A joint test of market power, menu costs, and currency invoicing

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Abstract

This article investigates exchange rate pass-through (ERPT) 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 have some market power 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.

JEL classification: C22, F12, F14

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1. Introduction

Since the pioneering work of Krugman (1987) on the concept of pricing-to-market (PTM), a considerable literature documenting evidence of price discrimination and incomplete exchange rate pass-through (ERPT) in international markets has emerged. The idea is that if export prices do not fully adjust to changes in the exchange rate (ERPT is incomplete), there must exist market power because mark-ups of price over marginal cost are not constant. The PTM concept refers to changes in the ratio of domestic to export prices following exchange rate variations and thus refers to price discrimination. 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. At the manufacturing level, Sasai (2002) and Uctum (2003) have analyzed the price discriminating behavior 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 behavior has also been observed in agri-food sectors. 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.\textsuperscript{1}

Even though trading firms are confronted by the currency invoicing issue every day as they negotiate prices in their own

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Data Appendix Available Online

A data appendix to replicate main results is available in the online version of this article. Please note: Wiley-Blackwell, Inc. is not responsible for the content or functionality of any supporting information supplied by the authors. Any queries (other than missing material) should be directed to the corresponding author for the article.

\textsuperscript{1}A voluminous literature addresses the issue of transmission of monetary and fiscal policies in the context of open economies and show that pricing to market behavior has important welfare implications (e.g., Betts and Devereux, 2000; Devereux et al., 2005; Dixit, 1989; Tille, 2001). Another strand of the literature to which the present paper is closely related aims at providing explanations for incomplete pass-through in import/export prices (e.g., Campa and Goldberg, 2005; Feenstra et al., 1996; and Froot and Klemperer, 1989).
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 literature review of Bowen et al. (1998), most studies attempt to rationalize the stylized fact that most international transactions are invoiced in the exporters’ currency by analyzing the role of currency invoicing as a hedging strategy against exchange rate risk (Donnenfeld and Haug, 2003; Friberg, 1998; Johnson and Pick, 1997). Others compared the role of macroeconomic factors and industry-specific features in explaining firms’ currency invoicing decisions (Bacchetta and van Wincoop, 2005; Goldberg and Tille, 2005). In these models, imperfect competition is intrinsically linked to currency invoicing, yet few authors have analyzed the two issues from an empirical standpoint. One notable exception is Sato (2003) who develops 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 ERPT decision 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.

Another important feature of export pricing decisions is that firms face costs in adjusting prices in response to changes in market conditions (such as the exchange rate). Flodén and Wilander (2006) propose a general theory of ERPT under the assumption that firms incur menu costs when they make changes to their pricing strategies. 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. The predictions of their theoretical model are consistent with the intuition that large enough menu costs discourage price adjustments that would otherwise be observed.

The objective of this article is to assess the plausibility of individual and joint hypotheses regarding the significance of menu costs, incomplete ERPT and the choice of currency for invoicing purposes. To shed some light on these issues, we propose in the next section a simplified version of Flodén and Wilander (2006)’s state-dependent pricing 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. Pork processing is a highly concentrated business that has been the object of many market power studies (e.g., Lopez et al., 2002; Morrison Paul, 1999). Accordingly, the extent to which large pork processors are able to exercise market power 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 U.S. and Japan are also the largest importers of Canadian pork.

The empirical model is based on Knetter (1989)’s insight to measure ERPT. It introduces a two-regime threshold model that is consistent with the existence of menu costs. The ERPT coefficients are conditional on the magnitude of the change in the exchange rate to capture the impact of menu costs. The results suggest Canadian pork exporting firms can exercise some market power in the Japanese and U.S. pork markets. The null hypothesis of no menu costs is soundly rejected by the data in three of the four ERPT 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 complete pass-through in two out of the four ERPT equations.

