Investigating Non-Linearities in the Relationship Between Real Exchange Rate Volatility and Agricultural Trade

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April 2004

Abstract: The article analyzes production and marketing lags in agri-food supply chains that force competitive producers and processors to commit to output targets before prices and exchange rates are realized. We show that export markets act as put options for exporters and an increase in the volatility of the real exchange rate will generally increase exports. Relaxing the assumptions about the real exchange rate distribution and risk preferences of producers and/or processors can introduce non-linearities in the relationship between exports and real exchange rate volatility. This relationship is investigated using the flexible non-linear inference framework of Hamilton (2001). Bilateral export equations for Canadian pork exports to the U.S. and Japan are specified. The empirical model shows that real exchange rate volatility has statistically significant non-linear effects on aggregate pork exports. Moreover, bilateral pork exports are less sensitive to country-specific variables than aggregate volatility in the real exchange rate.

Keywords: Real exchange rate volatility, non-linear flexible inference, production lags, pork exports.

J.E.L. Classification: F31, Q17, C32

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1 - Introduction

Despite the widespread view that increases in the volatility of financial variables have significant impacts on trade, empirical evidence is mixed (McKenzie, 1999). A number of theoretical models have been proposed to explain the impacts of exchange rate volatility on trade. Under the assumption of risk neutrality, a common belief is that firms' behavior is not affected by uncertainty and increases in volatility do not impact trade flows. When firms are risk-averse, exchange rate volatility generally has negative impacts on trade flows (McKenzie, 1999). A large body of empirical studies found evidence of significant negative impacts of exchange rate volatility on bilateral or aggregate trade flows (*e.g.,* Cushman, 1983; Kenen and Rodrik 1986; Chowdhury, 1993; Arize *et al.*, 2000; Sauer and Bohara, 2001; and Cho *et al.*, 2002). On the other hand, some authors have argued that it is theoretically possible to find evidence of positive correlation between exchange rate volatility and exports. Giovaninni (1988) argued that trade opportunities are similar to a put option held by firms. An increase in the volatility of the exchange rate raises the payoff of the option which induces a proportional increase in trade. In a general equilibrium setting, Bacchetta and Van Wincoop (2000) showed that exchange rate volatility can lead to a larger volume of trade for a large class of preferences and monetary policy rules. On the empirical side, Hooper and Kohlhagen (1978) and Asseery and Peel (1991) have uncovered evidence of a positive correlation between exchange rate uncertainty and trade.

The objectives of this paper are twofold. First, a theoretical trade model that accounts for production and marketing lags in agricultural supply chains is developed to analyze the effect of exchange rate volatility on trade. Production of primary agricultural goods and processed food products is inherently risky since it is characterized by biological and marketing lags that force agricultural producers and processors to commit to output targets before prices and exchange rates are realized. These lags are especially lengthy in livestock and grain industries whose production decisions precede marketing decisions by several months.[1]

The second objective is to analyze empirically the impacts of real exchange rate volatility on aggregate Canadian pork exports and bilateral trade flows to the U.S. and Japan. The estimation procedure considers potential non-linearities between pork trade flows and real exchange rate volatility. The theoretical ambiguity regarding the effect of exchange rate volatility on trade flows justifies the search for empirical evidence. Noticing that this ambiguity has not been resolved neither theoretically nor empirically, Baum *et al.* (2004) use aggregate export data from 13 developed countries to investigate if non-linearities in the relationship between exports and volatility may explain the existence of so many contradicting empirical results in the literature. They consider a rather stringent form of non-linearity by modeling the interaction between exchange rate volatility and the volatility of economic activity in the importing country. The current paper investigates non-linearities between pork exports and real exchange rate volatility using Hamilton's (2001, 2003) flexible non-linear estimation procedure. The estimation allows for unconstrained forms of non-linearity and thus provides a more powerful empirical test of non-linearity than in Baum *et al.* (2004).

Different provinces in Canada use different hog marketing institutions and some even use several institutions concurrently (Larue *et al.,* 2002). It is thus important to account for these particularities and estimate a disaggregated model that will account for these institutional features. Under general conditions, the theoretical model of hog marketing institutions reveals that the impact of real exchange rate volatility on pork exports can not be determined a priori. Export markets act as put options for Canadian pork meat exporters. Under risk neutrality, an increase in the volatility of export prices denominated in Canadian currency (or equivalently, the real exchange rate) increases total pork supplies and increases (expected) exports. Relaxing the theoretical assumptions about the exchange rate distribution and risk preferences of producers and/or processors introduces non-linearities in exports that are difficult to track theoretically. The empirical results show significant non-linearities in the relationship between exports and real exchange rate volatility. Gervais and Larue (2002) analyzed bilateral exports from the province of Quebec in Canada to the U.S. using the auto-regressive distributed lag framework of Pesaran and Shin (1999). Their linear model found that enhanced exchange rate volatility decreased exports of processed pork from Quebec to the U.S. when the exchange rate volatility is measured over a long-term horizon. On the other hand, our empirical model shows that there are values of the real exchange rate volatility that will increase exports.

The rest of the paper is structured as follows. The next section introduces the theoretical model to characterize the dynamic nature of hog marketing mechanisms and to highlight how marketing lags influence processors and producers' output decisions. The third section begins by describing the pattern of bilateral pork exports and exchange rate volatility. This is followed by the presentation of the empirical model and the results of the estimation. The final section offers concluding remarks and suggests further avenues of research.

2 – The Theoretical Model

We developed an analytical framework that explains the relationship between pork exports and real exchange rates. The model accounts for the dynamic nature of the hog/pork supply chain and the vertical marketing structure between hog producers and pork processors in a two-stage game. For analytical convenience, it is assumed that there is a single processor in the domestic market. It has monopoly power on the domestic market but its exports have a negligible effect on the terms of trade (*i.e.,* the small country assumption). The assumption of monopoly behavior is reasonable given the significant literature documenting the increasing concentration at the processing level.[2] While the current model can be applied to different agricultural commodities, its assumptions are mainly based on the stylized facts pertaining to the Quebec hog/pork industry.

In the first stage of the game, the processor must commit to a price paid to hog producers. Given the hog producers' supply, the price commitment determines how many live animals will be processed domestically in the second period. At the beginning of the second period, uncertainty about the foreign pork price is resolved and the processor market the hogs raised in the past period. This simple structure mirrors rather well the marketing institutions in the Quebec hog/pork supply chain. Since 1989, a single-desk selling board is responsible for marketing domestically produced hogs to processors. Although marketing institutions have constantly evolved in Quebec, the cornerstone of the marketing system remains a pre-attribution supply mechanism. In a few words, pre-attribution implies that a large percentage of total hog supplies are assigned to processors based on their historical share of pork sales at a predetermined price. This price has historically been set in relation to the U.S. price.[3]

As usual, the strategic game is solved by backward induction. Denote the total output (capacity) resulting from the 1st stage of the game by q^T. Consider that there is a single export market and a single processed pork commodity. Domestic and foreign pork prices are denoted by p^d and p^x respectively and domestic and foreign pork quantities supplied by the processor in the 2nd period are respectively q^d and q^x such that $q^T = q^d + q^x$. All prices are denominated in Canadian dollars and thus p^x is the foreign price multiplied by the value of the Canadian dollar per unit of the foreign currency. The processor faces the inverse demand function $p^d\left(q^d\right) = 1 - q^d$ on the domestic market; but is a price taker on the foreign market.

It is assumed that the export price is composed of a systematic component $\left(\overline{p}^x\right)$ and a random component ε such that $p^x = \overline{p}^x + \lambda\varepsilon$; with $\lambda > 0$. Uncertainty in the model is captured by the random term ε. Furthermore, it is assumed that ε follows a uniform distribution on the interval $\left[\theta, \eta\right]$ with density $\dfrac{1}{\eta - \theta}$. Hence, if $\eta = -\theta$, the unconditional mean of the export price is $\overline{p}^x$ and the parameter λ is a mean preserving spread (Rothschild and Stiglitz, 1970). At the beginning of the second period, the processor has full knowledge of the foreign price and there is no uncertainty. The processor's profit is defined as:

$$\pi = \left(1 - q^d\right)q^d + \left(\overline{p}^x + \lambda\varepsilon\right)\left(q^T - q^d\right) - r^d q^T, \tag{1}$$

where r^d is the domestic price of live hogs. Without loss of generality, it assumed that average processing costs are constant and are normalized to zero for simplicity.

