[12] It could be that their investment in a solid business relationship allows them to price discriminate.

The null hypothesis of domestic (or own) currency invoicing is a test about the significance of the coefficient of the exchange rate in the first regime $\left(H_0 : \theta_{1,1} = 0\right)$. This hypothesis is rejected in all equations at the 90 percent confidence level. The foreign currency invoicing hypothesis is a test of: $H_0 : \theta_{1,1} = -1$. The bootstrap simulations yield strong evidence of foreign currency invoicing for Quebec and Manitoba pork exports to Japan (respective *p-values* are 0.725 and 0.384). However, the statistical evidence rejects both the domestic and foreign currency invoicing hypotheses for exports from the two Canadian provinces to the U.S. We dismiss the possibility that Canadian products exported to the U.S be invoiced in a third country currency. Instead, it is hypothesized that the inconclusive "currency invoicing" results for the U.S. destination are attributable to a split among Canadian firms as to their choice of currencies (Can$ or US$) when invoicing.[13] To explore the consequences of this hypothesis on the parameters of our econometric model, let us assume that there are two types of firms that sell a homogenous pork product on the U.S. market. They are identical in all respects, except in their choice of currency used for invoicing. What we wish to ascertain is whether sequential least square estimates obtained from a data generating process averaging over the two types of firms could generate pass-through coefficients similar the ones reported in Table 2 ($-1<\theta_{1,1}<\theta_{2,1}<0$).

To thoroughly investigate the implications of this hypothesis, we build a small simulation model based on the Quebec to U.S. exports. We use the actual price of live hogs and exchange rate as independent variables. We set the true threshold parameter equal to the estimate in table 2, $\hat{\gamma}=0.001345$ and set the pass-through parameter of the first-regime to differ across firms' type (denoted *a* and *b*) depending on whether they invoice their sales in U.S. or Canadian dollars. For type *a* firms which invoice their export sales in Can$, we set the true parameters of the pass-through equation according to: $\theta^a \equiv \left[\theta_0^a \quad \theta_{1,1}^a \quad \theta_{1,2}^a \quad \theta_{2,1}^a \quad \theta_{2,2}^a\right]=\left[0.7 \quad 0 \quad 0.4 \quad -0.2 \quad 0.4\right]$. The behaviour of type *b* firms is driven by: $\theta^b=\left[0.7 \quad -1 \quad 0.4 \quad -0.2 \quad 0.4\right]$. Hence, the two types only differ in their currency invoicing decision as they exercise the same degree of price discrimination. The pricing decisions of each firm are generated by the equations:

$$p_t^j=\theta_0+\theta_{1,1}e_t+\theta_{2,1}c_t+v_t^j \text{ if } \left|\Delta e_t\right|\le\gamma\,; \quad p_t^j=\theta_0+\theta_{1,2}e_t+\theta_{2,2}c_t+v_t^j \text{ if } \left|\Delta e_t\right|>\gamma\,;\ j=a,b$$

where v^a and v^a are random error terms drawn from a normal distribution with mean zero and standard error 0.077. It is assumed that the relative importance of each type of firm in the industry is the same. Because firms within a type have identical sales, the reported export price at the aggregate level must be equal to a weighted sum of p^a and p^b. For this data generating process characterized by an even split between firms that invoice in Can$ and firms that invoice in US$, the sequential least squares procedure yields the following pass-through coefficients at the aggregate level: $\hat{\theta}_{1,1}^{a,b}=-0.47$ and $\hat{\theta}_{1,2}^{a,b}=-0.16$. These coefficients are surprisingly close to the Quebec-U.S. first-regime and second-regime pass-through coefficients reported in Table 2 (-

0.38, -0.21). Even though our simulation does not constitute a formal hypothesis test, it provides a reasonable explanation of the pass-through coefficients for exports to the U.S. in Table 2.