The next section lays out the theoretical model that rationalizes price rigidity when exporters are facing menu costs. 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 on domestic and export prices. The theoretical findings are then exploited in the following section to measure ERPT in export prices and test various hypotheses about 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

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

(1)

where $p_{i,j,t}$ and $q_{i,j,t}$ 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 $c_1(\cdot)$ 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

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

(2)
It is assumed that \( c_{1Q} \equiv \frac{\partial c_i(\cdot)}{\partial Q_i} > 0 \) where \( Q_i \) is the total quantity produced by firm \( i \). It is also assumed that marginal cost is constant, i.e., \( c_{1QQ} \equiv \frac{\partial^2 c_i(\cdot)}{\partial Q_i^2} = 0 \).

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 (due to the Bertrand conjectures) on its own turf by announcing its price last in our game. It also seems natural to assume that marginal culture can put an exporter at a disadvantage \( \text{vis-à-vis} \) 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, firm 1’s first-order conditions for firm 1’s profit maximization are

\[
\frac{\partial \pi_{1,t}}{\partial p_{11,t}} = q_{11,t} + (p_{11,t} - c_1 Q_1) \frac{\partial q_{11,t}}{\partial p_{11,t}} = 0 \quad (3)
\]

\[
\frac{\partial \pi_{1,t}}{\partial p_{12,t}} = q_{12,t} + \left( \frac{p_{12,t} - c_1 Q_1}{e_t} \right) \times \left( \frac{\partial q_{12,t}}{\partial p_{12,t}} + \frac{\partial q_{12,t}}{\partial p_{22,t}} \frac{\partial p_{22,t}}{\partial p_{12,t}/e_t} \right) = 0. \quad (4)
\]

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: \( p_{11,t}(1 + 1/e_{11,t}) = c_1 Q_1 \), where \( e_{11,t} \) is the price elasticity of demand faced by 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 asymmetry in its exchange rate. 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 \( \frac{dp_{11,t}}{de_t} = 0 \) 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 pass-through literature.

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

\[
\frac{dp_{12,t}}{de_t} = \frac{1}{|H|} \left( \frac{\partial q_{11,t}}{\partial p_{11,t}/e_t} \delta \right) (2p_{12,t} - c_1 Q_1) > 0 \quad (5)
\]

where \( |H| > 0 \) from the second order condition. Furthermore, given that \( p_{12,t} - c_1 Q_1 > 0 \), 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 et al. (1998). It is also possible to show that firm 1’s export price expressed in country 2’s currency actually follows as country 1’s currency depreciates (i.e., \( \partial (p_{12,t}/e_t)/\partial e_t < 0 \)), an outcome usually referred to as incomplete pass-through.

Let us now assume that when firm 1 wants to change \( p_{12,t} \), it must incur a fixed menu cost \( m \). In the second 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

\[
\pi_{12}(p_{12,1}; e_t) \geq \pi_{12}(p_{12,1}; e_t) - m. \quad (6)
\]

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: \( e_{12}^{\text{min}}(p_{12,1}; m) \) and \( e_{12}^{\text{max}}(p_{12,1}; m) \). The existence of these boundaries follows from the concavity of profit with respect to price and the spread between them is increasing with the menu cost.

Formally, the optimization problem of firm 1 in period 1 is the following. Firm 1 knows that it will keep its period 1 price in period 2 as long as \( e_2 \in [e_{12}^{\text{min}}, e_{12}^{\text{max}}] \). We assume that the firms’ period 1 expectation of the exchange rate in period 2 is \( E_1[e_2] \). For simplicity, let us assume that the exchange rate is drawn from a uniform distribution with support \([e, \bar{e}]\), a mean of \((\bar{e} - e)/2\), and that the parameter values are such that \( e < e_{12}^{\text{min}} < e_{12}^{\text{max}} < \bar{e} \). Hence, there is a probability \( \text{prob}(e_{12}^{\text{min}} < e_2 < e_{12}^{\text{max}}) = (e_{12}^{\text{max}} - e_{12}^{\text{min}})/(\bar{e} - e) \). If firm 1 will keep its period 1 price in period 2. Under such circumstances, it expects to earn \( E[\pi_{12}] \) as opposed to \( E[\pi_{12}'(\cdot)] \) which changes its price. Therefore, firm 1’s optimization in period 1, given discounting parameter \( \phi \), is as follows:

\[
\max \pi_{12}(p_{12,1}; e_1) + \phi \text{prob}(e_2^{\text{min}} \leq e_2 \leq e_2^{\text{max}}; \bar{e}, \phi, m) E[\pi_{12}(p_{12,1}, p_{12,2})]
\]

\[+ \phi (1 - \text{prob}(e_2^{\text{min}} \leq e_2 \leq e_2^{\text{max}}; \bar{e}, \phi, m)) \times E[\pi_{12}(p_{12,1}, p_{12,2}, m)]. \quad (7)
\]

The first order conditions are

\[
\frac{\partial \pi_{12}}{\partial p_{11,1}} = 0; \quad \frac{\partial \pi_{12}}{\partial p_{12,1}} + \phi \text{prob}(\cdot) \frac{\partial E[\pi_{12}]}{\partial p_{12,1}}
\]

\[+ \phi \text{prob}(\cdot) (E[\pi_{12}] - E[\pi_{12}']) = 0. \quad (8)
\]

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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.
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 to which firm 1’s profit in the second period must be taken into account in its first 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., $e_2 \notin [e_{2, \min}, e_{2, \max}]$), it would simply set $\partial \pi_{1,t}/\partial p_{12,t} = 0$ in choosing $p_{12,t}$.

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., $p_{12,t} = p_{12,2}$) 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 $p_{12,t}/e_2$ falls. As a result, we should not observe an incomplete ERPT in spite of our uncompétitive 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 econometrics is 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, assume the export price of firm 1 at time $t(p_{12,t})$ is now denominated in country 2’s currency while the currency denomination of other prices is left unchanged. The profit of firm 1 is

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

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

$$\frac{\partial \pi_{1,t}}{\partial p_{11,t}} = q_{11,t} + (p_{11,t} - c_1 e_t) \frac{\partial q_{11,t}}{\partial p_{11,t}} = 0 \tag{10}$$

$$\frac{\partial \pi_{1,t}}{\partial p_{12,t}} = e q_{12,t} + (e p_{12,t} - c_1 e_t) \left( \frac{\partial q_{12,t}}{\partial p_{12,t}} + \frac{\partial q_{12,t}}{\partial p_{p22,t}} \frac{\partial p_{22,t}}{\partial p_{12,t}} \right) = 0. \tag{11}$$

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). For the export price denominated in country 2’s currency, we find

$$\frac{dp_{12,t}}{de_t} = \frac{1}{|G|} \left[ \frac{\partial q_{11,t}}{\partial p_{11,t}} \right] \left[ -\frac{2c_1 e_t}{e_t^2} \right]. \tag{12}$$

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., $\partial ep_{12}/\partial e > 0$). 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 be quoted in a third-country’s currency. For example, Canadian exports to Japan could be invoiced in U.S. dollars. In this case, a large appreciation of the yen relative to the U.S. dollar would induce an increase in the U.S. dollar price chosen by the Canadian firm. If the yen appreciated relative to the Canadian dollar by the same percentage, the Canada–U.S. exchange rate would remain the same and the export price expressed in Canadian dollars would rise. As such, the adjustment is not unlike those described for the two other cases. However, a less pronounced appreciation of the yen relative to the Canadian dollar would cause an appreciation of the Canadian dollar relative to the U.S. dollar that would mitigate and perhaps even reverse the increase in the export price expressed in Canadian dollars. When shocks are small (or when menu costs matter), the export price, expressed in Canadian dollar, may remain the same if the Canada–U.S. exchange rate does not change or it may
increase (decrease) if the Canadian dollar appreciates (depreciates) relative to its U.S. counterpart. In short, the “third-country” currency case is much more complex because the dynamics between the three currencies must also be modeled.