Sales of the processor in each market are determined by maximizing (1) subject to the first period capacity constraint: $q^d + q^x \le q^T$. Given that the first-stage cost to invest in a capacity is sunk, then it follows that the processor maximizes revenue by selling in either or both markets as:

$$1 - 2q^d \underset{>}{\overset{<}{=}} \bar{p}^x + \lambda\varepsilon \tag{2}$$

There exits three distinct possibilities emanating from (2): *i*) if $\varepsilon < \left(1 - \bar{p}^x - 2q^T\right)/\lambda$, then exports will be zero $\left(q^x = 0, q^d = q^T\right)$ and the processor's profit is $\pi = \left(1 - q^T\right)q^T - r^d q^T$; *ii*) if the export price realization is such that $\left(1 - \bar{p}^x - 2q^T\right)/\lambda < \varepsilon < \left(1 - \bar{p}^x\right)/\lambda$, both exports and domestic sales will be positive $\left(q^x > 0, q^d > 0\right)$ and the processor's profit is: $\pi = \left(1 - q^d\right)q^d + \left(\bar{p}^x + \lambda\varepsilon\right)\left(q^T - q^d\right) - r^d q^T$; and finally *iii*) the export price realization can be so high $\left(\left(1 - \bar{p}^x\right)/\lambda < \varepsilon\right)$ that it may be more profitable for the monopolist not to serve the domestic market $\left(q^x = q^T, q^d = 0\right)$. In the latter case, the processor's profit function is $\pi = \left(\bar{p}^x + \lambda\varepsilon\right)q^T - r^d q^T$.

In the first stage of the game, hog production decisions are made. The 2nd period realization of the real exchange rate (equivalently the export price in Canadian dollars) is not known; but all agents know the distribution of the random variable. For further reference, it is useful to define the following bounds on the exchange rate.

Assumption I: Define the minimum random shock on the exchange rate that guarantees that exports will be positive in equilibrium as $\theta^e \equiv \left(1 - \bar{p}^x - 2q^T\right)/\lambda > \theta$. Similarly, we define the maximum random shock on the exchange rate that guarantees domestic sales will be positive in equilibrium as $\eta^d \equiv \left(1-\overline{p}^x\right)/\lambda < \eta$.

The processor is assumed to be risk-neutral and expected profits are computed by substituting for the decision rule of domestic sales in (1):

$$\begin{aligned} E[\pi] = & \int_{\theta}^{\theta^e} \left(1-q^T\right) q^T \frac{1}{\eta-\theta} d\varepsilon \\ & + \int_{\theta^e}^{\eta^d} \left(\left(\frac{1+\overline{p}^x + \lambda\varepsilon}{2} \right) \frac{1-\overline{p}^x - \lambda\varepsilon}{2} + \left(\overline{p}^x + \lambda\varepsilon\right) \left(q^T - \frac{1-\overline{p}^x - \lambda\varepsilon}{2} \right) \right) \frac{1}{\eta-\theta} d\varepsilon \\ & + \int_{\eta^d}^{\eta} \left(\overline{p}^x + \lambda\varepsilon\right) q^T \frac{1}{\eta-\theta} d\varepsilon - r^d q^T \end{aligned} \tag{3}$$

As mentioned earlier, hog marketing institutions play an important role in determining output capacity of processors. The parameters of the pre-attribution system in Quebec are negotiated on an irregular basis between the processor and the producers' marketing board. The negotiation process has similarities with a bilateral monopoly framework given that there is a single buyer of live animals and sellers are represented by a marketing board. However, the marketing board does not have supply management power and thus does not control supply. Hence, it is assumed that the processor commit to a price in the first period to target a specific level of total hog production supplied by perfectly competitive hog producers. The profit of a representative hog producer is:

$$\pi^{prod} = r^d q^T - \mu\left(q^T\right) \tag{4}$$

where $\mu\left(q^T\right)$ is a twice-differentiable cost function that satisfies $\mu', \mu'' > 0$.

The first-order condition for profit maximization determines total hog supply: $q^T = \phi\left(r^d\right)$. The processor must commit to a price in the first period that determines its supply of live animals to market in the second period. Although the model is cast in terms of two distinct time periods, the reality is that hog production is a lengthy process that can involve up to 10 months between the moments sows are inseminated and the time piglets attain the ready-to-market hog weight. For further reference, define the sum of all three integrals in (3) as total revenue of the processor, $RT\left(q^{T}\right)$. Total cost of the processor is: $CT = r^{d}q^{T}$. The processor's capacity is determined by the first-order condition with respect to the hog price commitment:[4]

$$\partial E[\pi]/\partial r^{d} = \frac{\partial RT}{\partial q^{T}}\frac{\partial q^{T}}{\partial r^{d}} - \left(q^{T} + r^{d}\frac{\partial q^{T}}{\partial r^{d}}\right) = 0 \tag{5}$$

Equation (5) determines the hog price commitment of the processor which in turn determines total capacity in the industry when using the producers' hog supply:[5,6]

$$q^{T*} = \chi\left(\overline{p}^{x}, \lambda, \eta, \theta\right) \tag{6}$$

Exports of pork products are defined by: $q^{x} = q^{T*} - q^{d}$ with $q^{d} = 0.5\left(1 - p^{x}\right)$. Exports are thus directly linked to output capacity of the industry.

Proposition 1: Exports are positively (negatively) related to the parameter λ such that $dq^{x}/d\lambda > (<) 0$ *if and only if* $\eta > (<) -\theta^{e}$. Moreover, exports are increasing in the bounds of the distribution of the random shock; hence $dq^{x}/d\eta > 0; dq^{x}/d\theta > 0$.

Proof: See the Appendix.

Proposition 1 relates exports of pork products to the parameter λ. Assumptions about the distribution of the export price implies that $E\left[p^{x}\right] = \overline{p}^{x} + 0.5\lambda\left(\eta + \theta\right)$ and

$\text{var}\left[p^x\right] = \lambda^2 (\eta - \theta)^2 / 12$. A change in the parameter λ has two effects. It increases the volatility of the export price but also potentially increases the expected export price depending on the parameters of the distribution. For example, when the distribution of the random shock is symmetric around zero, an increase in λ can be interpreted as an increase in the mean preserving spread of the export price. In other words, an increase in λ increases the variance of the export price but leaves unchanged the first moment of the distribution. Under this assumption $(i.e., \eta = -\theta)$, an increase in the mean preserving spread is a sufficient condition to increase exports from an *ex-ante* perspective. Note that realized exports are function of the realized export price and export levels from an *ex-post* perspective could be low. Proposition 1 is very general in that it allows the distribution of the export price to be non-symmetric. Suppose that there is a positive bias in the distribution of the export price such that $\eta > -\theta$. Proposition 1 states that *ex-ante* exports will increase following an increase in the parameter λ given that $\eta \geq -\theta > -\theta^e$.

Alternatively, let us assume that there is a negative bias in the export price such that $\eta < -\theta$. Two cases emerge. First, an increase in λ will increase exports if the bias is small such that $-\theta^e < \eta < -\theta$. The intuition is that the increase in λ increases volatility but decreases the expected average price by a small proportion. On the other hand, if there is a sufficiently large negative bias such that $\eta < -\theta^e < -\theta$, an increase in λ increases the volatility of exchange rate whose positive effect on exports is more than offset by a significant decrease in $E\left[p^x\right]$. It results in an anticipated decrease in exports.

The fact that increases in the volatility of the real exchange rate can boost exports, even when agents are risk-neutral, rests on the presence of production and marketing lags.

The impact of volatility critically hinges on the variable θ^e. As mentioned earlier, this variable establishes a lower bound on the random shock such that it is still profitable to export for the processor. Hence, for (*ex-ante*) exports to be negatively affected by an increase in λ, it must be that the upper bound on the exchange rate is lower in absolute value than the critical bound that allows positive exports. Table 1 provides a simple numerical example to illustrate Proposition 1. When $\eta = -\theta = 0.6$, exports are positively related to λ. Conversely, if the higher bound of the distribution of the real exchange rate is below the absolute value of the threshold $\left(\theta^e\right)$ that guarantees positive exports, total capacity will be decreasing in the value of λ and so will *ex-ante* exports.

Table 1. Numerical example

	Value of η				
	0.3	0.4	0.5	0.6	0.7
θ^e	-0.377	-0.417	-0.459	-0.504	-0.551
$dq^x/d\lambda$	-0.012	-0.003	0.008	0.020	0.035

Parameter values used in the numerical example are: $\bar{p}^x = 0.8$, $\theta = -0.6$, $\lambda = 1.2$ and $\mu\left(q^T\right) = 0.5\left(q^T\right)^2$.

