

# Is Exchange Rate Pass-Through in Pork Meat Export Prices Constrained by the Supply of Live Hogs?<sup>1</sup>

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*Abstract:* This paper investigates the effects of lags in production and marketing of agricultural products on Exchange Rate Pass-Through (ERPT) in export prices of processed agri-food commodities. The short-run inelastic supply of agricultural goods (such as grains and livestock) is tied to processors' export pricing decisions under certain marketing arrangements. The impacts of pre-determined supplies on the level of ERPT can be measured by regressing export prices on a destination-specific exchange rate but also on processed output and exchange rates in other export markets. The theoretical predictions of the model are tested by investigating ERPT in Canadian pork export prices in the U.S. and Japan. The empirical methodology accounts for potential unit root and cointegration using a dynamic seemingly unrelated regression framework. Pre-determined hog supplies have a statistically significant impact on ERPT. In the case of exports to Japan, the degree of misspecification involved with standard ERPT equations is quite large as they fail to reject the null hypothesis of no ERPT while significant ERPT coefficients are found in the full system approach that accounts for production lags.

*Keywords:* Exchange rate pass-through, Production lags, Dynamic seemingly unrelated regression, Canadian pork exports.

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

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# **Is Exchange Rate Pass-Through in Pork Meat Export Prices Constrained by the Supply of Live Hogs?**

## **Introduction**

Broad globalization pressures and increased concentration in downstream levels of agri-food supply chains set off a keen interest on the relationship between the nature of competition in international markets and observations of export price and exchange rates. One segment of the literature deals with Exchange Rate Pass-Through (ERPT) which can be broadly defined as the export price movement following changes in exchange rates (Bowen, Hollander and Viaene, 1998). Some notable studies include Pick and Carter's (1994) wheat study, Griffith and Mullen's (2001) analysis about Australia's rice exports, Brown's (2001) analysis of Canadian canola exporters and Gervais and Larue's (2004) analysis of Canadian pork exports. Not surprisingly, export price responses differ across commodities and countries and depend on many factors such as market structure, consumer preferences, product characteristics, etc.

One issue that has been neglected in the literature is whether firms have unlimited capacity when adjusting export prices following changes in the exchange rate. Knetter (1994) suggested that export price adjustments were likely to be linked to whether firms face capacity constraints in distribution networks or quantitative restrictions in export markets. In his framework, bottlenecks at the border are revealed through asymmetric adjustment in export prices. Bughin (1996) used panel data from Belgian manufactures to estimate a cost function under potential capacity constraints. He finds that the degree of mark-up adjustment following currency movements is significantly linked to each firm's capacity constraint.

In the context of agricultural supply chains, capacity constraints in downstream markets can stem from the usual short-run fixity of an input (*e.g.*, stock of capital) or lengthy lags between production and marketing of primary agricultural goods. These lags are especially

lengthy in livestock and grain industries whose production decisions precede marketing decisions by several months. For example, a currency depreciation may not trigger immediate increases in exports of processed commodities because the supply of upstream producers can be perfectly inelastic in the short-run. Moreover, processing firms and agricultural producers have developed different marketing strategies (*e.g.*, contracts) that often imply binding commitments with respect to price and output. Hence, even though a processing firm is faced with unexpected and unfavourable movement in its terms of trade, it may sell more output than it would otherwise choose to because of binding marketing agreements.

The objectives of the paper are twofold. First, a theoretical model that accounts for production and marketing lags is used to explain the pricing decisions of an exporting firm that sells a processed good into two export markets. At the marketing stage, the downstream firm's capacity may be sunk due to the inelastic supply of the primary agricultural commodity or because of predetermined purchases of the raw input. We refer to these scenarios as a capacity constraint from the processing firm's perspective. If the capacity constraint is not binding, the mark-up of the firm in a given export market reduces to the standard monopoly pricing rule under the assumption of constant returns to scale in processing. If the capacity constraint is binding, the destination-specific mark-up of the firm is function not only of the exchange rate in that particular market, but also of the exchange rate in the other export market.

The second objective is to measure ERPT in pork meat export prices from three Canadian provinces (Ontario, Quebec and Manitoba) to two destinations (U.S., and Japan) over the 1988-2003 period. The empirical model of Knetter (1989, 1993) is modified to test the theoretical finding that if export prices of processed pork meat are constrained by the supply of live hogs,

the number of hogs slaughtered and the exchange rate in the other market should explain the export pricing decisions.

There are two main empirical challenges when estimating ERPT in Canadian pork export prices. First, it is usual to find that export prices and exchange rates possess a unit root (see for example Carew and Florkowski, 2003; Choudhri, Faruqee and Hakura, 2005). As such, the empirical model must account for the potential non-stationarity of the data when estimating the model. Second, efficiency gains in the estimation can be captured by estimating the ERPT equation of each province to a given market simultaneously. These two issues are addressed by using the Dynamic Seemingly Unrelated Regression (DSUR) model proposed by Mark *et al.* (2005) and Moon and Perron (2004). Their strategy is to add leads and lags of the independent variables and use Generalized Least Squares (GLS) estimation to correct for potential endogeneity bias and autocorrelation in the residuals, respectively.

The results suggest that pork packers' volumes in Quebec and Ontario have significant impacts on the degree of exchange rate pass-through while hog supplies in Manitoba do not influence export prices. The empirical model also reveals significant differences between estimates of ERPT obtained in the current framework and the ones obtained using a standard specification that implicitly assumes instantaneous adjustment in output. The ERPT elasticities are statistically different than zero and thus suggest that Canadian pork exporters adjust their margin in response to fluctuations in the relative value of currencies. Finally, our results suggest that previously reported ERPT effects may be subject to some statistical bias when lags in production and marketing decisions are ignored.

The remainder of the paper is structured as follows. The next section presents a theoretical model that explains pricing decisions in a framework that accounts for marketing lags

in agri-food supply chains. Section 3 introduces the data used in the empirical model and investigates the statistical properties of the variables. Section 4 presents the empirical model and discusses the estimation results. Concluding remarks are presented in the last section.

### The Theoretical Model

Consider a processing firm that exports pork meat to two segmented foreign markets, identified by  $a$  and  $b$ . The demand in each market for the domestic firm's output is:  $D^a(e^a p^a, \bar{p}^a)$  and  $D^b(e^b p^b, \bar{p}^b)$ ; where  $p^j$  is the price set in market  $j$  by the domestic firm (in the domestic country's currency) and  $e^j$  is the exchange rate defined as the value of country  $j$ 's currency per unit of domestic currency. The variable  $\bar{p}^j$  is the price level of foreign substitute products in market  $i$ . The model uses the Armington assumption and purposely avoids modeling the interaction between the domestic and foreign firms.<sup>1</sup> This assumption is made for analytical convenience but does not qualify the result given that the emphasis of the paper is not to relate the degree of ERPT to market structure. For simplicity, it is also assumed that there are no domestic sales.

The processing firm maximizes profits defined as:

$$(1) \quad \pi = (p^a - t^a)D^a(e^a p^a, \bar{p}^a) + (p^b - t^b)D^b(e^b p^b, \bar{p}^b) - r^p Q^p;$$

where  $t^j$  measures the transportation cost between the domestic market and destination  $j$ ,  $r^p$  is the price of live hogs paid to domestic hog producers and  $Q^p$  represents live hogs that are purchased by the firm. Processing marginal costs are assumed constant and normalized to zero.

There are many ways to secure a desired supply of live hogs for the processing firm. The current analysis will focus on two hog marketing mechanisms: 1) the processor relies on the spot market to purchase hogs; and 2) contracts between the processor and individual hog suppliers

specify quantities to be delivered and the price that will be paid upon delivery. Additional assumptions about hog marketing arrangements are that: 1) live hogs are homogenous products (unlike the processed commodity); 2) hog producers are price takers in the world market; 3) that market  $a$  is closer to the domestic country than market  $b$  (such that  $t^a < t^b$ ); and 4) the domestic firm has monopsony power in the domestic market when purchasing live hogs.

The above assumptions are not unrealistic in the context of the Canadian hog/pork industry (Larue, Gervais and Lapan, 2004); but also apply to numerous other sectors that experienced increased concentration in downstream market levels. If the processing firm relies on the spot market to buy live hogs, it can capture all of the available output (denoted by  $Q^r$ ) at a price of  $e^a r^a - \mu t^a$ ; where  $r^a$  is the price of live hogs country  $a$ 's currency and  $\mu$  is a proportionality factor between transportation costs of the processed and raw commodities. As Larue, Gervais and Lapan (2004) argue, the possibility of a hold-up by the processing firm is constrained by the existence of an export market for the raw commodity. There is the possibility however that not enough hogs are produced from the processor's perspective because rational hog producers anticipate that the best price they receive is the price in market  $a$  adjusted for transportation costs.