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 market power, 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 ERPT equations are based on Knetter (1989)’s popular specification:

\[ p_i = \theta_0 + \theta_{1,1}e_i + \theta_{2,1}c_i + u_t, \quad \text{if } |\Delta e_i| \leq \gamma \]  \hspace{1cm} (13)

\[ p_i = \theta_0 + \theta_{1,2}e_i + \theta_{2,2}c_i + u_t, \quad \text{if } |\Delta e_i| > \gamma \]  \hspace{1cm} (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 \( \Delta e_i \) 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).\(^5\) The parameters \( \theta_0, \theta_1 = [\theta_{1,1} \theta_{2,1}], \theta_2 = [\theta_{1,2} \theta_{2,2}] \), and \( \gamma \) 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 \( d_i(\gamma) = [|\Delta e_i| \leq \gamma] \) such that \( X_i(\gamma) = X_i d_i(\gamma) \); where \( X_i \) is the vector of independent variables in (13) and (14). The pass-through model reduces to

\[ p_i = \theta' X_i + \delta_T X_i(\gamma) + u_t. \]  \hspace{1cm} (15)

Let us define \( \hat{\theta} \) and \( \hat{\delta} \) as OLS estimators conditional on \( \gamma \). The parameter \( \gamma \) is restricted to a bounded set \( \Gamma \equiv \{\gamma, \tilde{\gamma}\} \) that is approximated by a grid of observed exchange rate changes defined by: \( \Gamma \cap \{\Delta e(1), \ldots, \Delta e(T)\} \). The estimation procedure requires \( N < T \) evaluations of Eq. (15); where \( N \) is selected such that the 10% upper and lower percentiles of \( \{\Delta e(1), \ldots, \Delta e(T)\} \) are not included in \( \Gamma \). A natural estimator for \( \gamma \) is to minimize the sum of squared errors \( S(\gamma) = \sum_{t=1}^{T} \hat{u}_t^2(\gamma) \) such that \( \hat{\gamma} = \arg \min_{\gamma \in \Gamma} S(\gamma) \) among the \( N \) candidates.

There are two different hypotheses to test with respect to the choice of currency invoicing. The first hypothesis is 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., \( |\Delta e_i| \geq \gamma \)), 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 market power and own-currency invoicing. When menu costs prevent the adjustment of the export price (i.e., \( |\Delta e_i| < \gamma \)), we must have \( \theta_{1,1} = 0 \) under the joint hypothesis in the first 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 market power 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. Still, 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 \( \delta_T \equiv \theta_{1,1} - \theta_{1,2} \), 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 nonstandard distribution can be conveniently computed in closed-form. On the other hand, there is no reason to believe in our context that menu costs will disappear as the sample size increases. If \( \delta_T \) 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.

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\(^4\)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.

\(^5\)The specification in Feenstra et al. (1996) and Flodén and Wilander (2006) 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 present case. 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 labor 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.
While Hansen (2000) made significant contributions to the asymptotic theory of threshold models, there are still gaps that make statistical inference in our setting particularly challenging. We elected to rely on bootstrap methods to estimate the distributions of the estimators and test statistics. However, the test statistics are not asymptotically pivotal because 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 discussed 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 ERPT equation (keeping in mind that \( \epsilon \) 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 market power \( H_0 : \theta_1,2 = 0 \); (b) the null hypothesis of invoicing in Canadian (importing country) currency or \( H_0 : \theta_1,1 = 0 \) \( (H_0 : \theta_1,1 = -1) \); and (c) the joint null hypothesis of no market power and invoicing in Canadian (importing country) currency \( H_0 : \theta_1,1 = \theta_1,2 = 0 \) \( (H_0 : \theta_1,2 = 0 ; \theta_1,1 = -1) \).