The solution defined in (6) yields the optimal capacity choice of processors: $q^T = q^T\left(\gamma\left(\tilde{p}^x\right);\boldsymbol{\beta}\right)$; where $\gamma\left(\tilde{p}^x\right)$ is a function mapping the different moments of the distribution of the real exchange rate and $\boldsymbol{\beta}$ is a vector representing all other exogenous variables of the model. Substituting the optimal capacity choice of producers in the first-order condition defined in (2) yields respectively the equilibrium exports and domestic sales $q^{d*}\left(p^x,\gamma\left(\tilde{p}^x\right);\boldsymbol{\beta}\right)$ and $q^{x*}\left(p^x,\gamma\left(\tilde{p}^x\right);\boldsymbol{\beta}\right)$. It is important to note that export and domestic sales are both function of the realized real exchange rate and also of the different moments of the real exchange rate distribution.

3 – The Empirical Model

The theoretical framework underlines two key factors conditioning export decisions of processors. First, even though a processor is risk-neutral, exports are function of the uncertainty surrounding the exchange rate and export price because of lags in the marketing of agricultural products. Second, relaxing assumptions about the functional form of consumers' demand, risk preferences of processors, and the producers' technology can generate significant non-linearities between exports and exchange rate volatility.

Figure 1 illustrates total monthly pork exports from Quebec along with exports to the two most important destinations (U.S. and Japan) for the period starting January 1989 and ending December 2002. The U.S. represents the most important destination for Quebec pork exports. Exports to Japan and the U.S. averaged more than 72% of all exports over the sample period considered. Quebec exports have been more diversified in the later years of the sample as the two most important destinations became relatively less important. Figure 2 presents the evolution of monthly export unit values in Canadian dollars between January 1989 and December 2002 for each destination. Unit values for Japan are significantly higher than for the U.S. as the product mix of exports is significantly different between the two destinations.

The theoretical framework provides the foundation of the empirical export equation:

$$q_t^{x*} = f\left(\gamma\left(\tilde{p}_t^x\right)\right) \tag{7}$$

where the function $\gamma(\cdot)$ subsumes the different moments of the distribution of the effective world price of Quebec exports. It should be emphasized that the present analysis focuses on the distribution of the real exchange rate defined as the export price denominated in Canadian currency. There is no consensus in the literature as to whether one should study the impacts of real exchange rate uncertainty or nominal exchange rate uncertainty (Mackenzie, 1999). Most would agree that the choice of concept depends upon the characteristics of the market being investigated. In the present case, real exchange rate defined as the nominal exchange rate multiplied by the ratio of export to domestic prices is not pertinent because processors are assumed to have market power on the domestic market; as such the domestic price is endogenous to the decisions of processing firms.

Mackenzie (1999) surveys the various indicators used in the literature to measure volatility of the real exchange rate. We define volatility as a moving average of the standard deviation of the export price:[7] $V_t = \left[1/m \sum_{i=1}^{m} \left(e_{t+i-1} p_{t+i-1}^{x} - e_{t+i-2} p_{t+i-1}^{x} \right)^2 \right]^{1/2}$. Various values of the parameter m were tested. Figure 3 presents the volatility measure of the real exchange rate at the aggregate level and for the two destinations when $m = 12$. Given the relative importance of U.S. exports, it is not surprising that the volatility measure of the U.S. real exchange rate follows closely the volatility of the aggregate exchange rate. There are however, significant differences between the two measures mainly due to sudden surges in the volatility of the effective export price in Japan. In order to better gauge the robustness of our volatility measure to the choice of the parameter m, different volatility measures were also computed $(m = 3, 6)$, but they produced similar qualitative results although the measure based on the longer lag generally yields higher estimates of volatility. In what follows, the parameter *m* is set to 12 throughout.

As it is usually the case with monthly time series, the degree of integration of 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 of the level of the series, a constant, a time trend and, if needed, lagged first differences of the dependent variables to make the residuals white noise:

$$\Delta y_t = \alpha + \beta t + \rho y_{t-1} + \sum_{j=1}^{w} \gamma_j \Delta y_{t-j} + \varepsilon_t \qquad (8)$$

The ADF test involves testing whether ρ differs significantly from zero. Failure to reject the null hypothesis of the ADF test indicates that the variables are non-stationary.

The ADF test was implemented on the logarithmic transformation of the real exchange rate, export sales and the volatility of real exchange rate. The results are reported in the second column of Table 1. The first column indicates whether a time trend (T) or no time trend (NT) were used in (8). Following Hall's (1994) recommendations, we used the SBC information criterion to select the lag length in (8) because it tends to make the ADF test more powerful in small samples than the AIC criterion. The null hypothesis of unit root is not rejected for exports to the U.S., the Japanese real exchange rate, and the 12-month volatility of the real exchange rate in the Japanese market.

To assess the reliability of the ADF test, the stationarity test developed by Kwiatkowski *et al.* [hereafter referred to as KPSS, (1992)] was performed on all series. The KPSS testing procedure complements standard unit root testing because its null hypothesis is stationarity. The KPSS test involves estimating the equation:

$$y_t = \delta t + \zeta_t + \varepsilon_t\,;\;\; \zeta_t = \zeta_{t-1} + u_t\,;\;\; u_t \sim iid\left(0, \sigma_u^2\right) \qquad (9)$$

The null hypothesis of trend stationarity can be ascertained by testing $\sigma_u^2 = 0$. Testing the null of level stationarity instead of trend stationarity involves regressing the series on a constant instead of trend variable. The KPSS test relies on the Bartlett kernel with a bandwidth for the spectral window selected with the formula: $l = trunc\left\{4\left(0.01T\right)^{0.25}\right\}$; where *T* is the number of observations in the sample. The third column of table 1 confirms that the null hypothesis of stationarity is rejected for a majority of the variables. In particular, all export quantities and price variables are identified as non-stationary processes. Unfortunately the ADF and KPSS tests yield conflicting evidence; an outcome documented in Maddala and Kim (1998). This is why Carrion-I-Silvestre *et al.* (2001) argue that simultaneous testing of the null hypotheses of stationarity and unit root should not be conducted using standard marginal critical values for each test. They implemented a Confirmatory Data Analysis (CDA) method by computing critical values for the joint confirmation hypothesis of a unit root. They show that using their set of critical values brings about significant improvements in the reliability of the test results when compared to marginal critical values if the data generation process is integrated of order one. The CDA shows that real exchange rate in the U.S. and Japanese markets are integrated of order one as well as exports to the U.S. market. The hypothesis of a unit root is also jointly confirmed by the two tests for the 12-month volatility measures of real exchange rate of aggregate exports and exports to Japan. Clearly, some series are non-stationary and the empirical methodology will need to account for that.

As mentioned previously, the pork export equation can exhibit significant non-linearities in the various moment of the distribution of the real exchange rate. To account for these potential non-linearities, the flexible non-linear inference framework of Hamilton (2001, 2003) is applied. Hamilton's approach begins by estimating a nonlinear regression model of the form:

$$x_t^* = \mu(\mathbf{z}_t) + \upsilon_t \tag{10}$$

where υ_t is a random error term distributed normally with mean zero and variance σ^2. The function $\mu(\mathbf{z}_t)$ is unknown and can accommodate non-linearities in the vector of independent variables, $\mathbf{z}_t$ of dimension $T \times k$. The empirical strategy is to view this function as the outcome of random fields.[8] For a given non-stochastic vector $\mathbf{z}$, the functions $\mu(\mathbf{z})$ is assumed to be normally distributed with mean $\gamma_0 + \boldsymbol{\gamma}_1\mathbf{z}$ and variance λ^2. If the variance is zero, equation (10) reduces to $x_t^* = \gamma_0 + \boldsymbol{\gamma}_1\mathbf{z}_t + \upsilon_t$; which is identical to the standard linear regression framework. However, when λ is large, (10) can substantially deviate from a linear regression model. A specification search is conducted over parameters that characterize the variability of the function $\mu(\mathbf{z})$. Hamilton (2001) shows that this approach is equivalent to specifying the correlation between two random realizations, $\mathbf{z}_1$ and $\mathbf{z}_2$. The empirical framework assumes that these two realizations are uncorrelated if they are sufficiently far apart. Specifically, the correlation is zero when $0.5\left(\sum_{j=1}^{k} g_j^2 \left(z_{j1} - z_{j2}\right)^2\right)^{0.5} > 1$; where the parameters g_j govern the variability of the nonlinear function as $\mathbf{z}_j$ vary. When the previous inequality is not satisfied, the correlation differs from zero and its exact form is given in Hamilton (2001, p. 542). Equation (10) can be rewritten as:

$$x_t^* = \gamma_0 + \boldsymbol{\gamma}_1\mathbf{z}_t + \lambda m(\mathbf{z}_t) + \upsilon_t \tag{11}$$

where $m(\cdot)$ is a stochastic process that characterizes the conditional expectation $\mu(\mathbf{z}_t)$. The function has a mean zero and unit variance. The parameters which need to be estimated are the linear regression coefficients $(\gamma_0, \boldsymbol{\gamma}_1)$, the parameter indicating the prevalence of a non-linear component (λ), the variance of the error term (σ^2) and the *k* parameters governing the non-linearities $(\mathbf{g})$. Given that the error term υ_t and the random field $\mu(\mathbf{z}_t)$ have finite variance, the dependent variable in (11) must be a stationary time series. Hence, it is important to recall that exports to the U.S. market did not satisfy this property, but that the null hypothesis of a unit root was rejected for aggregate exports and exports to Japan.