One option would be for the processing firm to commit its price (to a level higher than the expected value of  $e^a r^a - \mu t^a$ ) before hog producers make their decision and thus effectively setting the quantity of hogs available in the marketing period. Even though contracting is possible, the processor *ex-ante* optimal demand for live hogs does not necessarily coincide with his *ex-post* optimal capacity choice once the values of the exchange rates are realized. Under price commitment, capacity is sunk at the marketing stage unless hog producers anticipate exporting hogs to market  $a$  in the marketing period.

In the spot market scenario, the processing firm can simply wait until the hogs attain market-ready weight to secure its supply of live hogs. It faces the possibility that its desired demand be higher than the available supply. The various hog marketing mechanisms in Canada provide us with a rich and diversified economic environment to test the theoretical predictions of the model. For example, hog marketing mechanisms in Quebec address coordination issues between packers and producers by relying on some hybrid marketing schemes. In short, a marketing board has exclusive rights to market hogs to processors. An important share of all hogs available in any given period is allocated to processors at a predetermined price based on their historical market shares while the remainder is auctioned off (Larue *et al.*, 2000). Hog marketing mechanisms in other provinces involve contracts between individual packers and hog producers as well as spot market transactions.

Going back to the profit maximization problem defined in (1), suppose that prior to export pricing decisions, the firm committed to buy a quantity  $Q^p$  of live hogs. The processing firm makes pricing decisions in the foreign market subject to the constraint that  $Q^p = D^a + D^b$ . Given the foreign price level of substitutes (denoted by  $\bar{p}^a$  and  $\bar{p}^b$ ), the first-order conditions are:

$$(2) \quad \frac{\partial \pi}{\partial p^a} = D^a(e^a p^a, \bar{p}^a) + e^a(p^a - t^a) \frac{\partial D^a}{\partial(e^a p^a)} - \lambda e^a \frac{\partial D^a}{\partial(e^a p^a)} = 0$$

$$(3) \quad \frac{\partial \pi}{\partial p^b} = D^b(e^b p^b, \bar{p}^b) + (p^b - t^b) e^b \frac{\partial D^b}{\partial(e^b p^b)} - \lambda e^b \frac{\partial D^b}{\partial(e^b p^b)} = 0$$

where  $\lambda$  is the Lagrange multiplier associated with the capacity constraint. Equations (2) and (3) define the domestic firm's reaction function  $p^a(e^a, e^b, Q^p; \bar{p}^a, \bar{p}^b)$  and  $p^b(e^a, e^b, Q^p; \bar{p}^a, \bar{p}^b)$  which can be substituted back into (1) to yield:

$$(4) \quad \pi(\bullet) = RT(e^a, e^b, Q^p; \bar{p}^a, \bar{p}^b) - r^p Q^p.$$

In the first scenario, the price commitment of the processing firm is made before hog producers make their sunk investment decisions. In the first stage, we assume that hog producers' supply is  $Q^r(r^p)$  with  $Q^{r'} > 0$ .<sup>2</sup> Because of its monopsony position in the purchase of domestic hogs, the risk-neutral processing firm maximizes:

$$(5) \quad E[\pi(\cdot)] = E\left[RT(e^a, e^b, Q^r(r^p); \bar{p}^a, \bar{p}^b)\right] - r^p Q^r(r^p).$$

The first-order condition to the optimization problem in (5) yields the optimal live hog price:  $r^{p*} = \phi(e^a, e^b; \bar{p}^a, \bar{p}^b)$  which is function of the various moments of the distribution of the exchange rates and the foreign firms' prices.

In the second case, the domestic firm uses the spot market to purchase live hogs and  $r^p$  is chosen when uncertainty about the exchange rate in world markets is resolved. However, hog supply is perfectly inelastic at that point; and the processor knows it can buy as many hogs as there are available ( $Q^r$ ) as long as it offers at least  $e^a r^a - \mu t^a$ . Let the parameter  $\theta$  be the Lagrange multiplier associated with the inequality  $Q^p \leq Q^r$ . If  $Q^r > Q^p$  ( $\theta = 0$ ), the processor does not face any constraint *ex-post* when setting export prices ( $D^a + D^b = Q^p < Q^r$ ) and equation (1) becomes:

$$(6) \quad \pi = (p^a - t^a)D^a(e^a p^a, \bar{p}^a) + (p^b - t^b)D^b(e^b p^b, \bar{p}^b) - r^p (D^a(e^a p^a, \bar{p}^a) + D^b(e^b p^b, \bar{p}^b)).$$

The processing firm maximizes (6) subject to the constraint  $r^p = e^a r^a - \mu t^a$ . The first-order conditions are:

$$(7) \quad \partial\pi/\partial p^a = D^a(e^a p^a, \bar{p}^a) + e^a (p^a - t^a - r^p) (\partial D^a / \partial (e^a p^a)) = 0,$$

$$(8) \quad \partial\pi/\partial p^b = D^b(e^b p^b, \bar{p}^b) + e^b(p^b - t^b - r^p)(\partial D^b/\partial(e^b p^b)) = 0.$$

The first-order conditions in (7) and (8) can be manipulated to yield the standard elasticity pricing formula of Knetter (1989). The equilibrium prices defined by (7) and (8) are:  $p^a(e^a; \bar{p}^a, r^p)$  and  $p^b(e^b; \bar{p}^b, r^p)$ .

However, if the processors' demand for live hogs is equal to the (perfectly inelastic) supply of hogs ( $Q^p = Q^r$ ), the optimization problem of the processor when selecting export prices reduces to equations (2) and (3). As Larue, Gervais and Lapan (2004) argued, if the processor does not commit its output price, it has no incentive to raise prices above the net marginal revenue that hog producers can obtain in the export market once hogs attain ready-to-market weight. Producers are rational and fully anticipate that outcome thus leading to a potential "low-price, low-capacity trap".

Export marginal costs are not identical across destinations because transportation costs differ. Moreover, Canadian products can have closer substitutes in one market relative to another. These factors are likely to impact the effect of the capacity constraint on the degree of exchange rate pass-through. Based on the previous theoretical set-up, the empirical model needs to distinguish between two general cases. In the first instance, production of live hogs will impact ERPT either through the price commitment made before hog marketing transactions or because hog supplies are pre-determined (*i.e.*, inelastic hog supply) when export pricing decisions are made and that not enough hogs are available. For example, consider favourable movements in the exchange rates that were unexpected when the processor's price commitment was made. The changes would normally induce additional sales in export markets but additional purchases on the spot market may not be possible due to the inelasticity of the short-run hog supply. Similarly, commitments made in the first stage can also influence ERPT when there are

unfavourable movements in the exchange rate because the domestic firm's purchases of live hogs are sunk at this stage. In the second general situation, the domestic firm relies exclusively on the spot market and the supply of live hogs does not constrain the domestic firm. At the observed domestic price, exports of live hogs should be positive.

Before formally investigating the exchange rate pass-through effects using comparative static tools, Figures 1 through 3 in the appendix illustrate the economic intuition of the model. In all instances, the segments  $ED^a$  and  $ED^b$  represent the (perceived) residual demand for the Canadian product in each market. There are two marginal revenue segments in each market; the lower segment accounts for transportation costs. Consider the case in which the processor relies entirely on the spot market to purchase live hogs and the quantity of available hogs does not put any constraint on the processor's exporting decisions. Under that scenario, the profit-maximizing solutions ( $p^{a*}$  and  $p^{b*}$  in figure 1) are defined by (7) and (8). Marginal revenue in each export market equals marginal cost. Hence, the export price in market  $j$  will be function of the exchange rate in  $j$  and the price of live hogs. Figure 2 assumes that the domestic firm faces a binding capacity constraint at  $Q^p$ . Marginal revenue functions in each market must be equal when evaluated at the profit maximizing solution and  $D^{a*} + D^{b*} = Q^p$ . Marginal revenue is higher than the price of live hogs because the processor is rationed. The export price will be function of the two markets' exchange rates and sunk capacity.

Most of the same intuition will apply in the case that live hogs are purchased under contract. However, there is one interesting case in which, because of the price commitment made before uncertainty about exchange rates is realized, marginal revenues in each market may end-up below marginal cost (*i.e.*, the price of live hogs). The processing firm maximizes revenue at this stage, and thus it is possible that *ex-post* demand for hogs be smaller than its *ex-ante*

commitment. Once again, the export price in each market will be function of the two exchange rates as well as (pre-determined) output.

