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

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Abstract: Production and marketing lags in agri-food supply chains force competitive primary producers and food processors to commit to output targets before prices and exchange rates are realized. A theoretical model with one processor and many price-taking primary producers is developed to show that an increase in the volatility of the export price generally increases exports under risk neutrality. Furthermore, relaxing the assumption that the processing firm is risk neutral introduces non-linearities in the relationship between exports and export price volatility. This relationship is empirically investigated using the flexible non-linear inference framework developed by Hamilton (2001). The theoretical model provides the foundation for empirical bilateral export equations for Canadian pork exports to the U.S. and Japan. The empirical investigation supports the hypothesis that export price volatility has statistically significant non-linear effects on Canadian pork exports.

Keywords: 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 changes in the volatility of financial variables have significant impacts on trade, empirical evidence is notoriously mixed (McKenzie, 1999). A number of theoretical models have been proposed to explain the ambiguous impact of exchange rate volatility on trade. A commonly-held view in the literature is that risk aversion is a sufficient condition for exchange rate volatility to exert a negative impact on trade flows (McKenzie, 1999). This belief is corroborated by a large body of empirical studies that 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, Franke (1991), Dellas and Zilberfarb (1993) and Broll and Eckwert (1999) have shown that it is theoretically possible to find a positive correlation between exchange rate volatility and exports. Franke (1991) presents export markets as being similar to a put option held by domestic firms. An increase in the volatility of the exchange rate raises the payoff of the option which induces a proportional increase in trade. Dellas and Zilberfarb (1993) assume that export decisions are made after the uncertainty about the exchange rate is dissipated. Under certain conditions with regard to the level of risk aversion, higher volatility leads to higher exports. At least two empirical studies (Hooper and Kohlhagen, 1978; and Asseery and Peel, 1991) uncovered empirical evidence of a positive correlation between exchange rate volatility and trade.

It is fair to say that the literature is unclear about the nature of the relationship between exports and exchange rate volatility and ambiguities remain, both theoretically and empirically.

These ambiguities are what motivated Baum *et al.* (2004) to use aggregate export data from 13 developed countries to ascertain whether non-linearities in the relationship between exports and volatility may explain the contradictory empirical results reported in the literature. They find non-linear relations between bilateral exports and exchange rate volatility that vary across country pairs. However, their model restricts the range of plausible non-linear responses by assuming that non-linearities arise from interactions between exchange rate volatility and the volatility of economic activity (GDP) in the importing country.

The issue of exchange rate volatility is probably of greatest concern for sectors characterized by limited short run adjustment capabilities. In such sectors, the investment/capacity decisions might have to be made long before production and consumption decisions. There are many sectors that are constrained in such a way, but primary agricultural goods and processed food products are particularly fitting examples because of significant 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 sectors whose production decisions precede marketing decisions by several months. As such, agriculture is inherently risky even when climate-related risks are not taken into account.

The objective of this paper is twofold. First, a theoretical trade model accounting for production and marketing lags in agricultural supply chains is developed to analyze the effect of exchange rate volatility on the volume of trade. The theoretical model uncovers potential non-linear export responses to volatility. These responses are driven by two assumptions: the existence of a market in which there is no uncertainty (the domestic market in this case) and risk aversion. Under general conditions, the impact of export price volatility on exports cannot be determined a priori. There are two offsetting effects. First, export markets act as put options for the exporting firm. Under risk neutrality, an increase in the volatility of export prices increases total supply and (expected) exports. Second, risk aversion introduces significant non-linearities because risk tends to reduce the capacity commitment of the downstream firm when volatility is increasing.

In a related paper, Broll and Eckwert (1999) analyzed firms' behavior under exchange rate uncertainty and various risk preferences. Under risk neutrality, they show that higher uncertainty always increases exports because their competitive setting inevitably produces a corner solution. The firms' decision is either to sell all output domestically or to export all production. Under risk aversion, the two effects (the option value of the export market and risk aversion) condition the relationship between exports and volatility. However, our theoretical model is the first modeling effort that considers the non-linearity in exports induced by these two simultaneous effects.

The second objective is to gauge to what extent trade flow responses to exchange rate/price volatility suggested by our theoretical framework are consistent with observed empirical responses. The empirical investigation focuses on how Canadian pork exports to the United States and to Japan are impacted by the volatility of the export price expressed in Canadian dollars. To achieve this end, we search for the sort of non-linearities uncovered by our theoretical model 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 test of non-linearity than the procedure adopted by Baum *et al* (2004). The empirical model detects significant non-linearities in the relationship between Quebec pork exports to the U.S. and export price volatility. It also clearly identifies significant non-linearities between Canadian pork exports to Japan and export price volatility; but the nature of this relationship is more difficult to reconcile with our theoretical results.