In the first stage of the empirical application, the impact of the real exchange rate uncertainty on aggregate pork exports is estimated. It is assumed that exports in time *t* are function of the realized real exchange rate in period *t*, the lagged real exchange rate and lagged volatility measures of real exchange rate. Lagged values of the exchange rate and volatility are proxies for the expected exchange rate and variance of the exchange rate respectively. As shown in the theoretical model, these latter variables are important determinants of the first stage capacity and thus determinants of exports. Hog production is characterized by production lags of ten months between the time sows are inseminated and pork meat is marketed. Hence, the selection of the appropriate lag structure is challenging considering the need to specify a parsimonious empirical model. A number of different lag specifications were experimented with; but a ten month lag was found to be adequate.

Following Hamilton (2003), the equation in (11) is rewritten as:

$$x_t^* = \gamma_0 + \boldsymbol{\gamma}_1 \mathbf{z}_t + \sigma \cdot \omega \cdot m(\mathbf{z}_t) + \sigma \varepsilon_t \qquad (12)$$

where the innovation υ_t is written as σ times ε_t and the parameter λ is written as $\lambda = \sigma \cdot \omega$ The maximum likelihood estimates of (12) and their standard error (between parentheses) are:

$$
\begin{aligned}
x_t^* = &\underset{(0.89)}{8.32} - \underset{(0.51)}{0.51} p_t^x + \underset{(0.53)}{1.39} p_{t-10}^x - \underset{(0.28)}{0.11} vol_{t-10}^x \\
&+ \underset{(0.02)}{0.27}\left[\underset{(0.23)}{0.91} m\left(\underset{(1.6)}{4.34} p_t^x + \underset{(1.99)}{4.08} p_{t-10}^x + \underset{(0.75)}{5.32} vol_{t-10}^x\right)\right]
\end{aligned}
\tag{13}
$$

A non-negligible advantage of the flexible non-linear framework is that it allows a direct test of the null hypothesis that the true relation in (11) is linear. This amounts to testing whether λ^2 is different from zero with a Lagrange Multiplier (LM) test. In the current application, the null hypothesis of a linear model is rejected since the *p-value* of the LM test is 0.00.

Of the outmost interest is the fact that all coefficients in the linear part of (13) have a relatively large standard error with respect to the coefficient estimate except for the constant of the regression. Moreover, all the parameters in the non-linear component of (13) are positive and significantly different than zero. To establish a point of comparison with the usual empirical applications, we also computed the OLS estimates of the linear component in equation (12). The coefficient estimates and their standard error are:

$$
x_t^* = \underset{(0.85)}{8.03} - \underset{(0.48)}{1.17} p_t^x + \underset{(0.57)}{1.56} p_{t-10}^x - \underset{(0.22)}{0.61} vol_{t-10}^x
\tag{14}
$$

The results in (14) demonstrate that ignoring the potential non-linearity of the export equation can result in severe misspecification issues. The coefficients in (14) are all statistically significant than zero and volatility is negatively correlated with exports. Although the results in (13) confirm that the relationship between exports and volatility is non-linear, it is difficult to determine what the non-linear relationship looks like. Hamilton (2003) suggests fixing all but one of the independent variables to their sample mean to examine the consequences on the conditional mean of $\mu(\mathbf{z})$ in (10) of letting one variable vary. Inference about the behavior of $\mu(\mathbf{z})$ can be conducted using a Bayesian framework. This entails selecting prior distributions for the linear parameters $\{\gamma_0, \boldsymbol{\gamma}_1, \sigma\}$ and non-linear parameters $\{\mathbf{g}, \omega\}$ of the model and generating values of the parameters whose mean and standard deviation are reported in (13) to simulate the posterior distribution of $\mu(\mathbf{z})$ given specific values of $\mathbf{z}$. The rationale to use a Bayesian framework is that the small sample properties of the function $\mu(\mathbf{z})$ are largely unknown. The selection of priors for the parameters is described in Hamilton (2001, pp. 552-553).[9] The posterior distribution of $\mu(\mathbf{z})$ was simulated based on 5,000 draws of the importance sampling distribution described in Hamilton (2001).

Figure 5 plots the predicted value of $\mu\left(\overline{p}_t^x, \overline{p}_{t-10}^x, vol^x\right)$ as a function of vol^x when p_t^x and p_{t-10}^x are set at their mean value. The posterior probability that $\mu(\mathbf{z})$ will fall between the bounds defined by the dotted lines is 95%. Variations in vol^x range from $\pm$ two times the standard deviation around its mean of 0.179. The point estimate of $\mu(\mathbf{z})$ is not monotonic in vol^x. Volatility impacts are strongest when volatility is slightly below its mean value. Lower volatility tends to reduce exports. The evidence suggests that although increases in volatility can potentially increase the expected payoff of exporting activities, there are levels of volatility for which export activities are less attractive. In any case, there are substantial differences between the linear and non-linear models.

The next step is to investigate bilateral export flows. Past studies have provided evidence that destination specific volatility measure of the exchange rate play an important role in determining exports to that market (see for example Baum *et al.,* 2004). Although our theoretical model did not explicitly account for such effects, it showed that global volatility of the real exchange rate (measured by the unit value of aggregate exports) must be an important determinant of bilateral exports due in large part to the fact that marketing lags force exporters to commit to capacity before uncertainty is resolved. However, destination-specific volatility can also play a role in determining bilateral exports. For example, significant differences in the preferences of consumers across importing markets can exist. Pork products are not homogenous and preferences in one country may be biased toward higher quality (more expensive) products than in another country. In that case, destination-specific volatility may play an important role in determining bilateral exports.

In summary, lags of the real exchange rate of aggregate exports and lags of the volatility of real exchange rate should be important determinants of processors' global capacity; and thus exert an indirect influence on bilateral trade flows. The lagged volatility of the real exchange rate to the specific market is also included to test whether the destination-specific volatility measure has an impact on bilateral trade flows. Finally, the current real exchange rate of a specific destination also enters the model specification as it is determines profitability in that market.

The bilateral export equations are specified for the American and Japanese markets. The maximum likelihood estimates of the equations along with their standard errors are:

$$x_t^{US} = 8.59 + 0.22 p_t^{US} + 0.13 p_{t-10}^{x} + 0.07 vol_{t-10}^{US} - 0.07 vol_{t-10}^{x}$$
$$(0.72)\ (0.31) \quad (0.38) \quad (0.20) \quad (0.18)$$
$$+ 0.058\left[4.45 m\left(-5.58 p_t^{US} + 3.77 p_{t-10}^{x} + 3.24 vol_{t-10}^{US} + 4.24 vol_{t-10}^{x}\right)\right] \tag{15}$$
$$(0.029)(2.61) \quad (1.14) \quad (1.85) \quad (0.71) \quad (1.05)$$

$$x_t^{Jap} = 8.07 - 2.17 p_t^{Jap} + 1.26 p_{t-10}^{x} - 0.07 vol_{t-10}^{Jap} - 0.85 vol_{t-10}^{x}$$
$$(1.00)\ (0.27) \quad (0.54) \quad (0.18) \quad (0.28)$$
$$+ 0.092\left[5.20 m\left(6.02 p_t^{Jap} + 3.38 p_{t-10}^{Jap} + 7.67 vol_{t-10}^{Jap} + 5.05 vol_{t-10}^{x}\right)\right] \tag{16}$$
$$(0.091)(5.48) \quad (1.10) \quad (1.69) \quad (1.29) \quad (1.71)$$