Comparative static exercises can be carried out on the set of first-order conditions in (2)-(3) or (7)-(8) which define equilibrium price. The latter first-order conditions are independent from each other and, provided that the demand in country  $j$  is not too convex, it is relatively easy to show that:

$$(9) \quad \frac{\partial p^j}{\partial e^j} = \frac{-(2p^j - t^j - r^p) \left( \frac{\partial D^j}{\partial (e^j p^j)} \right) - e^j p^j (p^j - t^j - r^p) \left( \frac{\partial^2 D^j}{\partial (e^j p^j)^2} \right)}{2e^j \left( \frac{\partial D^j}{\partial (e^j p^j)} \right) + (e^j)^2 (p^j - t^j - r^p) \left( \frac{\partial^2 D^j}{\partial (e^j p^j)^2} \right)} < 0.$$

Equation (9) illustrates the standard result that depreciation (appreciation) of the domestic currency will lower (increase) the export price, albeit in a lesser proportion.

The comparative static exercise for the binding capacity constraint case is a little more involved. Assume for simplicity that the demand in each market is linear in its arguments. Totally differentiate the set of first-order conditions in (2) and (3) and the constraint to obtain:

$$(10) \quad (2p^a - t^a) \frac{\partial D^a}{\partial (e^a p^a)} de^a - \lambda \frac{\partial D^a}{\partial (e^a p^a)} de^a + 2e^a \frac{\partial D^a}{\partial (e^a p^a)} dp^a - e^a \frac{\partial D^a}{\partial (e^a p^a)} d\lambda = 0$$

$$(11) \quad (2p^b - t^b) \frac{\partial D^b}{\partial (e^b p^b)} de^b - \lambda \frac{\partial D^b}{\partial (e^b p^b)} de^b + 2e^b \frac{\partial D^b}{\partial (e^b p^b)} dp^b - e^b \frac{\partial D^b}{\partial (e^b p^b)} d\lambda = 0$$

$$(12) \quad \frac{\partial D^a}{\partial (e^a p^a)} (p^a de^a + e^a dp^a) + \frac{\partial D^b}{\partial (e^b p^b)} (p^b de^b + e^b dp^b) - dQ^p = 0$$

The system in (10), (11) and (12) can be solved using Cramer's rule to obtain:  $\partial p^j / \partial e^j < 0$  and  $\partial p^j / \partial e^i < 0$ ;  $i \neq j$ . A depreciation (appreciation) of the domestic currency with respect to country  $a$ 's currency will increase (decrease) the export price in market  $a$  (albeit in a lesser proportion than when pass-through is unconstrained). The incomplete pass-through will increase

(decrease) sales in market  $a$  and will decrease (increase) sales in market  $b$  because supply is sunk at this stage. The decrease (increase) in sales to market  $b$  must necessarily be induced by an increase (decrease) in the export price in market  $b$ .

### **Data**

Hog production data from January 1988 to November 2003 for the provinces of Manitoba, Ontario and Quebec were obtained from the Red meat market division of Agriculture and Agri-food Canada. The three provinces accounted for more than 75% of all hogs marketed in Canada in 2003. Figure 4 presents total monthly hog slaughterings in each province. It clearly illustrates the growth in the Canadian hog industry beginning in the mid-90s. There are some significant differences between total hog marketings and hog slaughterings within a province in some instances. The monthly ratio of hogs slaughtered over total hogs marketed in each province (on a scale of one hundred) is presented in figure 5. Processors in Quebec almost always slaughtered all available hogs in the province while a significant portion of total hog production in Ontario is sold to packers outside Ontario. The ratio for Manitoba is less indicative because there is a period in the mid nineties when total hogs marketed exceeded total hogs slaughtered in the province. These simple stylized facts will prove to be important when empirically analyzing ERPT in the next section.

Data on monthly pork exports from Quebec, Ontario and Manitoba between January 1988 and November 2003 were collected from Statistics Canada. Figure 6a, 6b and 6c presents pork exports from Quebec, Ontario and Manitoba respectively to the two most important export destinations: the U.S. and Japan. Export prices are proxied by export unit values. Figure 7a, 7b and 7c presents pork export unit values (in Can\$) to each destination from Quebec, Ontario and Manitoba respectively. Export unit values differ substantially across destinations; suggesting that

the export product mix could be quite different across each destination and thus could be a source of bias in the price indexes (Kravis and Lipsey, 1974). Lavoie and Liu (2004) present Monte Carlo simulations that illustrate the caveats associated with using unit values to proxy export prices when commodities within a product category are significantly differentiated. Nevertheless, pork meat is not believed to be a significantly differentiated commodity and the analysis proceeds with unit values given the absence of a credible alternative to proxy prices.

Data on exchange rates between the Canadian dollar and the export market currency and food price indexes were obtained from the Central banks of the importing countries. Each exchange rate series is weighted by the food price index of the importing country to account for foreign firms' pricing strategies. Figure 8 presents an index (1988 = 100) of the exchange rate between the Can\$ and the currency in each destination weighted by the consumer food price index in that destination. Finally, live hog prices in all three provinces follow similar patterns over the sample period as illustrated in Figure 9.

The first step in the empirical model is to determine the stochastic properties of the data. For further reference, let superscripts  $QB, MB$  and  $ON$  indicate the source of exports as Quebec, Ontario and Manitoba respectively and superscripts  $US$  and  $JP$  indicate the destination markets as the United States and Japan respectively. The theoretical model suggests estimating the relation between the export price from a province to a specific destination and the price of live hogs, the supply of live hogs available to processors and exchange rates in each market. Hence, the variables used in the study are the export unit values (denoted by  $p^{j,m}$ ;  $j = QB, MB, ON$  and  $m = US, JP$ ), the exchange rate weighted by the food price index for each destination ( $e^m$ ;  $m = US, JP$ ), the hog price in each province ( $r^j$ ;  $j = QC, MB, ON$ ) and total hogs slaughtered in each province ( $Q^j$ ;  $j = QC, MB, ON$ ). At this stage, it is perhaps

instructive to discuss the proxy used to measure pre-determined hog supplies. There are significant movements in live animals between the three provinces. As such, the supply of live hogs in one province may yield a poor approximation of the total supply of live hogs available to processors in that province because hogs can be transferred from one province to another. Hence, the empirical model uses the total quantity of hogs slaughtered in the province as a proxy to measure if marketing arrangements and production lags have any impact on export pricing decisions. In the case in which these factors have no effects, total hogs slaughtered will not influence export prices as slaughters can be adjusted freely by relying on the spot market.

The standard Augmented Dickey-Fuller (ADF) test is performed to assess the degree of integration of the variables. The usual regression equation is (Hamilton, 1994):

$$(13) \quad \Delta y_t = \alpha + \beta t + \rho y_{t-1} + \sum_{j=1}^k \gamma_j \Delta y_{t-j} + \varepsilon_t; \quad t = 1, \dots, T;$$

where  $t$  is a time trend,  $T$  is the sample length and  $k$  measures the length of the lag in the dependent variable. The selection of this parameter is carried out using Ng and Perron (2001) modified Akaike Information Criterion (MAIC). The objective is to minimize:

$$MAIC(k) = \ln(\hat{\sigma}_k^2) + \frac{2(k + \tau)}{T - k_{\max}}; \text{ where } \hat{\sigma}_k^2 \text{ is the variance estimator and } k_{\max} \text{ is defined as the}$$

maximum degree of augmentation in (13) and is set to  $k_{\max} = \text{int}\{12(T/100)^{0.25}\} = 14$  (Cook and Manning, 2004). Visual inspection of the series determines whether a time trend is included in (13) or if the latter equation is estimated with the restriction:  $\beta = 0$ . The critical values of the ADF unit root tests are provided in McKinnon (1991).

Table 1 presents the ADF test results. Three variables out of the 14 variables used in the empirical analysis have a statistic that exceeds the critical value of the test at the 5% significance

level. It is however notorious that unit root tests suffer from low power and tend to under-reject the null hypothesis of a unit root (Maddala and Kim, 1998). Hence, table 1 also reports the stationarity test of Kwiatkowski *et al.* (KPSS, 1992). The KPSS test rejects the null hypothesis of stationarity for 10 out of the 14 possible cases. However, there are conflicting results between the two tests in some instances; a conclusion that is not unusual for this type of Confirmatory Data Analysis (CDA). Carrion-I-Sylvestre *et al.* (2001) tabulated different sets of critical values to account for the jointness in the distribution of the two test statistics under the CDA approach. Their procedure yields a more powerful unit root investigation than standard unit root testing. The joint hypothesis of a unit root cannot be rejected for 10 of the 14 variables. The appealing feature of the empirical model in the next section is that it does not require establishing *a priori* the degree of integration of the variables and the procedure is valid under a number of different hypotheses about the statistical behaviour of the series (Moon and Perron, 2004).