The individual hypotheses about the absence of price discrimination and currency invoicing can also be tested jointly. For example, $H_0 : \theta_{1,1} = \theta_{1,2} = 0$ is a joint hypothesis about the absence of price discrimination and domestic currency invoicing. 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 $\mathbf{R} = \begin{bmatrix} 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \end{bmatrix}$ and $\mathbf{r} = \begin{bmatrix} 0 & 0 \end{bmatrix}'$. The test statistic is $F = \left(\mathbf{R}\hat{\Theta} - \mathbf{r}\right)' \left[\mathbf{R}\,\mathrm{cov}\left(\hat{\Theta}\right)\mathbf{R}'\right]^{-1} \left(\mathbf{R}\hat{\Theta} - \mathbf{r}\right) \Big/ 2$. The inference is made possible by using bootstrap samples as described previously. The joint hypothesis about the absence of price discrimination and domestic currency invoicing is 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 hypothesis about the absence of price discrimination and foreign currency invoicing. This joint hypothesis is squarely rejected in cases involving exports to the U.S. It is also rejected, but less strongly, in the Quebec-Japan pass-through equation. In the case of pork exports from Manitoba to Japan, the evidence suggests that Manitoba pork exporters are unable to price discriminate.

5. Conclusion

This paper developed a theoretical exchange rate pass-through framework accounting for menu costs and different choices of currency for invoicing purposes. Menu costs make it costly for exporters to revise their prices in response to exchange rate changes. This introduces a non-linearity between the exchange rate and the export price. This non-linearity motivates the empirical specification of a two-regime pass-through model to analyze the pricing decisions of pork exporters from two Canadian provinces to the U.S. and Japan. The choice of currency used for invoicing purposes imposes theoretical restrictions on the pass-through in the first regime (i.e., when menu costs are high relative to the profits arising from a price change) which can be tested empirically.

The empirical model rejects the null hypothesis of no menu costs in three of the four equations. Statistically significant menu costs are identified in the export pricing decisions of Quebec and Manitoba exporters in their dealings with U.S. buyers. Manitoba pork exporting firms also appear to face menu costs in their dealings with Japanese buyers. We argue that the non-rejection in the case of the Quebec-Japan pass-through equation is more likely attributable to the small length of our sample than to the actual significance of menu costs faced by Quebec firms. Overall, the empirical evidence favours threshold pass-through models over linear ones.

The evidence of price discriminating behaviour is weak for Manitoba pork exports to Japan. The individual null hypothesis about no price discrimination could no be rejected and the same can be said about the joint hypothesis about the absence of price discrimination and foreign currency invoicing. However, foreign currency invoicing and price discrimination appear to characterize the behaviour of Quebec exporters in their dealings with Japanese importers. We also found evidence of price discrimination in Canadian exports to the United States, but it was not possible to validate the hypothesis that export sales are invoiced in U.S. dollars. A simulation showed that the pass-through coefficients for exports shipped to the U.S. are consistent with the concurrent use of the Canadian and U.S. currencies by Canadian exporters. Donnenfeld and Haug (2003) argue that even when firms export similar products they are likely to choose different currencies for invoicing purposes because of heterogeneity in the firms' cost structures and attitudes toward risk. However, the across-destinations differences in our results suggest that there are also destination-specific factors that influence the choice of currencies beside heterogeneity among exporting firms. Overall, our results are consistent with the stylized facts of world pork trade, Canada being a major player on the world scene, and with the differences in the domestic market structures of the Quebec and Manitoba processing sectors.

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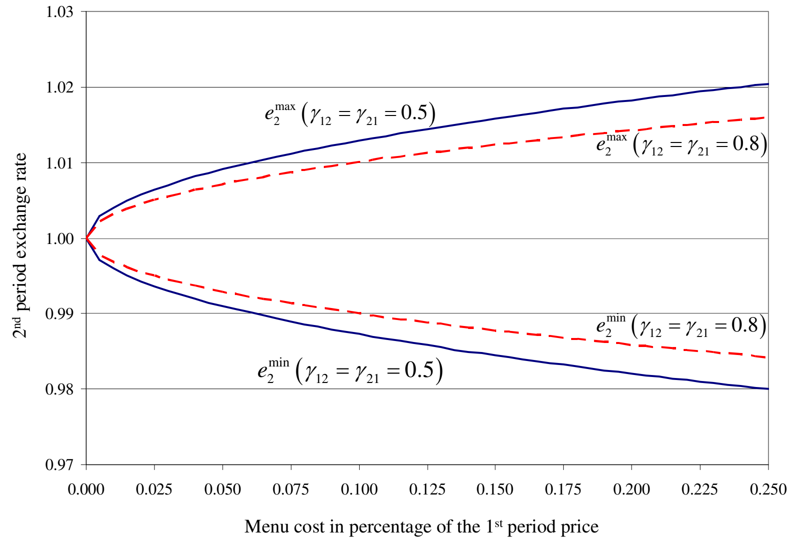