The Lagrange multiplier test did not reject the null hypothesis of non-linearity for each bilateral equation. Before interpreting the results in (15) and (16), we computed the OLS estimates of the linear component in (11) and their standard error:

$$x_t^{US} = 8.99 + 0.29 p_t^{US} - 0.35 p_{t-10}^{x} + 0.29 vol_{t-10}^{US} - 0.33 vol_{t-10}^{x} \tag{17}$$
$$(0.77)\ (0.33) \quad (0.53) \quad (0.20) \quad (0.19)$$

$$x_t^{Jap} = 7.04 - 2.22 p_t^{Jap} + 2.19 p_{t-10}^{x} + 0.02 vol_{t-10}^{Jap} - 0.94 vol_{t-10}^{x} \tag{18}$$
$$(1.03)\ (0.24) \quad (0.59) \quad (0.16) \quad (0.27)$$

The coefficients of the bilateral U.S. export equation in (17) are not statistically different than zero at the 90% confidence level expect for the coefficient of the lagged aggregate exchange rate volatility. Lagged real exchange rate volatility measured over aggregate trade flows negatively affects exports to the U.S. while destination specific volatility has no significant impact. A similar story holds for the bilateral exports to Japan. Moreover, the coefficients of the export price (both country specific and aggregate) are strongly significant. Overall, the linear specifications in (17) and (18) suggest that it is global volatility that has a significant and negative impact on bilateral trade flows.[10]

The non-linear specification of the U.S. equation in (15) tells a different story than its linear counterpart in (17). The coefficients of the linear component do not seem to be significantly different than zero while the coefficients of the non-linear component are statistically significant. The coefficients of the linear component in the exports to Japan equation are quite similar to the coefficients in (18). Although the Lagrange multiplier test suggested that there was a significant non-linearity in (16), the estimate of σ and ω are quite close to their standard error. This sheds doubts that there is a significant non-linearity in the relationship between bilateral exports from Quebec to Japan and real exchange rate volatility.

In order to explore further the potential non-linearities reported in (15) and (16), figure 6a presents the marginal impact on Quebec exports to the U.S. market of changes in the lagged volatility of the aggregate exchange rate holding all other independent variables fixed at their mean. There are significant non-linearities in the lagged volatility of the aggregate real exchange rate. Figure 6b plots the value of $\mu\left(\overline{p}_t^{US}, \overline{p}_{t-10}^{x}, \overline{vol}_{t-10}^{x,12}, vol_{t-10}^{US,12}\right)$ as a function of $vol_{t-10}^{US,12}$. Lagged volatility of the real exchange rate in the U.S. market does not appear to be an important determinant of exports to the U.S. Figure 7a and 7b present respectively the marginal impact on Quebec exports to Japan of a change in the lagged volatility of the aggregate real exchange rate and the lagged volatility of the real exchange rate in the Japanese market. It further confirms that destination-specific volatility does not have as strong an effect in the Japanese market as it does in the U.S. market and that aggregate volatility seems to be negatively correlated with exports.

4 - Concluding Remarks

The literature on the impacts of real exchange volatility on exports is voluminous and conflicting results about the relationship between volatility and export flows at the aggregate and disaggregate levels abound. The current paper shows that production and marketing lags in agri-food supply chains can have ambiguous effects on the relationship between real exchange rate volatility and trade. The theoretical model illustrates that, even under simple market and behavioral assumptions, one cannot sign the relationship between exchange rate volatility and exports as it depends in this simple case on the distribution assumption of the real exchange rate. The model explains the relationship between exchange rate volatility and exports in the context of the Quebec hog/pork industry; but it applies to numerous settings in which there are long lags in marketing or production; or more generally when capacity decisions must be made before marketing decisions. If the distribution of the real exchange rate is symmetric, an increase in its variance will cause an increase in exports because the existence of a foreign market acts as a put option for domestic firms. Hence, more volatility increases the expected payoffs of the option *ex-ante*. Introducing non-linear cost and demand functions as well as risk-aversion creates ambiguities that are extremely difficult to resolve theoretically.

Based on the theoretical model, we specified an empirical model that accounts for potential non-linearities and has a flexible approach to inference because of the many possibilities. Hamilton's (2001) framework was used to diagnose whether volatility of the real exchange rate of the Canadian currency has a positive or negative impact on Quebec pork exports. The empirical results strongly reject the hypothesis of linearity in the relationship between exports and volatility. Moreover, it is shown that the relationship seems positive for low levels of volatility but negative for higher levels. Note that the impact of volatility on exports is through the capacity constraint pork processors face in the model. An increase in the volatility of the real exchange rate increases the expected payoff of the processors but this increase must be sufficiently large to induce additional investments in primary input purchases.

Bilateral export equations to the two most important destinations were also estimated. Volatility of the aggregate export price seems to exercise a greater influence on bilateral trade flows than destination-specific measures of real exchange rate volatility. Volatility of the former variable is perhaps more important than the volatility in the latter variable when pre-committing to capacity levels before exporting decisions are made. The empirical model detects non-linearities that are less important when analyzing bilateral trade flows than aggregate exports. Due to features of agricultural commodities such as the jointness in production of different pork products and heterogeneous consumers' preferences across export markets, one would expect destination-specific variables to be more important than what is currently portrayed by the current empirical results. This topic could constitute an interesting avenue to explore in future research.

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Appendix - Proof of Proposition 1

To prove the proposition, we use the fact that total capacity and exports are positively linked. Multiply the first order condition defined in (5) by $\partial r^d / \partial q^T$ to obtain: $\partial E[\pi]/\partial q^T = \dfrac{\partial RT}{\partial q^T} - \left(q^T \dfrac{\partial r^d}{\partial q^T} + r^d \right) = 0$; which states that marginal revenue with respect to capacity equals marginal cost, $\partial E[\pi]/\partial q^T = MR - MC = 0$. Comparative static on the previous equation yields: $\dfrac{dq^T}{d\lambda} = \dfrac{\partial MR}{\partial \lambda} \bigg/ \left(\dfrac{\partial MC}{\partial q^T} - \dfrac{\partial MR}{\partial q^T} \right)$; where $MC \equiv \left(q^T \bigg/ \dfrac{\partial \phi}{\partial r^d} + r^d \right) > 0$. The second order condition for a maximum requires that $\partial MC/\partial q^T - \partial MR/\partial q^T > 0$. It follows that $sign\left(dq^x/d\lambda\right) = sign\left(\partial MR/\partial \lambda\right)$. The derivative of marginal revenue with respect to λ, $\partial MR/\partial \lambda = \dfrac{\eta^2 \lambda^2 - \left(\overline{p}^x - 1 + 2q^T\right)^2}{2(\eta - \theta)\lambda^2}$, is greater than zero if $\left(\eta - \left(\dfrac{\overline{p}^x - 1 + 2q^T}{\lambda} \right) \right) \left(\eta + \left(\dfrac{\overline{p}^x - 1 + 2q^T}{\lambda} \right) \right) > 0$. Since $\eta > \left(1 - \overline{p}^x - 2q^T\right)/\lambda$ when there are positive exports, it follows that $sign\left(\partial MR/\partial \lambda\right) = sign\left(\eta - \left(\overline{p}^x - 1 + 2q^T\right)/\lambda\right)$. Hence, $dq^x/d\lambda > (<) 0$ if and only if $\eta > (<) -\left(1 - \overline{p}^x - 2q^T\right)/\lambda \equiv -\theta^e$.