### **The Empirical Model**

Given that the joint null hypothesis of a unit root in most of the variables can not be rejected, we use the Dynamic Seemingly Unrelated Regression (DSUR) model of Mark *et al.* (2005) to estimate Exchange Rate Pass-Through (ERPT) effects.<sup>3</sup> The DSUR approach considers the possibility that the integrated variables are cointegrated and accounts for potential contemporaneous correlation between each province ERPT equation. It corrects the potential endogeneity in the regressors and correlation in the error terms by adding leads and lags of the independent variables and using feasible GLS respectively. The DSUR estimators are normally distributed asymptotically and thus inference can be carried out in the usual way. More importantly, the authors show that the DSUR estimator is more efficient in small samples than the class of non-parametric cointegration estimators for SUR models (*e.g.*, Moon, 1999).

The cointegration approach is appealing in the present context for many reasons. It is possible that small variations of the exchange rate can have temporary impacts on pricing decisions but would error-correct in the long-run because either processors make adjustments to their hog purchases or that the exchange rate moves in a different direction and thus cancels out previous disequilibrium pricing strategies. Cointegration among the variables specifically assumes that there exists a stable long-run relationship between the variables and admits the possibility that there are deviations from the long-run pricing decision in the short-run. Second, the potential endogeneity bias between export unit values and output is explicitly accounted for in the cointegration model.

We illustrate the DSUR approach using the ERPT equations in market  $m$  from origin  $j = QC, MB, ON$ . The ERPT equations are based on a linear approximation to the solutions defined in (7) and (8). There is one cointegrating regression equation for each origin:

$$(14) \quad \begin{aligned} p_t^{QC,m} &= \alpha^{QC} + \beta^{QC,US} e_t^{US} + \beta^{QC,JP} e_t^{JP} + \lambda^{QC} Q_t^{QC} + \phi^{QC} r_t^{QC} + u_t^{QC} \\ p_t^{MB,m} &= \alpha^{MB} + \beta^{MB,US} e_t^{US} + \beta^{MB,JP} e_t^{JP} + \lambda^{MB} Q_t^{ON} + \phi^{MB} r_t^{MB} + u_t^{MB} \\ p_t^{ON,m} &= \alpha^{ON} + \beta^{ON,US} e_t^{US} + \beta^{ON,JP} e_t^{JP} + \lambda^{ON} Q_t^{MB} + \phi^{ON} r_t^{ON} + u_t^{ON} \end{aligned}$$

It is assumed that  $u_t^j$  is a stationary autoregressive process of order  $p$  such that

$$u_t^j = \rho^j u_{t-1}^j + \sum_{h=1}^{p-1} \eta^{j,h} \Delta u_{t-h}^j + \omega_t^j \quad \text{with} \quad \omega_t^j = \chi^j \kappa_t^j + \xi_t^j.$$

The latter terms model the cross-sectional covariance between error terms. Potential correlation between the error terms in (14) and the first difference of the regressors (i.e. the endogeneity problem) is addressed by augmenting the system in (14) by leads and lags of the independent variables:

$$(15) \quad \begin{aligned} p_t^{j,m} &= \alpha^j + \beta^{j,US} e_t^{US} + \beta^{j,JP} e_t^{JP} + \lambda^j Q_t^j + \phi^j r_t^j \\ &+ \sum_{l=-k}^k \gamma_l^{j,US} \Delta e_{t-l}^{US} + \sum_{l=-k}^k \gamma_l^{j,JP} \Delta e_{t-l}^{JP} + \sum_{l=-k}^k \theta_l^j \Delta Q_{t-l}^j + \sum_{l=-k}^k \kappa_l^j \Delta r_{t-l}^j + u_t^j \end{aligned}$$

for  $j = QC, MB, ON$  and  $m = US, JP$ .

When (predetermined) hog supplies have an impact on pricing, the export price in a given market is independent of the exchange rate in the other destination and the quantity of hogs marketed in the province. Hence, testing the non-significance of hog production on pork meat pricing decisions of province  $j$  is a test of the null hypothesis:  $\beta^{j,-m} = \lambda^j = 0$ ; where  $\beta^{j,-m}$  is the parameter measuring the other market's exchange rate effect. Under the null hypothesis, the number of hogs slaughtered in the province should not be cointegrated with the export price.

Table 2 presents the estimation results for the ERPT equations in the U.S. market.<sup>4</sup> The coefficients of the lead and lagged variables and the constant are not reported out of concern for space. The number of leads and lags in (15) was selected using the MAIC and is equal to one for both the U.S. and Japan systems. The parameters were estimated using the two-step procedure of Mark *et al.* (2005). First, the residuals  $\hat{u}_t^j$  are obtained from applying OLS to each equation individually in (15). An AR(2) model was found to model adequately the autocorrelation in the predicted residuals of each equation. The AR(2) parameter estimates are used to obtain the predicted residuals  $\hat{\omega}_t^j$ . The latter series yields the long-run variance-covariance matrix used in the second step to get the DSUR estimator through GLS. Goodness of fit statistics were used to make sure that autocorrelation in the residuals has been purged by the estimation procedure and that the distribution of the residuals is consistent with normality. The Jarque-Bera normality test (Greene, 1998) does not reject the null hypothesis that the residuals of each series are normally distributed at conventional levels of significance. The autocorrelation coefficients of the residuals of the final model were small relative to their standard error.

Following Mark and Sul (2003), we first tested homogeneity restrictions in the system. If the homogeneity restrictions are not rejected by the data, the three equations in (15) should be

pooled and the estimation procedure should proceed with dynamic OLS. The homogeneity restriction for the ERPT coefficients ( $\beta^{QC,US} = \beta^{MB,US} = \beta^{ON,US}$ ) cannot be rejected. The Wald tests for the other homogeneity restrictions all yield *p-values* lower than 0.001. Hence, the estimation proceeded with the DSUR estimator. The ERPT coefficients in table 2 for each of the three provinces have the expected sign and are statistically different than zero at the five percent significance level. The coefficient of the live hog price is positive and significant (*p-value* less than 0.000) in all three provinces. The coefficient of hogs marketed is statistically significant in Quebec and Ontario and has the expected sign. However, the *p-value* of the null hypothesis  $\beta^{MB,US} = 0$  is larger than 0.10; thus suggesting that output in Manitoba does not influence ERPT. The coefficients of the yen exchange rate are statistically significant in the Quebec and Ontario ERPT cointegrating equations. However, these coefficients do not have the expected sign; a devaluation of the Canadian currency with respect to the Japanese yen should cause an increase in the export price in the U.S. market. The coefficient for the yen exchange rate in the Manitoba equation is not significant; thus strengthening the case that hog supplies do not constrain exchange rate pass-through behaviour in Manitoba.

When the joint null hypothesis of zero coefficients on the yen exchange rate and the capacity variable is tested, the *p-value* is 0.242 for Manitoba but less than 0.001 in Quebec and Ontario. The latter results provide additional evidence that the supply of live hogs constrains exchange rate pass-through in Quebec and Ontario pork prices for exports to the U.S. The results are also consistent with anecdotal evidence in the Canadian hog/pork industry. Quebec packers historically processed all hogs that were marketed by Quebec hog producers. While exports of live hogs from Ontario are significant (figure 5), marketing arrangements rely heavily on private contracts between producers and packers. Hence, output has an important effect on pricing

decisions even though packers have access to additional supplies on the spot market after uncertainty about exchange rate is resolved.

Table 3 presents the estimation results in the Japanese market.<sup>5</sup> The ERPT coefficient in the Japanese market is negative for all three provinces but has a relatively large standard error in the Quebec export price equation. The price of live hogs has the expected impact on Quebec and Ontario export prices and but its coefficient in the Manitoba equation is not significant and does not have the expected sign. As in the U.S. case, the coefficient of the number of hogs slaughtered within the province is significant in Quebec (and has the correct sign) but the number of hogs slaughtered in Manitoba and Ontario is however not significant in explaining the export price in Japan. This suggests that hog supplies do not significantly impact export prices. The coefficient for the U.S. dollar exchange rate is positive in the Ontario equation but insignificant in the other two equations. The Wald test of the joint null hypothesis  $\beta^{j,US} = \lambda^j = 0$  is strongly significant (*p-value* less than 0.000) for Ontario and Quebec; which indicates hog supplies have an important impact on the export price response. Once again, hog supplies in Manitoba do not seem to have a significant impact on export pricing decisions.

It is useful at this point to investigate pass-through coefficients using standard empirical techniques that do not account for lags between hog production and marketing decisions. The existence of lags can introduce some significant bias in ERPT coefficients if output and other currency exchange rates are omitted in the ERPT equations. The framework of Knetter (1989, 1993) is used to investigate potential misspecification problems associated with “traditional” ERPT equations. It is adapted to account for unit roots and cointegration and uses a somewhat different specification. Knetter’s ERPT equation assumes that marginal cost is exclusively function of time while processors’ marginal costs in this paper are proxied by hog prices.