4. Data and estimation

ERPT equations are specified for pork meat exports from two different Canadian provinces to two destinations over the period beginning in January 1992 and ending in December 2003. Export data were obtained from Statistics 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 of pork processors 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 meat exporters. Pork exports of each province are depicted in Fig. 1(a) and (b). Historically, Quebec has been the dominant exporter of pork products among Canadian provinces while Manitoba firms produce and export large volumes of both ready-to-slaughter hogs and feeder pigs. Despite Manitoba’s significant exports of live hogs, the current study focuses on the ERPT associated with pork meat exports.\(^6\) Fig. 2(a) and (b) present the unit export values by province 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.

Fig. 3 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 differences in hog marketing institutions. In short, it is mandatory for Quebec hog producers to sell their hogs through the producers’ controlled marketing board while Manitoba producers are free to deliver hogs to the buyer of their choice. Transactions are either made on the spot market or through private contracts with packers (Larue et al., 2000). Fig. 4 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 larger variations in the value of the Canadian dollar with respect to the Japanese yen (weighted by the Japanese food price index). These observations are summarized in Table 1, which presents summary statistics for the variables of the ERPT equations. There are significant differences in the distribution of export prices across origin and destinations indicating potential differences in the basket of pork products exported by firms in each province. The variability in the weighted U.S. dollar exchange rate is larger than in the weighted Japanese yen.

As 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 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 2. The first column indicates whether a time trend (T) or no time trend (NT) was used. Following Booth and 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 favor of stationarity, the stationarity test developed by Kwiatkowski et al. (KPSS, 1992) was also carried out for each series to reinforce our confidence in the stationarity hypothesis. The KPSS test involves estimating the equation: \( y_t = \delta t + \xi_t + e_t; \xi_t = \xi_{t-1} + u_t; u_t \sim iid(0, \sigma^2_u) \). The null hypothesis of trend stationarity is about the validity of a zero restriction on \( \sigma^2_u \). As is well known, the ADF and KPSS tests often yield conflicting evidence, and the results presented in Table 2 provide a vivid example of this notorious inconsistency. However, Carrion-i-Silvestre et al. (2001) proposed to conduct 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 in Table 2 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 nonstandard 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 $t_j = \hat{\theta}_j / \sigma(\hat{\theta}_j)$ 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 $\hat{e}$ 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., \( t_j^* = \hat{\theta}/\sigma(\hat{\theta}) \). 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, \( \hat{e}^{**} \), 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 \( t_k^{**} = \hat{\theta}^{**}/\sigma(\hat{\theta}^{**}) \) is computed. This procedure is repeated \( K \) times. Thus \( G^{**} \) is defined as the statistic that counts the number of times that \( t_k^{**} \) falls below \( t_k^* \) divided by the number of draws in the second bootstrap \( (K) \). Finally, the

\[ p\text{-value} = 1 - \# \{ G^{**} < G^* \}/J, \]

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 \( \psi L^{2/3} \) and \( \psi^{-1} L^{1/3} \), respectively, where \( \psi \equiv [0.5(0.95)^{-2}0.05(5-0.05)]^{1/3} \) when constructing confidence intervals at the 95% confidence

\[ \text{Fig. 2.} \] (a) Quebec export unit values to the U.S. and Japan from January 1992 to December 2003. (b) Manitoba export unit values to the U.S. and Japan from January 1992 to December 2003.
level. A number of practical considerations must also be taken into account. Due to the discrete nature of the empirical distribution, it is desirable that \((J + 1)/K\) 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 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 3 presents the OLS estimates of the ERPT 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 ERPT equations has the expected algebraic sign. The first regime ERPT coefficients being larger in the threshold value. Hence, there are 115 OLS regressions computed for each bootstrap sample leading to a total of \(115 \times 143^2 = 336,283,805\) regressions.

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\(7\) 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

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Fig. 3. Hog prices in Quebec and Manitoba from January 1992 to December 2003.