The second part of the proposition uses: $\dfrac{dq^x}{d\eta} = \dfrac{\partial MR}{\partial \eta} \bigg/ \left(\dfrac{\partial MC}{\partial q^T} - \dfrac{\partial MR}{\partial q^T} \right)$. It follows that $sign\left(\dfrac{dq^x}{d\eta}\right) = sign\left(\dfrac{\partial MR}{\partial \eta}\right)$; with $\dfrac{\partial MR}{\partial \eta} = \dfrac{\left(\overline{p}^x - 1 + 2q^T + \eta\lambda\right)\left(1 - \overline{p}^x - 2q^T + \eta\lambda - 2\theta\lambda\right)}{2(\eta - \theta)\lambda^2}$. Substituting η by its lowest admissible value that guarantees positive domestic sales, $\eta^d = \left(1 - \overline{p}^x\right)/\lambda$, in $\left(\overline{p}^x - 1 + 2q^T + \eta\lambda\right)\left(1 - \overline{p}^x - 2q^T + \eta\lambda - 2\theta\lambda\right)$, a sufficient condition for

$dq^x / d\eta > 0$ is that $1 - q^T > \overline{p}^x + \theta\lambda$. Because $\theta < \left(1 - \overline{p}^x - 2q^T\right) / \lambda$ (see assumption I), it must be that $1 - q^T > \overline{p}^x + \theta\lambda$. This establishes the second part of proposition. Finally, as $\frac{\partial MC}{\partial q^T} - \frac{\partial MR}{\partial q^T} > 0$ and $\frac{\partial MR}{\partial \theta} = \frac{\left(\overline{p}^x - 1 + 2q^T + \eta\lambda\right)^2}{2(\eta - \theta)^2 \lambda} > 0$, it follows that $dq^x / d\theta > 0$.

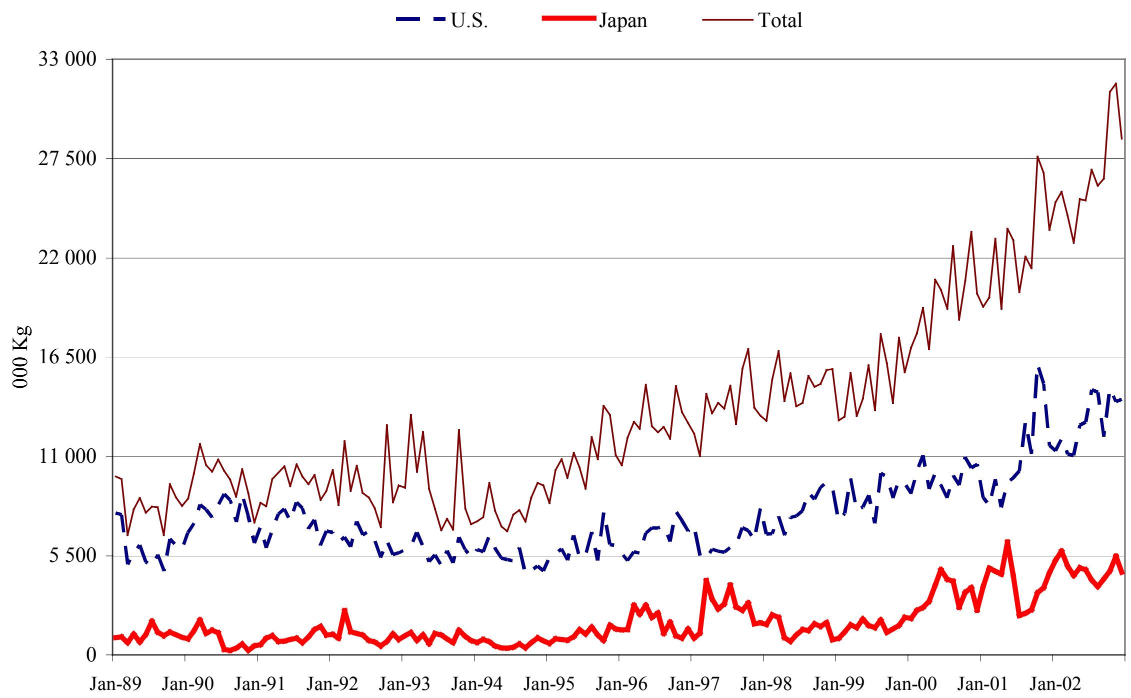

Figure 1. Total monthly pork exports from Quebec and bilateral exports to the U.S. and Japan from January 1989 to December 2002

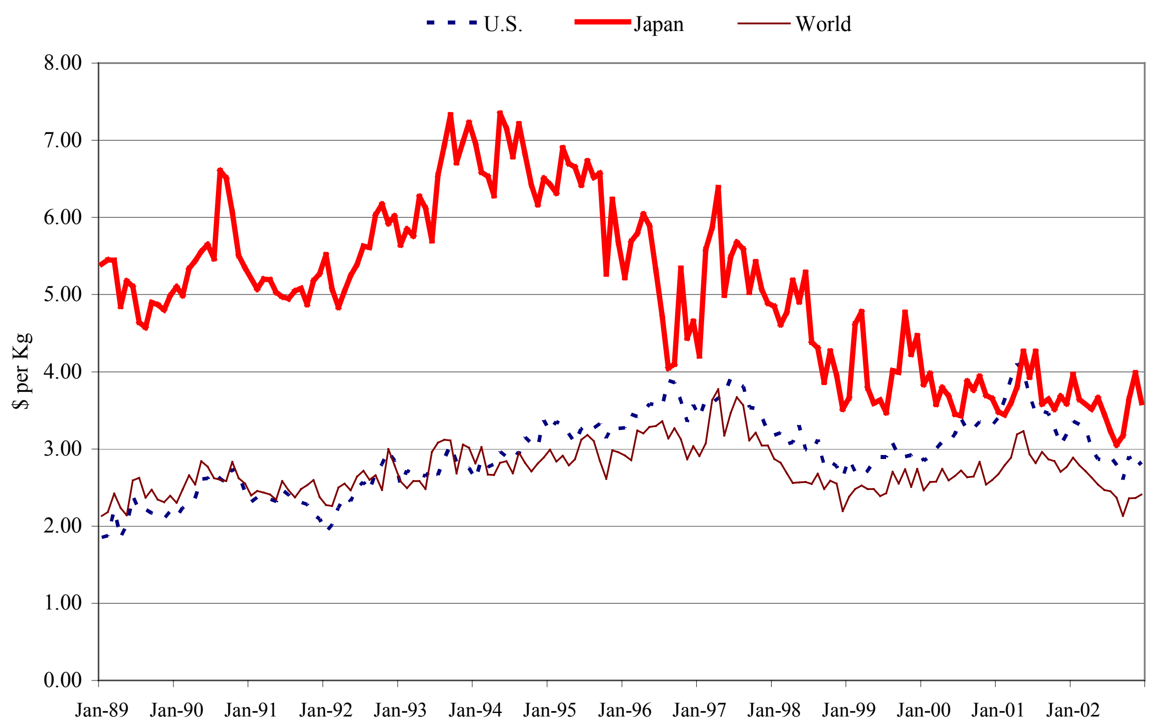

Figure 2. Monthly price (in $Can) of Quebec total pork exports oand U.S. and Japan pork exports from January 1989 to December 2002

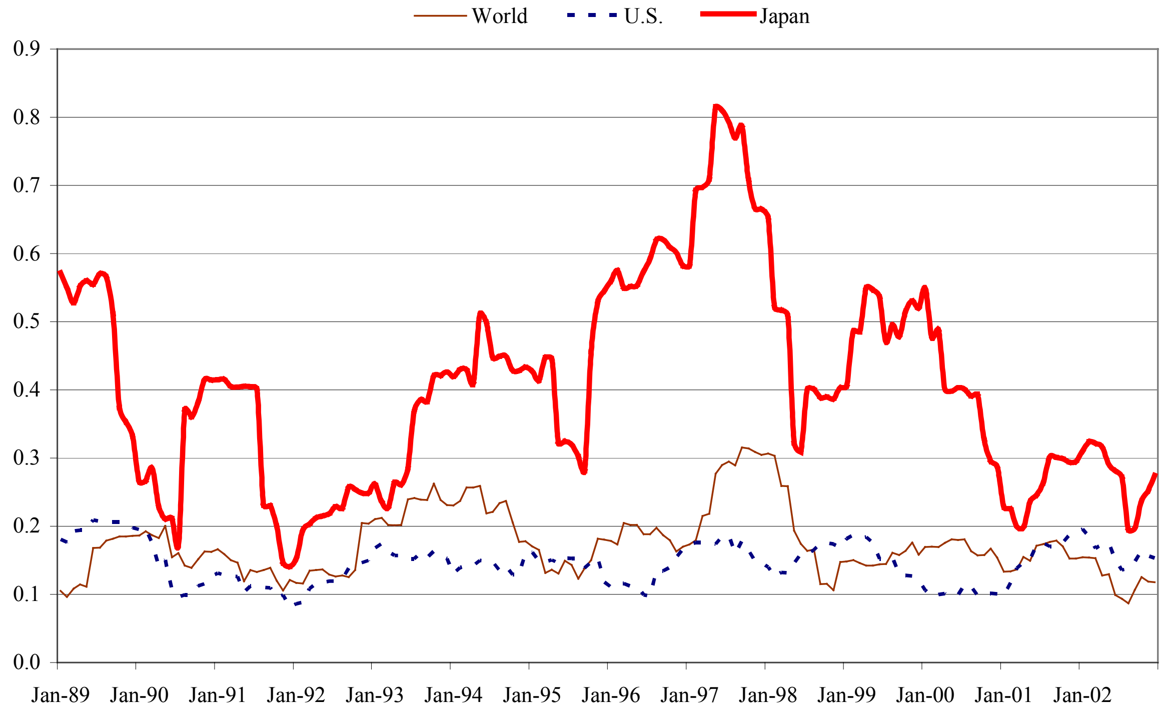

Figure 3. Decomposition of the monthly volatility measure of real export prices to the U.S., Japan and aggregate exports.