Moreover, his latter study addresses non-stationarity in the data by first-differencing each variable while the 1989 study ignores the stochastic properties of the data. These are evident shortcomings that are tackled here with the estimation procedure.

Table 4 reports the ERPT coefficients for exports from one province to each market and the  $t$ -statistic of the null hypothesis that the coefficient is not statistically different than zero. ERPT coefficients for the U.S. are different from the estimates in Table 2 but not substantially so. Although, the previous results established that production lags have significant effects on ERPT, the basic Knetter equation does not find substantial misspecification problems when looking at the magnitude of ERPT. Conversely, the inference for the ERPT coefficients in the Japanese market in Table 4 is strikingly different than the inference reported in Table 3. When potential marketing lags are not considered, the empirical framework finds no evidence of significant exchange rate pass-through in the Canadian export price in Japan at the 95% confidence level. The empirical model that accounts for pre-determined live hog supplies finds significant ERPT coefficients at the 10% significance level in all three provinces and the ERPT elasticities are large for Ontario and Manitoba. Hence, the failure to account for lags in production and marketing activities in agri-food supply chains can create substantial biases in the estimation of ERPT effects.

### **Concluding Remarks**

Exchange Rate Pass-Through (ERPT) is broadly defined as the export price response following a movement in the relative value of the domestic currency over the currency in the export market. The analysis investigated how ERPT for processed agri-food commodities can be impacted by pre-determined supplies of raw agricultural goods. It has been customary in the literature to assume constant returns to scale (*e.g.*, Knetter, 1989) in order to separate out the firms' pricing

decisions across export markets. The current framework assumed that there exist significant lags between production and marketing decisions for goods such as grains and livestock. Under the assumption that processing firms commit to purchase raw agricultural products before marketing decisions occur, export pricing decisions in all markets for processed commodities are tied. Even in the case when a processor relies on the spot market to purchase raw agricultural commodities, the available supply of raw commodities can be lower than the desired demand of processors. The theoretical model leads to simple testable predictions about the impact of the pre-determined supply of the primary good on the degree of exchange rate pass-through in export prices.

Canadian pork meat export prices from three provinces to two destinations (U.S., and Japan) were collected to investigate ERPT. The empirical model tested the ERPT implications of pre-determined supplies by regressing the export price in a given market on the exchange rate, the price of live hogs, total processed output and the other export market's exchange rate. Potential non-stationary in the data and endogeneity bias are addressed using the Dynamic Seemingly Unrelated Regression (DSUR) framework proposed by Mark *et al.* (2005) and Moon and Perron (2004). The DSUR procedure uses feasible generalized least squares to account for potential autocorrelation in the residuals and it corrects the endogeneity bias by introducing leads and lags of the independent variables in the equations. It also captures the contemporaneous correlation between exports of the three Canadian provinces.

The estimation results strongly support the hypothesis that pre-determined supplies have a significant impact on the degree of ERPT for two out of three Canadian provinces. ERPT elasticity for Canadian exports to the U.S. is in the range of -0.42 to -0.60. In the case of exports to Japan, the degree of misspecification involved with standard ERPT equation is quite large. Standard ERPT equations usually include the Canadian dollar to yen exchange rate as well as a

marginal cost proxy. The standard specification fails to reject the null hypothesis of no ERPT while significant ERPT coefficients are found in the full system approach that includes pre-determined hog supplies. For Ontario and Manitoba, the ERPT elasticities are quite large. Hence, failure to account for the dynamic nature of agricultural supply chains may result in significantly biased estimates of ERPT.

One interesting extension to the current framework would involve using pre-determined supplies to investigate the selection of export markets. In some periods, exports to a particular destination are zero and thus no export unit values are available. This seriously impedes the ability to analyze exchange rate pass-through behaviour in emerging markets. Zero trade flows are not uncommon in the empirical trade literature but researchers have struggled to properly address the issue (Helpman, Melitz and Rubinstein, 2005). In the current context, zero exports imply that there are no proxies for the export price for that particular destination. A promising research avenue would perhaps involve using a two-stage estimation procedure in which: 1) trade flows are first explained by a set of independent variables (as in gravity models); and 2) the first stage results are used to correct the selection bias related to missing values when estimating the ERPT equation. Once again, the existence of production and marketing lags in agri-food supply chains is likely to influence the selection rule.

## References

- Bowen, H. P., A. Hollander and J-M. Viaene. 1998. *Applied International Trade Analysis*. Michigan University Press.
- Bughin, J. 1996. "Capacity Constraint and Export Performance: Theory and Evidence from Belgian Manufacturing." *Journal of Industrial Economics* 44:187-204.
- Brown, J. 2001. "Price Discrimination and Pricing to Market Behaviour of Canadian Canola Exporters." *American Journal of Agricultural Economics* 83: 1343-49.
- Carew, R., and W. J. Florkowski. 2003. "Pricing to Market Behaviour by Canadian and U.S. Agri-food Exporters: Evidence from Wheat, Pulse and Apples." *Canadian Journal of Agricultural Economics* 51: 139-59.
- Carrion-i-Silvestre, J. L., A. Sanso-i-Rossello and M. A. Ortuno. 2001. "Unit Root and Stationarity Tests' Wedding." *Economics Letters* 70: 1-8.
- Choudhri, E. U., H. Faruquee, and D. S. Hakura. 2005. "Explaining the Exchange Rate Pass-Through in Different Prices." *Journal of International Economics* 65:349-374.
- Cook, S., and N. Manning. 2004. "Lag Optimisation and Finite-Sample Size Distortion of Unit Root Tests." *Economics Letters* 84: 267-274.
- Froot, K. A., and P. D. Klemperer. 1989. "Exchange Rate Pass-Through When Exchange Rate Matters." *American Economic Review* 79: 637-654.
- Gervais, J-P., and B. Larue. 2004. "A Joint Test of Pricing-to-Market, Menu Costs and Currency Invoicing." Presented at the 2004 AAEA meetings in Denver.
- Greene, W. H. 1998. *Econometric Analysis*. NJ: Prentice Hall.
- Griffith, G. and J. Mullen. 2001. "Pricing-to-Market in NSW Rice Export Markets." *Australian Journal of Agricultural Economics* 45: 323-334.

- Hamilton, J. D. 1994. *Time Series Analysis*. Princeton, NJ: Princeton University Press.
- Helpman, E., M. Melitz, and Y. Rubinstein. 2005. *Trading Partners and Trading Volumes*, Working paper.
- Knetter, M. M. 1989. "Price Discrimination by U.S. and German Exporters." *American Economic Review* 79: 198-210.
- Knetter M. M. 1993. "International Comparisons of Pricing-to-Market Behaviour." *American Economic Review* 83: 473-486.
- Knetter, M. 1994. "Is Export Price Adjustment Asymmetric? Evaluating the Market Share and Marketing Bottlenecks Hypotheses." *Journal of International Money and Finance* 13: 55-70.
- Kravis, I. B., and R. E. Lipsey. 1974. "International Trade Prices and Price Proxies." In the *Role of the Computer in Economic and Social Research in Latin America*, ed. N. Ruggles, NBER, New York.
- Kwiatkowski, D., P. C. B. Phillips, P. Schmidt and Y. Shin. 1992. "Testing the Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure are we that Economic Time Series Have a Unit Root?" *Journal of Econometrics* 54: 159-178.
- Larue, B. J-P. Gervais, and H. E. Lapan. 2004. "Low-Price Low-Capacity Traps and Government Interventions in Agricultural Markets." *Canadian Journal of Agricultural Economics* 52: 237-256.
- Larue, B., R. Romain, J-P. Gervais and S. Ben Salha. 2000. "The Collusion Deterring Effect of Pre-Attributed Supplies and the Hog Auction in Quebec." *Canadian Journal of Agricultural Economics* 48: 607-622.

- Lavoie, N., and Q. Liu. 2002. *Pricing-to-Market: Price Discrimination or Product Differentiation?* Working paper 2004-011, Department of Resource Economics, University of Massachusetts at Amherst.
- Maddala, G. S., and I-M. Kim. 1998. *Unit Roots, Cointegration and Structural Change*, Cambridge University Press.
- Mark, N. C., M. Ogaki, and D. Sul. 2005. "Dynamic Seemingly Unrelated Cointegrating Regression." *Review of Economic Studies* 72: 797–820
- Mark, N. C., and D. Sul. 2003. "Cointegration Vector Estimation by Panel DOLS and Long-Run Money Demand." *Oxford Bulletin of Economic and Statistics* 65: 655-680.
- Moon, H. R., and B. Perron. 2004. "Efficient Estimation of the Seemingly Unrelated Regression Cointegration Model and Testing for Purchasing Power Parity." *Econometric Reviews* 23: 293-323.
- Moon, H. R. 1999. "A Note on Fully-Modified Estimation of Seemingly Unrelated Regressions Models with Integrated Regressors." *Economics Letters* 65: 25-31.
- Ng, S., and P. Perron. 2001. "Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power." *Econometrica* 69: 1519–1554.
- Pick, D. H., and C. A. Carter. 1994. "Pricing to Market with Transactions Denominated in a Common Currency." *American Journal of Agricultural Economics* 76: 55-60.