The rest of the paper is structured as follows. The next section introduces the theoretical model by characterizing the dynamic nature of the primary input marketing mechanism. The emphasis is on the manner with which marketing lags influence the downstream firm and upstream producers' output decisions. As mentioned before, we chose to focus on one particular agri-food sector in developing our theoretical model and empirical application, but our model and the conclusions derived from it generalize easily to other situations in which production/marketing lags exist. The third section begins by describing the pattern of bilateral pork exports and export price volatility. This is followed by the presentation of the empirical model and the results of the estimation. The final section offers concluding remarks.

2 – The Theoretical Model

This section develops 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.[1] It has monopoly power on the domestic market, but its exports have a negligible effect on its country's terms of trade (*i.e.*, the small country assumption). The assumption of monopoly behavior is reasonable in our setting given the significant literature documenting the increasing concentration at the processing level in agri-food markets (see for example Lopez *et al.*, 2002).

While the current model can be applied to many 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/technology, the price commitment determines how many live animals will be processed domestically in the second period. At the beginning of the 2nd period, uncertainty about the foreign pork price is resolved and the processor markets 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. Under such a mechanism, a large percentage of total hog supply is 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.[2] The marketing assumptions are also consistent with a situation in which supply of live animals is secured through contracting.

The two-stage 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 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 $[\theta, \eta]$ with density $\frac{1}{\eta - \theta}$. We assume that $\eta = -\theta > 0$, so the unconditional mean of the export price is $\bar{p}^x$ 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:

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

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: $q^d + q^x \leq q^T$. Given that the first-stage cost to invest in capacity is sunk, it follows that the processor maximizes revenue by selling on either or on 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 large, $\left(\left(1 - \bar{p}^x\right) / \lambda < \varepsilon\right)$, that it may be more profitable for the monopolist to ignore the domestic market $\left(q^x = q^T, q^d = 0\right)$. In the latter case, the processor's profit function is: $\pi = \left(\overline{p}^x + \lambda\varepsilon\right)q^T - r^d q^T$.

It should be emphasized that when hog production decisions are made in the first stage of the game, the 2nd period realization of the export price denominated in Canadian dollars is not known; but all agents know the distribution of the random variable. For future reference, it is useful to define the following bounds on the export price.

Definition I: *The minimum random shock on the export price that guarantees that exports will be positive in equilibrium is* $\theta^e \equiv \left(1 - \overline{p}^x - 2q^T\right)/\lambda$ *. Similarly, the maximum random shock on the exchange rate that guarantees that domestic sales will be positive in equilibrium is defined as* $\eta^d \equiv \left(1 - \overline{p}^x\right)/\lambda \geq \theta^e$.

Definition 1 is important because it establishes thresholds on the distribution of the export price that yield six possible cases related to exporting and domestic sale decisions. The first case is rather uninteresting in that if $\theta < \eta < \theta^e < \eta^d$, the processing firm supplies only the local market as the inequalities prevent exports from ever occurring. The opposite case with the processing firm being present only on the export market requires $\theta^e < \eta^d < \theta < \eta$. Third, the inequalities $\theta^e < \theta < \eta < \eta^d$ guarantee an interior solution characterized by the equalization of marginal revenues from domestic and exports sales. Fourth, the inequalities $\theta^e < \theta < \eta^d < \eta$ imply that for some values of $\varepsilon \in \left(\theta, \eta^d\right)$, there will be an arbitrage between domestic sales and exports while for values of $\varepsilon \in \left(\eta^d, \eta\right)$, the processor will only sell on the foreign market. In the fifth case, the conditions $\theta < \theta^e < \eta < \eta^d$ imply that for $\varepsilon \in \left(\theta, \theta^e\right)$, the processor sells only to local consumers while $\varepsilon \in \left(\theta^e, \eta\right)$ implies sales are arbitraged between domestic and foreign markets. Finally, the following inequalities $\theta < \theta^e < \eta^d < \eta$ imply that all previously discussed situations are possible.