Fig. 4. Value of the foreign currency per Can$ weighted by the consumer food price index.
absolute value than their second-regime counterparts are consistent with hypotheses about menu cost and invoicing in the importing country’s currency. At first glance, invoicing in the importing country’s currency seems an especially likely hypothesis for exports to Japan because the first regime pass-through coefficients are close to 0. The second regime pass-through coefficients being between 0 and −1 are consistent with exercising market power. The coefficient of the hog price is significant except in the first regime of the Quebec–Japan and Manitoba–Japan ERPT equations. Statistically significant coefficients have the expected algebraic sign as an increase in the processors’ marginal cost is expected to induce an increase in 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 the processors’ marginal cost.

It should be noted that the threshold values reported in Table 3 do not provide a direct estimate of menu costs because the threshold variable is actually conditioned by menu costs and by the structural demand parameters 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 hypothesis of no market power, which is a test about the statistical significance of the coefficient of the exchange rate in the second regime of the ERPT equation ($H_0 : \theta_{1,2} = 0$). The null hypothesis 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. 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. It could be that their investment in a solid business relationship allows them to exercise market power.

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 ($H_0 : \theta_{1,1} = 0$). This hypothesis is rejected in all equations at the 90 percent confidence level. The hypothesis about invoicing in the importing country’s currency is a test of: $H_0 : \theta_{1,1} = -1$. The bootstrap simulations yield strong evidence in favor of this hypothesis for Quebec and Manitoba pork exports to Japan (respective $p$-values are 0.725 and 0.384). However, the statistical evidence rejects the invoicing in either Canadian or U.S. dollars 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$). 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. A highly stylized

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Note: Trade data which are used to compute export prices come from Statistics Canada while data on domestic prices come Agriculture and Agri-food Canada. Data on exchange rates come from Canada and Japan central banks.

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Table 1
Summary statistics of the variables in the exchange rate pass-through (ERPT) equations

<table>
<thead>
<tr>
<th>Variables</th>
<th>Average</th>
<th>Standard Error</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>Quebec</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Export price to the U.S. (Can$/kg)</td>
<td>3.07</td>
<td>0.41</td>
<td>1.92</td>
<td>4.11</td>
</tr>
<tr>
<td>Export price to Japan (Can$/kg)</td>
<td>4.89</td>
<td>1.23</td>
<td>2.98</td>
<td>7.35</td>
</tr>
<tr>
<td>Hog price (Can$/kg)</td>
<td>1.57</td>
<td>0.31</td>
<td>0.68</td>
<td>2.30</td>
</tr>
<tr>
<td>Manitoba</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Export price to the U.S. (Can$/kg)</td>
<td>2.46</td>
<td>0.41</td>
<td>1.50</td>
<td>3.49</td>
</tr>
<tr>
<td>Export price to Japan (Can$/kg)</td>
<td>5.15</td>
<td>0.96</td>
<td>2.83</td>
<td>8.04</td>
</tr>
<tr>
<td>Hog price (Can$/kg)</td>
<td>1.50</td>
<td>0.27</td>
<td>0.61</td>
<td>2.16</td>
</tr>
<tr>
<td>U.S. weighted exchange rate</td>
<td>0.45</td>
<td>0.07</td>
<td>0.35</td>
<td>0.62</td>
</tr>
<tr>
<td>Japan weighted exchange rate</td>
<td>0.80</td>
<td>0.09</td>
<td>0.58</td>
<td>1.04</td>
</tr>
</tbody>
</table>

Note: Trade data which are used to compute export prices come from Statistics Canada while data on domestic prices come Agriculture and Agri-food Canada. Data on exchange rates come from Canada and Japan central banks.