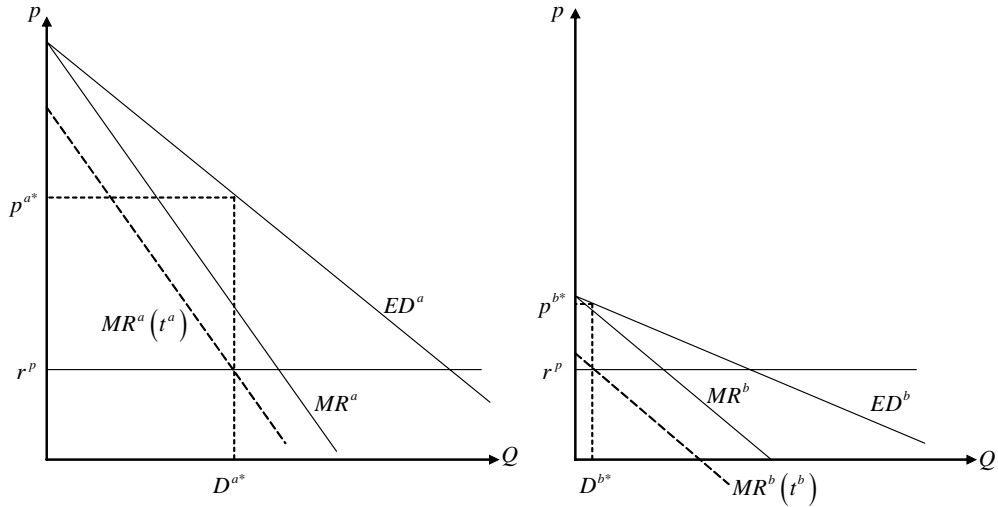


Figure 1.

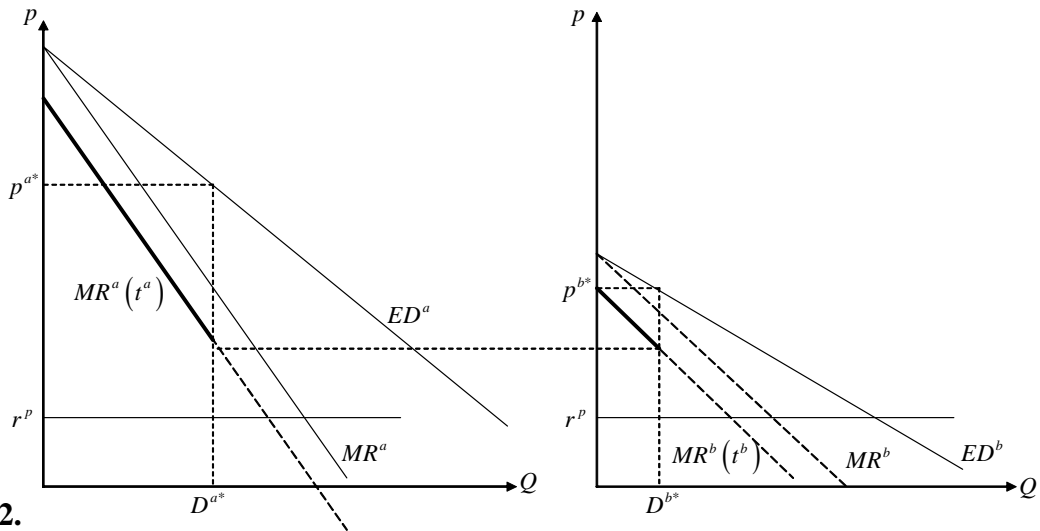


Figure 2.

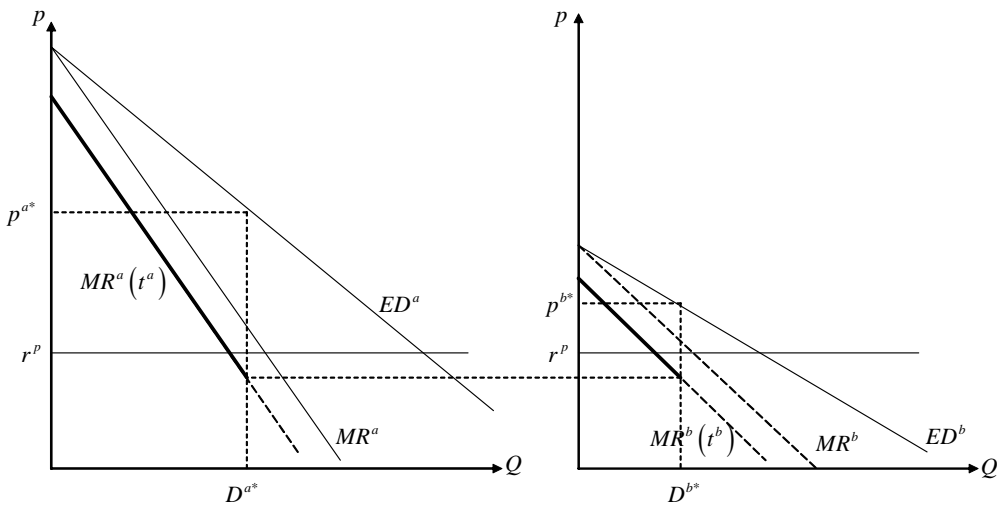
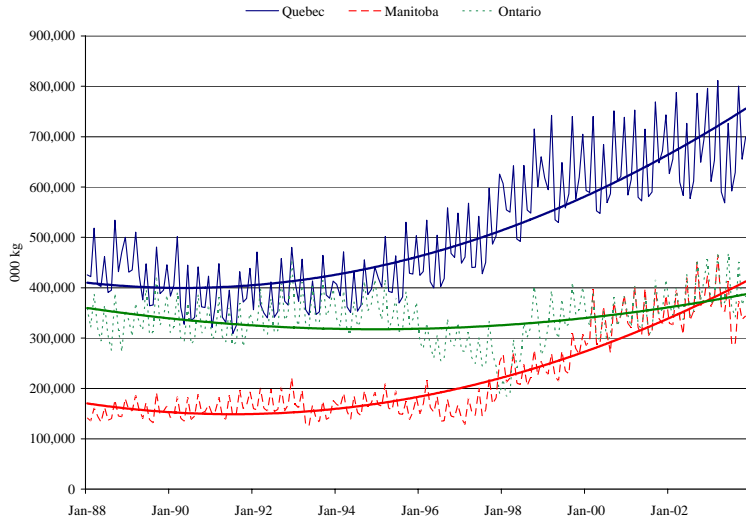
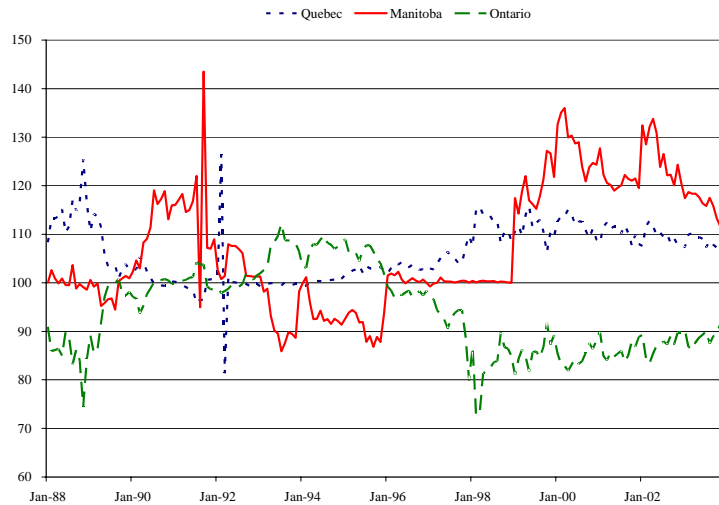


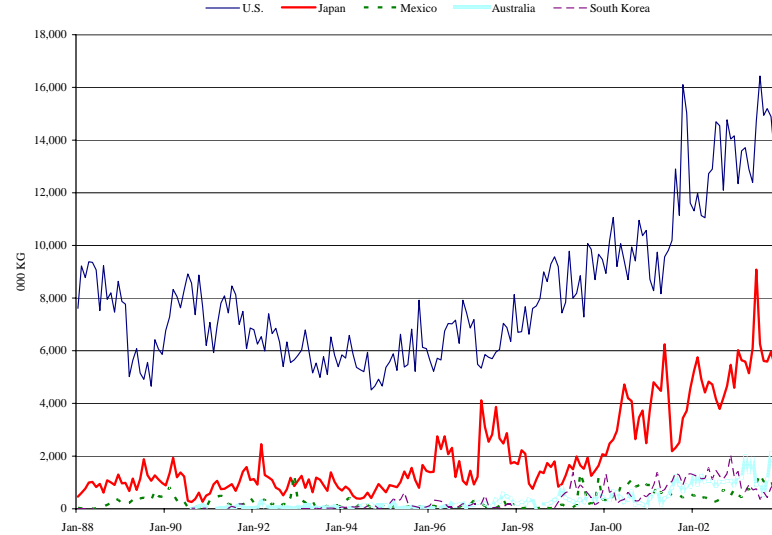
Figure 3.