Table 1. Unit root testing

Variables	ADF test		KPSS test	Joint confirmation of a unit root
	Lag	Statistic		
Total exports (T)	1	-3.93*	0.48*	No
U.S. exports (T)	4	-1.86	0.57*	Yes
Japan exports (T)	0	-5.19*	0.23*	No
Aggregate real exchange rate (NT)	1	-3.42*	0.53*	No
U.S. real exchange rate (NT)	0	-2.83**	1.66*	Yes
Japan real exchange rate (NT)	3	-0.97	1.78*	Yes
12-month Aggregate Vol (NT)	0	-2.64**	0.25	Yes
U.S. 12-month Vol (NT)	0	-2.61**	0.09	No
Japan 12-month Vol (NT)	0	-2.25	0.36**	Yes

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

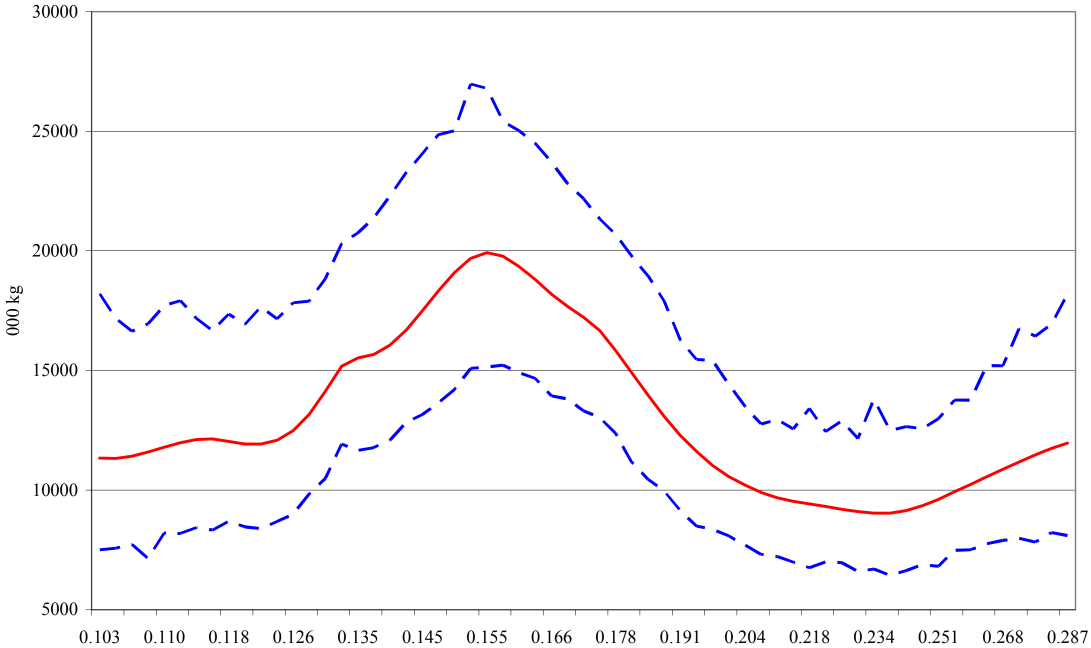

Figure 5. Impact of the real exchange rate volatility on aggregate exports holding all other independent variables at their sample mean.

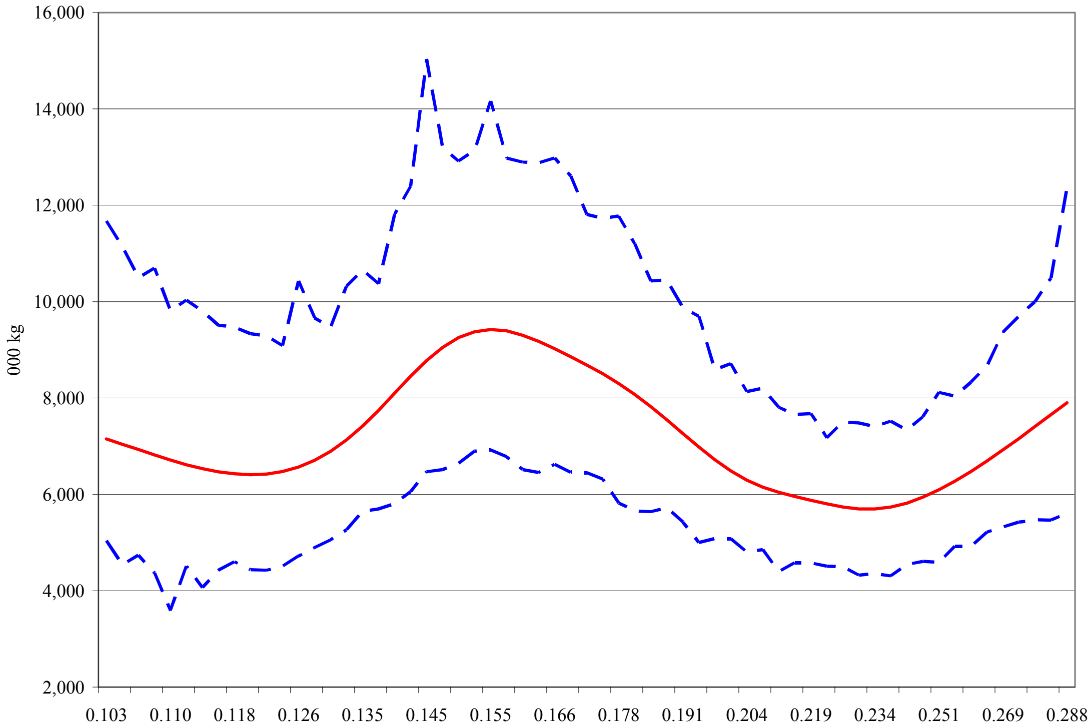

Figure 6a. Impact of the real aggregate exchange rate volatility on exports to the U.S. holding all other independent variables at their sample mean.

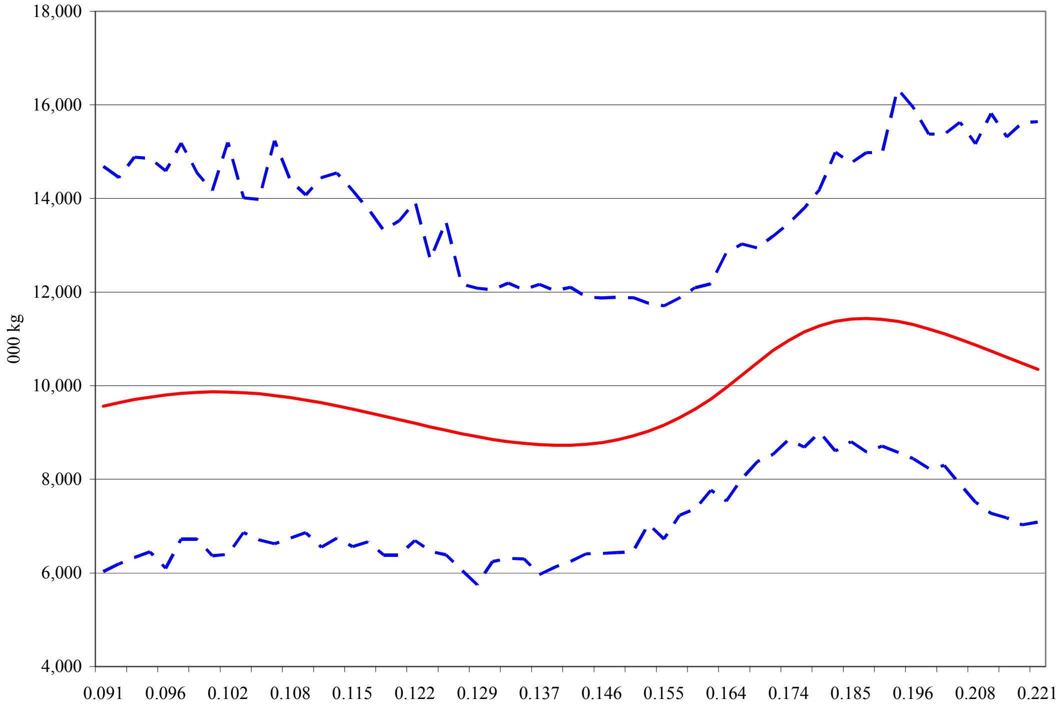

Figure 6b. Impact of the real U.S. exchange rate volatility on exports to the U.S. holding all other independent variables at their sample mean.