Table 2
Unit root testing

<table>
<thead>
<tr>
<th>Variables</th>
<th>ADF test</th>
<th>KPSS Test</th>
<th>Joint confirmation of a unit root</th>
</tr>
</thead>
<tbody>
<tr>
<td>Quebec</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Export price to the U.S. (NT)</td>
<td>0</td>
<td>3.37*</td>
<td>0.46* No</td>
</tr>
<tr>
<td>Export price to Japan (T)</td>
<td>0</td>
<td>3.61*</td>
<td>0.22* No</td>
</tr>
<tr>
<td>Hog price (NT)</td>
<td>0</td>
<td>3.06*</td>
<td>0.22 No</td>
</tr>
<tr>
<td>Manitoba</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Export price to the U.S. (NT)</td>
<td>0</td>
<td>3.45*</td>
<td>0.84* No</td>
</tr>
<tr>
<td>Export price to Japan (T)</td>
<td>2</td>
<td>3.13**</td>
<td>0.34* No</td>
</tr>
<tr>
<td>Hog price (NT)</td>
<td>0</td>
<td>3.45*</td>
<td>0.13 No</td>
</tr>
<tr>
<td>U.S. weighted exchange rate (NT)</td>
<td>0</td>
<td>2.70**</td>
<td>2.78* Yes</td>
</tr>
<tr>
<td>Japan weighted exchange rate (NT)</td>
<td>1</td>
<td>2.63**</td>
<td>0.23 No</td>
</tr>
</tbody>
</table>

Notes: 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).
The numerical model also confirmed that sequential least square estimates obtained from a data generating process averaging over two types of firms (invoicing in U.S. or Can$) can generate pass-through coefficients similar the ones reported in Table 3.

The null of no threshold is tested using the likelihood ratio statistic proposed by Hansen (2000), \( LR(y = 0) = T(S(y = 0) - S(\hat{y}))/S(\hat{y}) \); where \( S(\hat{y}) \) and \( S(y = 0) \) 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.\(^9\) The inability to reject 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 nonrejection 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 market power and currency invoicing, taking as valid the specification of our two-regime ERPT equations.

The individual hypotheses about the absence of market power 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 market power and domestic currency invoicing. The inference strategy is to write the restrictions on the parameter in (15) as \( H_0 : R\Theta = r \), where \( \Theta = [\theta \delta_T] \) is a \( 6 \times 1 \) vector and the

\(^9\) 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–U.S. 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.
matrix $\mathbf{R}$ selects the appropriate elements from the vector $\theta$ to
be restricted according to $\mathbf{r}$. Under the null hypothesis, we have
that $\mathbf{R} = \begin{bmatrix} 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$ and $\mathbf{r} = \begin{bmatrix} 0 \\ 0 \\ 0 \end{bmatrix}$. The test statistic is
$F = (\mathbf{R} \hat{\theta} - \mathbf{r}) (\mathbf{R} \mathbf{cov} (\hat{\theta}) \mathbf{R}^{-1} (\mathbf{R} \hat{\theta} - \mathbf{r}) / 2$. The inference is made
possible by using bootstrap samples as described previously.

The across-destinations differences in our results suggest
that the choice of currency for invoicing might be conditioned
by destination-specific factors beside heterogeneity among ex-
porting firms. Overall, our results are consistent with the styl-
ized facts of world pork trade, Canada being a major player
on the world scene, and with the differences in the domes-
tic market structures of the Quebec and Manitoba processing
sectors.

5. Conclusion

This article developed a theoretical ERPT framework ac-
counting 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 nonlinearity between the exchange rate and the ex-
port price. This nonlinearity motivates the empirical specifica-
tion 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

pursposes 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
nonrejection in the case of the Quebec–Japan ERPT equation is
more likely attributable to the small length of our sample than
to the actual significance of menu costs faced by Quebec firms.

Invoicing in the importing country’s currency and market
power appear to characterize the behavior of Quebec exporters
in their dealings with Japanese importers. We also found evidence
of market power 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 atti-
dudes toward risk. The evidence of price discriminating behavior
is weak for Manitoba pork exports to Japan. The individual null
hypothesis about no market power could not be rejected and the
same can be said about the joint hypothesis about the absence
of market power and foreign currency invoicing.

The across-destinations differences in our results suggest
that the choice of currency for invoicing might be conditioned
by destination-specific factors beside heterogeneity among ex-
porting firms. Overall, our results are consistent with the styl-
ized facts of world pork trade, Canada being a major player
on the world scene, and with the differences in the domes-
tic market structures of the Quebec and Manitoba processing
sectors.

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