Figure 1. Simulated exchange rate band as a function of a fixed menu cost *m*.

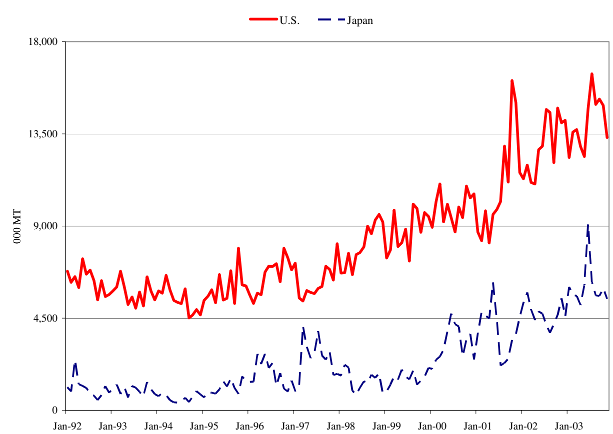

Figure 2a. Pork meat exports from Quebec to the U.S. and Japan from January 1992 to December 2003

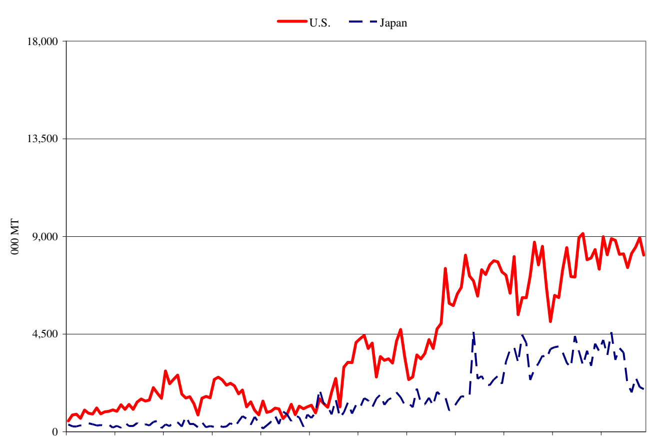

Figure 2b. Pork meat exports from Manitoba to the U.S. and Japan from January 1992 to December 2003

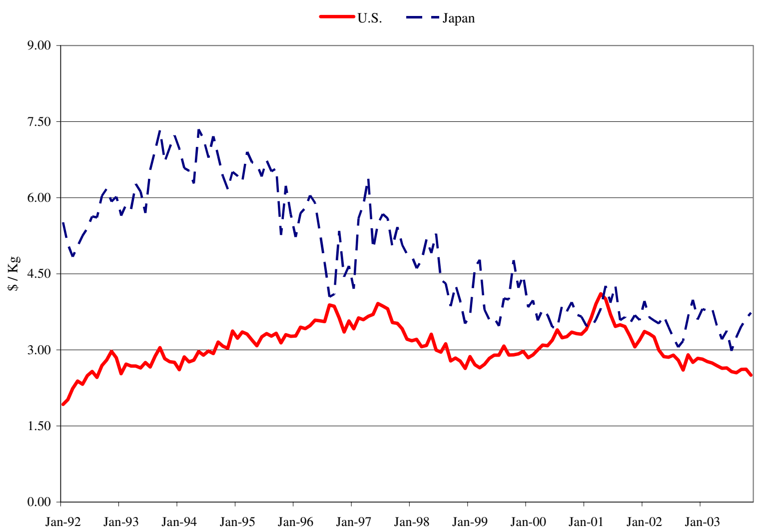

Figure 3a. Quebec export unit values to the U.S. and Japan from January 1992 to December 2003