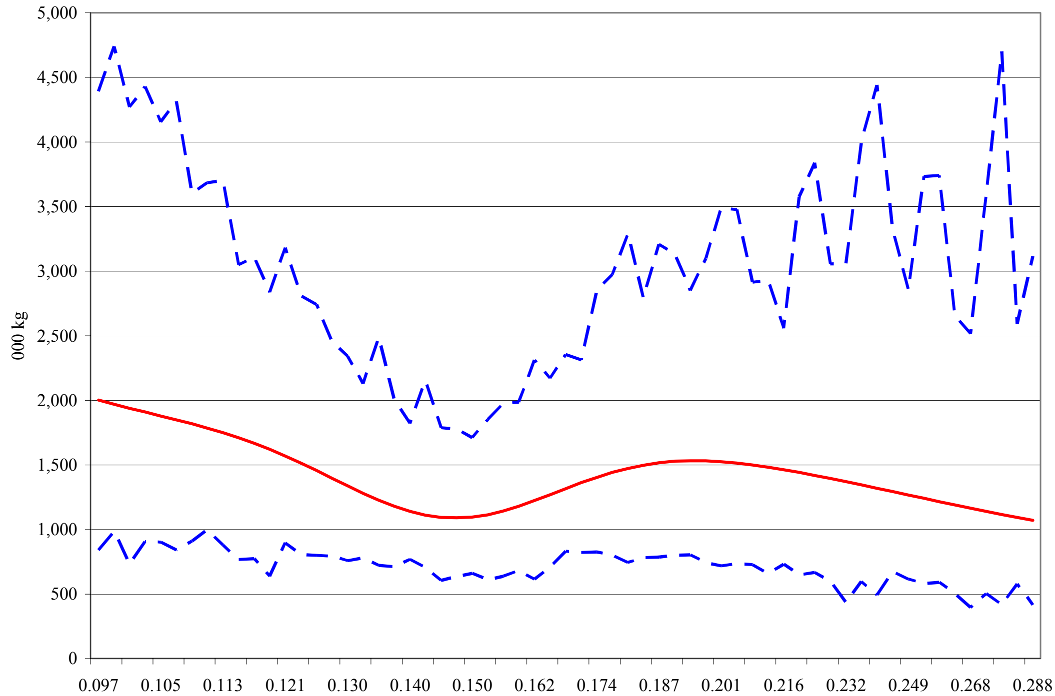

Figure 7a. Impact of the real aggregate exchange rate volatility on exports to Japan holding all other independent variables at their sample mean.

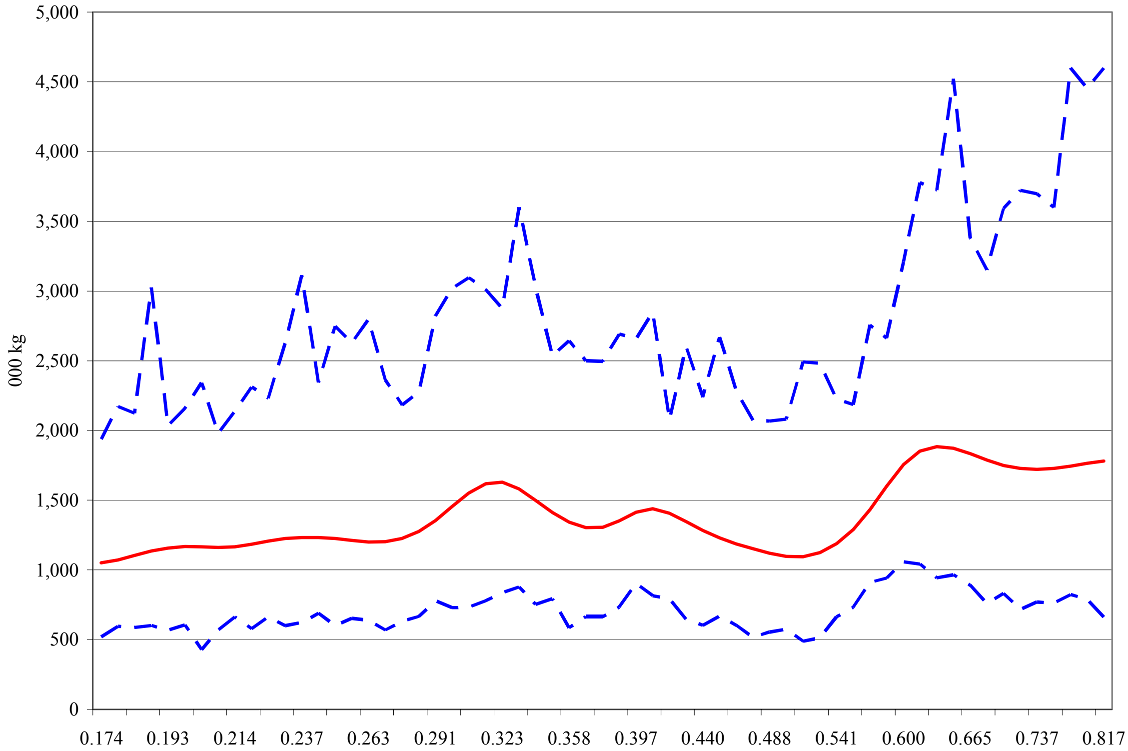

Figure 7b. Impact of the Japanese real exchange rate volatility on exports to Japan holding all other independent variables at their sample mean.

Endnotes

[1] In animal production, the length of time between birth and slaughtering ranges from 6 weeks for chicken to 14-18 months for beef. For hogs, 5 months must elapse to bring a newborn piglet to market, but the full production lag is roughly 10 months when the gestation period is accounted for. In contrast, the production of processed agricultural products usually involves much shorter lags.

[2] For example, Lopez *et al.* (2002) recently documented significant market power in U.S. food processing industries. Schroeter *et al.* (2000) found that there are significant departures from perfect competition in the U.S. beef packing industry. Liu, Sun, and Kaiser (1995) found evidence of market power exercised by U.S. fluid and manufactured dairy processors. Larue, Gervais and Lapan (2004) developed a theoretical model of the hog/pork industry to explain how a hold-up of producers can occur due to imperfect competition at the processing level.

[3] Hog marketing institutions in Quebec are described in greater details in Larue *et al.* (2000).

[4] At this stage, the choice variable of producers is irrelevant given its monopsony position. As is well known, the decision variable would be important under different market structures such as an oligopsony. However, this would unduly clutter the analytical model because it would involve equilibria in mixed strategies.

[5] It is easily verified that the second order condition for a maximum is respected.

[6] There could be a possibility that processors' capacity might be constrained by the number of hogs below their desired level. Larue, Gervais and Lapan (2004) study this case in greater details.

[7] McKenzie (1999) terms the volatility measure used in our study a measure of "changeableness" in the real export price. Therefore, it may fail to capture the uncertainty in the exchange rate and/or the export price, as the movements in at least one variable may be at least partially predictable. McKenzie (1999) suggests using a measure based upon prediction errors such as ARIMA and ARCH models. The latter models also surfer from one serious flaw in that they are usually estimated over the whole sample and thus includes information that is not available to agents.

[8] It is worth emphasizing that this specification entails nature generating a single realization of $\mu(\cdot)$ prior to generating the observed data $\{x_t, \mathbf{z}_t\}_{t=1}^{T}$. The objective econometrician's task is to form inference about the nature of the realized value for $\mu(\cdot)$ based on the properties of the observed data.

[9] Certain conditions must be respected such as that non-diffuse priors must be specified for the non-linear component of the model.

[10] Given the definition of the volatility variables and their inherent correlation, a multicollinearity problem can arise in the estimation of (17) and (18). Hence, parameter estimates may not be very precise and can have large standard error.