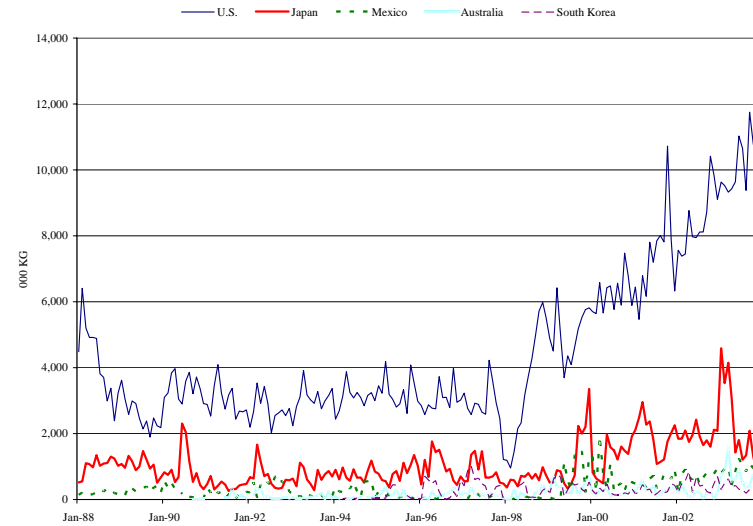
**Figure 4. Monthly hog slaughtering in Quebec, Ontario and Manitoba from January 1988 to November 2003.**



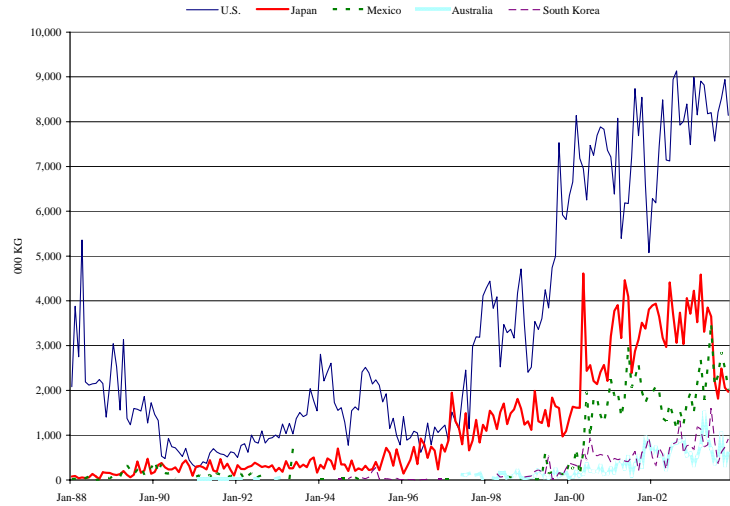
**Figure 5. Ratio of total hog slaughtered to total hog marketing in Quebec, Ontario and Manitoba from January 1988 to November 2003.**



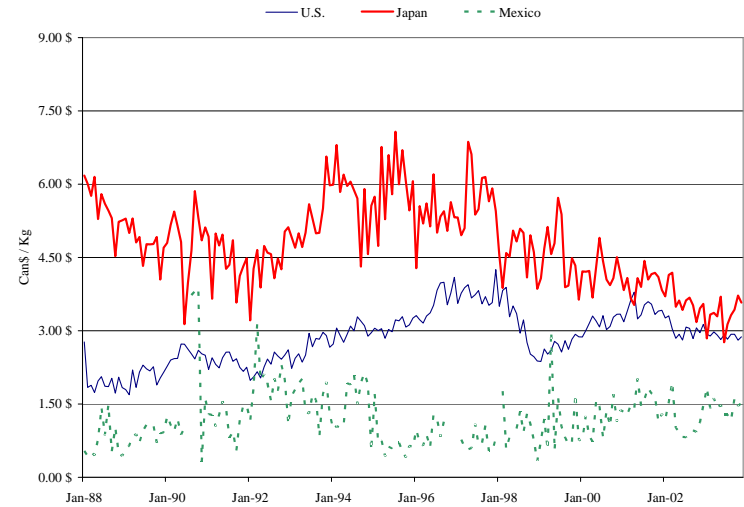
**Figure 6a. Quebec monthly pork exports to its five most important destinations from January 1988 to November 2003.**



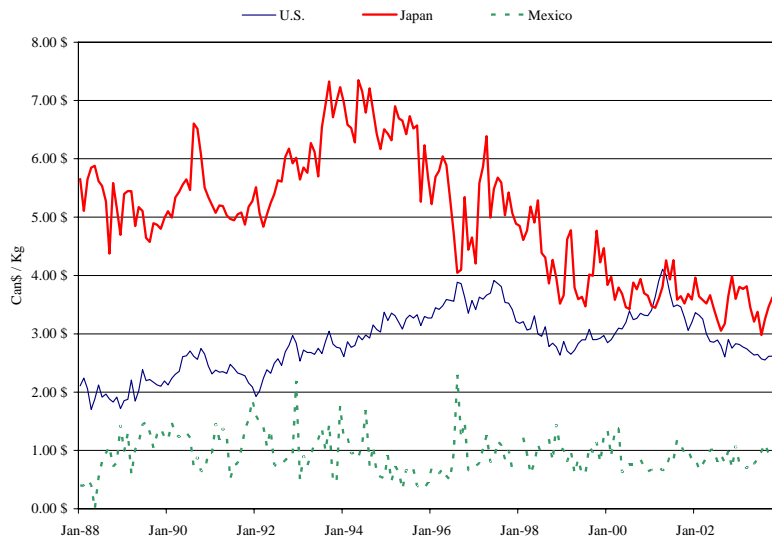
**Figure 6b. Ontario monthly pork exports to its five most important destinations from January 1988 to November 2003.**



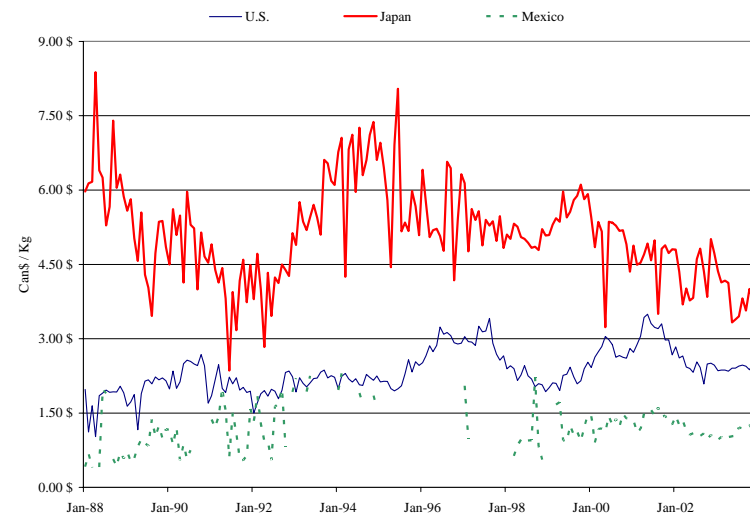
**Figure 6c. Manitoba monthly pork exports to its five most important destinations from January 1988 to November 2003.**