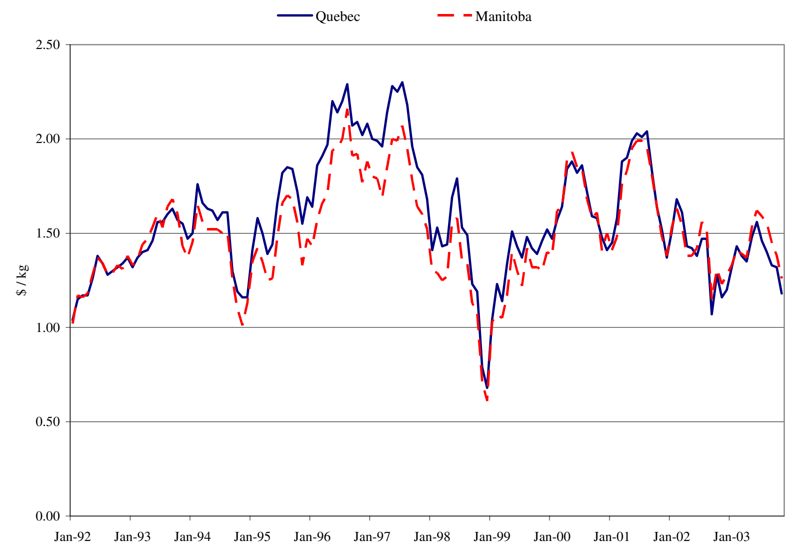

Figure 4. Hog prices in Quebec and Manitoba from January 1992 to December 2003

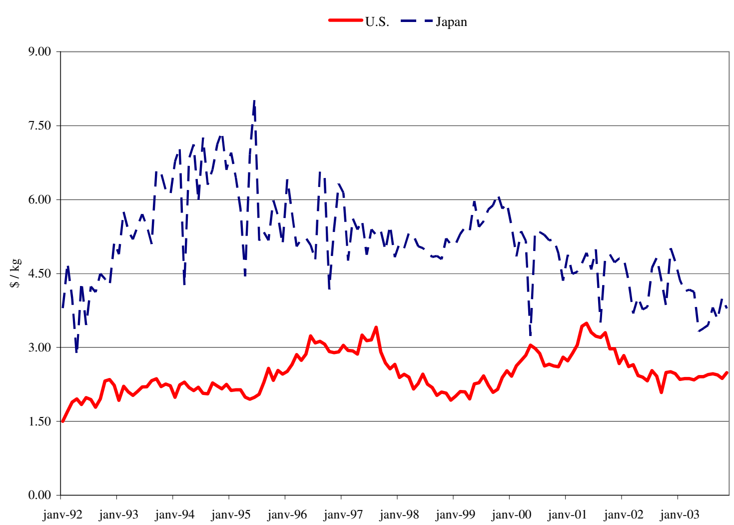

Figure 3b. Manitoba export unit values to the U.S. and Japan from January 1992 to December 2003

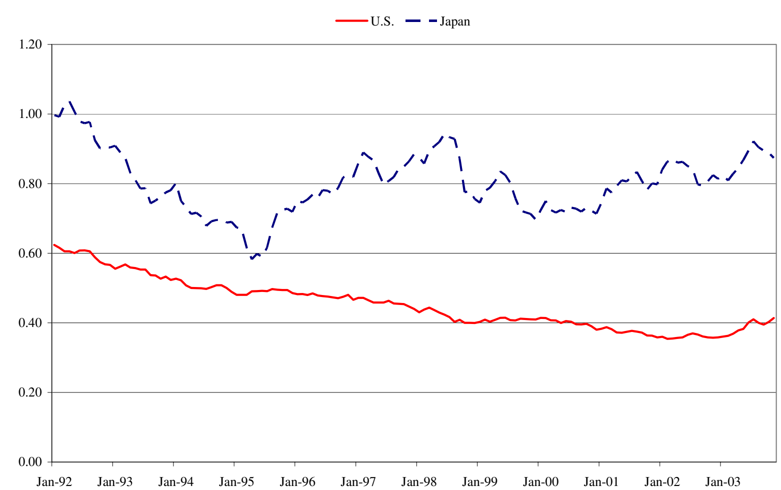

Figure 5. Value of the foreign currency per Can$ weighted by the consumer food price index

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
Manitoba				
Export price to the U.S. (NT)	0	-3.45**	0.84*	No
Export price to Japan (T)	2	-3.13**	0.34*	No
Hog price (NT)	0	-3.45*	0.13	No
U.S. weighted exchange rate (NT)	0	-2.70**	2.78*	Yes
Japan weighted 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 from Carrion-i-Silvestre *et al.* (2001).

Table 2. Estimates of the Pass-Through equation and inference

	Quebec		Manitoba	
Parameters	**U.S.**	**Japan**	**U.S.**	**Japan**
constant	0.70 (0.04) 0.000	1.35 (0.06) 0.000	0.28 (0.04) 0.000	1.40 (0.04) 0.000
1st regime				
weighted x-rate, $\theta_{1,1}$	-0.38 (0.06) 0.000	-0.81 (0.37) 0.099	-0.60 (0.06) 0.000	-1.15 (0.12) 0.000
hog price, $\theta_{2,1}$	0.34 (0.06) 0.000	0.25 (0.21) 0.405	0.41 (0.06) 0.000	-0.07 (0.07) 0.453
2nd regime				
weighted x-rate, $\theta_{1,2}$	-0.21 (0.04) 0.000	-0.52 (0.19) 0.000	-0.40 (0.05) 0.000	-0.20 (0.12) 0.232
hog price, $\theta_{2,2}$	0.55 (0.04) 0.000	0.18 (0.11) 0.079	0.71 (0.04) 0.000	0.50 (0.07) 0.000
Threshold estimate	0.001	0.005	0.001	0.043
Hypothesis testing	**Statistic** ***p-value***	**Statistic** ***p-value***	**Statistic** ***p-value***	**Statistic** ***p-value***
Foreign currency invoicing $\left(\theta_{1,1} = -1\right)$	10.33 0.000	0.35 0.725	6.67 0.000	-1.25 0.384
Likelihood ratio test of no menu costs $\left(\theta_{1,1} = \theta_{1,2}; \theta_{2,1} = \theta_{2,2}\right)$	15.99 0.029	2.54 0.940	23.95 0.001	14.97 0.036
Joint no PTM and domestic currency invoicing $\left(\theta_{1,1} = 0; \theta_{2,1} = 0\right)$	21.65 0.000	4.80 0.032	58.82 0.000	43.03 0.000
Joint no PTM and foreign currency invoicing $\left(\theta_{1,1} = -1; \theta_{2,1} = 0\right)$	211.75 0.000	4.05 0.094	244.94 0.000	0.95 0.666

Table 3. Linear PTM equations

Parameters	Quebec U.S.	Quebec Japan	Manitoba U.S.	Manitoba Japan
constant	0.73	1.36	0.31	1.41
	(0.04)	(0.06)	(0.04)	(0.04)
weighted x-rate	-0.23	-0.55	-0.43	-0.97
	(0.05)	(0.19)	(0.05)	(0.12)
hog price	0.48	0.16	0.59	-0.03
	(0.04)	(0.10)	(0.04)	(0.07)

Endnotes

[1] Until recently, Quebec was the largest hog producing province in Canada. Quebec remains the largest exporting province of pork products as every hog produced is processed within the province. However, differences in the evolution of provincial environmental regulations have given Manitoba a significant comparative advantage in hog production. Pork processing is a highly concentrated sector. The largest firm in Quebec was processing over 65% of all the hogs produced in that province in the early 1990s. While its share has decreased during throughout the 1990s and early 2000s, it recently jumped due to a merger with the second largest firm. The processing sector in Manitoba is less concentrated and as such is likely to generate contrasting results.