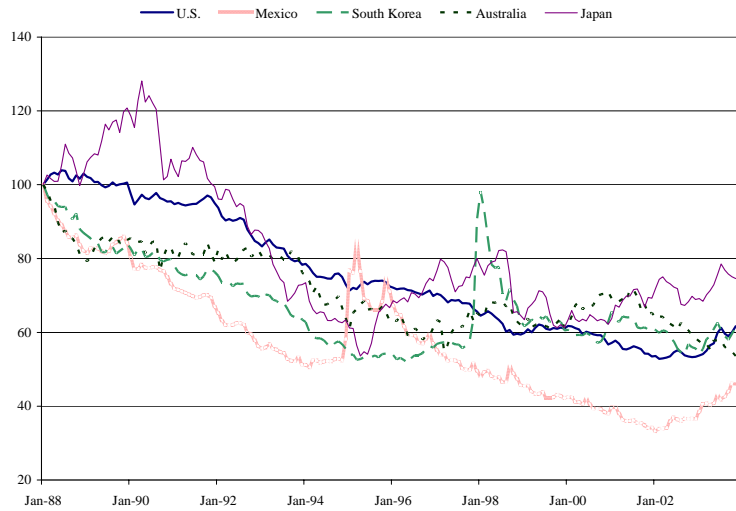
**Figure 7b. Monthly unit values of Ontario pork exports to each destination from January 1988 to November 2003.**



**Figure 7a. Monthly unit values of Quebec pork exports to each destination from January 1988 to November 2003.**



**Figure 7c. Monthly unit values of Manitoba pork exports to each destination from January 1988 to November 2003.**



**Figure 8. Monthly exchange rate index (Jan-1988 = 100) of the Can\$ with respect to foreign currencies weighted by food price indexes.**



**Figure 9. Monthly hog prices in Quebec, Ontario and Manitoba between January 1988 and November 2003.**

**Table 1. Unit root and stationarity tests**

Variables	Deterministic specification	ADF		KPSS	Joint confirmation of a unit root at the 5% significance level
		Lags	Test value	Test value	
$e^{US}$	T	1	-1.84	0.14	Yes
$e^{JP}$	T	2	-4.22	0.31	No
$Q^{QC}$	T	14	-3.41	0.60	Yes
$Q^{MB}$	T	11	-1.51	0.73	Yes
$Q^{ON}$	T	12	-1.77	0.37	Yes
$p^{QC,US}$	NT	2	-1.80	2.10	Yes
$p^{MB,US}$	NT	3	-3.86	1.77	No
$p^{ON,US}$	NT	7	-2.12	2.20	Yes
$p^{QC,JP}$	NT	7	-0.65	2.45	Yes
$p^{MB,JP}$	NT	6	-1.84	0.56	Yes
$p^{ON,JP}$	NT	10	-0.63	1.54	Yes
$r^{QC}$	NT	9	-2.12	0.23	Yes
$r^{MB}$	NT	9	-2.45	0.16	No
$r^{ON}$	NT	9	-2.24	0.07	No

Note: Critical values for the ADF test without a trend are: -3.481, -2.884 and -2.574 at the 1%, 5% and 10% significance levels respectively. Critical values for the ADF test with a trend are: -4.011, -3.439 and -3.139 at the 1%, 5% and 10% significance levels respectively. Critical values for the KPSS stationarity test without a trend are: 0.739, 0.463 and 0.347 at the 1%, 5% and 10% significance levels respectively. Critical values for the KPSS stationarity test with a trend are: 0.216, 0.146 and 0.119 at the 1%, 5% and 10% significance levels respectively. Finally, the joint critical values of the ADF and KPSS tests to perform the confirmatory analysis at the 5% significance level are respectively -3.126 and 0.171 when there is no trend and -3.74 and 0.073 when there is a trend. Finally, T and NT stand respectively for trend and no trend in the unit root test specified in (13).

**Table 2. Exchange rate pass-through coefficients in the U.S. market**

Variable	Quebec		Manitoba		Ontario	
	Coefficient	<i>t</i> -statistic	Coefficient	<i>t</i> -statistic	Coefficient	<i>t</i> -statistic
U.S. exchange rate, $e^{US}$	-0.596	-3.55	-0.510	-3.42	-0.423	-4.11
Japan exchange rate, $e^{JP}$	-0.220	-2.17	0.139	1.57	-0.270	-2.68
Capacity, $Q^j$	-0.275	-2.53	-0.023	-0.36	-0.328	-3.77
Hog price, $r^j$	0.346	5.26	0.629	10.69	0.365	5.40
Homogeneity restrictions	Wald test	<i>p</i> -value				
$H_0 : \beta^{QC,US} = \beta^{MB,US} = \beta^{ON,US}$	0.97	0.617				
$H_0 : \beta^{QC,JP} = \beta^{MB,JP} = \beta^{ON,JP}$	15.06	0.001				
$H_0 : \lambda^{QC} = \lambda^{MB} = \lambda^{ON}$	9.92	0.001				
$H_0 : \phi^{QC} = \phi^{MB} = \phi^{ON}$	19.25	0.000				
$H_0 : \beta^{QC,US} = \beta^{MB,US} = \beta^{ON,US};$ $\beta^{QC,JP} = \beta^{MB,JP} = \beta^{ON,JP};$ $\lambda^{QC} = \lambda^{MB} = \lambda^{ON};$ $\phi^{QC} = \phi^{MB} = \phi^{ON}$	60.01	0.000				
Significance of pre-determined hog supplies	Wald test	<i>p</i> -value	Wald test	<i>p</i> -value	Wald test	<i>p</i> -value
$H_0 : H_0 : \beta^{j,JP} = \lambda^j = 0$	17.11	0.000	2.83	0.242	20.67	0.000

**Table 3. Exchange rate pass-through coefficients in the Japanese market**

Variable	Quebec		Manitoba		Ontario	
	Coefficient	<i>t</i> -statistic	Coefficient	<i>t</i> -statistic	Coefficient	<i>t</i> -statistic
U.S. exchange rate, $e^{US}$	-0.334	-0.49	0.207	0.63	0.883	6.46
Japan exchange rate, $e^{JP}$	-0.426	-1.86	-0.554	-2.87	-0.717	-5.31
Capacity, $Q^j$	0.702	3.14	-0.203	-1.36	-0.183	-1.15
Hog price, $r^j$	0.320	2.54	-0.065	-0.52	0.259	2.71
Homogeneity restrictions	Wald test	<i>p</i> -value				
$H_0 : \beta^{QC,US} = \beta^{MB,US} = \beta^{ON,US}$	6.56	0.038				
$H_0 : \beta^{QC,JP} = \beta^{MB,JP} = \beta^{ON,JP}$	1.45	0.484				
$H_0 : \lambda^{QC} = \lambda^{MB} = \lambda^{ON}$	12.91	0.002				
$H_0 : \phi^{QC} = \phi^{MB} = \phi^{ON}$	5.75	0.056				
$H_0 : \beta^{QC,US} = \beta^{MB,US} = \beta^{ON,US};$ $\beta^{QC,JP} = \beta^{MB,JP} = \beta^{ON,JP};$ $\lambda^{QC} = \lambda^{MB} = \lambda^{ON};$ $\phi^{QC} = \phi^{MB} = \phi^{ON}$	28.07	0.000				
Significance of pre-determined hog supplies	Wald test	<i>p</i> -value	Wald test	<i>p</i> -value	Wald test	<i>p</i> -value
$H_0 : H_0 : \beta^{j,US} = \lambda^j = 0$	6.74	0.034	1.01	0.605	48.86	0.000

**Table 4. Exchange rate pass-through coefficients for standard specifications that do not account for pre-determined supplies**

Variable	Quebec		Manitoba		Ontario	
	Coefficient	<i>t</i> -statistic	Coefficient	<i>t</i> -statistic	Coefficient	<i>t</i> -statistic
U.S.	-0.55	-4.98	-0.41	-6.94	-0.64	-7.25
Japan	-0.51	-1.15	-0.30	-1.90	-0.22	-0.98

## Endnotes

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<sup>1</sup> Alternatively, one could assume that the domestic firm is a price follower in that  $\bar{p}^i$  represents the price announcements of firms in market  $i$ . We later solve equilibrium prices as function of competitors' prices (or anticipations about these prices) to be consistent with the empirical specification generally employed in the literature (*e.g.*, Knetter, 1989, 1993).

<sup>2</sup> For simplicity, it was assumed that all hogs are either sold through contracts or through the spot market. In reality, there is always the possibility that hog producers sell their production through both channels as exemplified in Quebec. This case however clutters the analytical model and has no bearing on the empirical model presented in the next section.

<sup>3</sup> Note that Moon and Perron (2004) propose a similar estimator but their asymptotic theory relies on the hypothesis that all regressors in the cointegrating regression equations be different. In case the latter assumption is not met, they propose a minimum distance estimator.

<sup>4</sup> The Breusch-Pagan statistic used to test the null hypothesis of diagonal variance-covariance matrix (Greene, 1998) yields a *p-value* of 0.000. Thus, there are significant efficiency gains in estimating the ERPT equation for each province jointly.

<sup>5</sup> The Breusch-Pagan statistic associated with a diagonal variance-covariance matrix yields a *p-value* of 0.000.