[2] See the comprehensive review of Wohlgenant (2001) for empirical evidence.

[3] This would be the case for instance when translating and legal services must be contracted to implement a price change. Furthermore, transactions involving Canadian food processors and U.S. food distributors and retailers are often implemented through intermediaries or middlemen. Therefore, price changes may be costly to communicate, especially if they cause interruptions in deliveries, even when all parties involved speak the same language and share a similar business culture.

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

[5] More generally, $\frac{dp_{11,t}}{de_t} = \frac{1}{|G|}\left[\frac{\partial q_{11,t}}{\partial p_{11,t}}\delta\right]\left[2c_{1Q\omega}\frac{\partial \omega_{1,t}}{\partial e_t}e_t - \frac{c_{1QQ}c_{1Q}\delta}{e_t}\right]$, where $|G| > 0$ from the second order condition. Under our technology assumptions, the last bracketed expressions vanish and markets can be analyzed separately, as is routinely done in empirical pass-through and PTM studies.

[6] As before let $\varepsilon_{1j} \equiv \left(\partial p_{1j,t}/\partial e_t\right)\left(e_t/p_{1j,t}\right)$, then $\partial\left(p_{11,t}/ep_{12,t}\right)/\partial e_t < 0$ implies $\varepsilon_{11} - \left(1+\varepsilon_{12}\right) < 0$.

[7] The price of live hogs seems a natural proxy because live animals represent a large share of processors' costs. Hence, marginal cost is not likely to be constant and could influence the pass-through outcome.

[8] The specification in (13) and (14) admits two different relationships between marginal costs and the export price conditioned on the first difference in the exchange rate, and thus menu costs. This specification works best in fitting the data and was preferred to a Knetter-like specification that excludes a specific proxy for marginal cost. We posit that marginal costs should have a lesser (larger) impact in the first (second) regime if the exporters' marginal cost is influenced by exchange rate variations, which is clearly a possibility in the case of the Canada/U.S. exchange rate. Given that price adjustments are justified under large exchange rate variations, one would expect a larger marginal cost coefficient under these circumstances. In a more general framework with multiple thresholds, changes in marginal costs arising from increases in labour costs or hog prices should not necessarily induce a change in the export price because of menu costs. Thus one might be tempted to estimate a multiple-thresholds model, but the difficulty in conducting statistical inference makes this option unattractive.

[9] 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: $y_t = \delta t + \zeta_t + \varepsilon_t$; $\zeta_t = \zeta_{t-1} + u_t$; $u_t \sim iid\left(0, \sigma_u^2\right)$. The null hypothesis of trend stationarity is about the validity of a zero restriction on σ_u^2.

[10] Given that the sample length is 143, the product of K and J must equal $143^3 = 2,924,207$. The latter number is the number of total bootstrap samples that are generated. The sequential least squares procedure requires leaving out the lower and upper ten percentile of the observations ordered according to the threshold value. Hence, there are 115 OLS regressions computed for each bootstrap sample leading to a total of $115 \times 143^3 = 336,283,805$ regressions.

[11] It is also interesting to compare the bootstrap critical values with the asymptotic critical values reported in Hansen (2000). Differences are expected because of our relatively small sample and because Hansen's asymptotic critical

values were calculated under the assumption that the threshold vanishes as sample size grows. The likelihood ratio of 15.99 for the no-menu cost hypothesis for the Quebec-US equation has a *p*-value of 2.9% while Hansen's asymptotic critical value at the 97.5 percent confidence level is 8.75 (Hansen, 2000, p. 582). The rather large difference between these critical values associated with very similar confidence levels demonstrates the importance of computing finite sample critical values.

[12] Quebec industry representatives mentioned that Japanese buyers had demands that were costly to meet and that transactions involved small volumes at first. As Quebec processors proved themselves reliable, the volume and value of the transactions increased and so did the margins, as in reputation models.

[13] Donnenfeld and Haug (2003) report that it is not unusual for exporting firms selling a homogenous product to use different currencies for invoicing purposes. Differences in cost structures and/or in attitudes toward risk can explain differences in currency invoicing decisions.