For the time being, it is assumed that the processor is risk-neutral. Its expected profits are computed by substituting the decision rule for domestic sales into (1). If we consider for example, the case in which $\theta < \theta^e < \eta^d < \eta$, expected profits are:

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

From (3), it is straightforward to analyze all cases presented following definition I by appropriately redefining the domains of integration. The first component in (3) measures expected profits when there are no export sales, whereas the second and third components represent respectively expected profits when domestic and foreign sales are positive and only foreign sales are observed. In what follows, we consider all possible cases.

As mentioned earlier, hog marketing institutions play an important role in determining the processor's output capacity. It is assumed that the downstream monopolist commits to a price in the 1st period to target a level of total hog production supplied by perfectly competitive hog producers in the 2nd period. The profit of a representative hog producer is assumed to be:

$$\pi^{prod} = r^d q^T - 0.5 c {q^T}^2 \qquad (4)$$

The first-order condition for profit maximization determines total hog supply, $q^T = r^d / c$. 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 time sows are inseminated and the time piglets attain the ready-to-market hog weight. For future reference, we define $RT\left(q^T\right)$ as the processor's expected revenue which corresponds to the sum of all three integrals in (3).

The processor's total costs are: $CT = r^d q^T$ and the processor's capacity is determined by the first-order condition with respect to the hog price commitment:[3]

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

Equation (5) determines the hog price commitment of the processor which in turn determines total capacity in the industry given the producers' hog supply:[4]

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

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 the processor's capacity. The focus of the paper is on the relationship between exports and the volatility of the export price denominated in domestic currency. Comparative static on equation (6) leads to the following proposition.

Proposition 1: *For all admissible values of* λ, $dq^x/d\lambda = 0$ *if* $\eta < \theta^e$ *or* $\theta^e < \theta$. *For all admissible values of* λ *such that* $\theta < \theta^e < \eta$, $dq^x/d\lambda > 0$.

Proof: See the technical appendix.

It should be noted that the comparative static exercise in proposition 1 is implemented from an *ex-ante* perspective. It argues that the volatility of the export price has an impact on exports only if it is more profitable for the processor to serve only the local market for some realizations of ε, *i.e.* $\theta < \theta^e < \eta$. In this case, an increase in the volatility of the export price induces an increase in the level of planned exports. Henceforth, we call this effect the λ-*effect*. An increase in λ represents an increase in the mean preserving spread of the export price, but it is not mean-preserving with respect to marginal revenue as it increases the expected marginal revenue of the firm which in turn increases planned exports. The ability of the firm to sell all of its capacity on the domestic market when the observed shock is just sufficient to prevent exports makes the firm immune to stronger negative shocks in the sense that the (corner solution) outcome will be the same for the firm. When $\theta^e > \eta$, it is not profitable to export and thus a change in λ does not affect expected marginal revenue. Its effect on planned exports is obviously nil in that case. The results in proposition 1 are consistent with Franke (1991)'s result which likens the export market to a put option. Due to our market structure assumption, exports are non linear in the volatility measure as there are volatility thresholds for which planned exports are increasing in volatility. Franke (1991) obtained a similar result by making the strong assumption that there are entry and exit costs on export markets that decline with volatility. The results in proposition 1 also generalize the findings of Broll and Eckwert (1999) under risk neutrality in that our model considers all possible equilibria based on the distribution of the export price.

The intuition behind the positive effect of volatility on the capacity choice of a risk neutral firm clearly emerges when one considers the effect of volatility on the firm's pricing decision once capacity is chosen as in Figure 1. At that stage, the average price received by the firm is a weighted average of the domestic price and the export price. As long as there are domestic sales, the domestic price exceeds the export price and the average price lies somewhere in between. If initially the export price is such that $p^x \in \left[\bar{p}^x - \lambda_1\varepsilon, \bar{p}^x + \lambda_1\varepsilon\right]$, the lowest possible average price is $p_{\min}$ at which $q^d = q^T$. When volatility increases $\left(\lambda_2 > \lambda_1\right)$,

$p^x \in \left[\bar{p}^x - \lambda_2\varepsilon, \bar{p}^x + \lambda_2\varepsilon \right]$, the minimum average price remains the same, but the maximum average price increases. Thus, the increased volatility increases the expected return on the q^T units to be marketed which in turn induces an upward adjustment in the chosen capacity level.[5]

We now entertain the possibility that the downstream firm be risk-averse by assuming that its preferences towards risk can be characterized by the first two moments of the distribution of its profits. This assumption can be reconciled with expected utility theory in the current context if the utility function is quadratic in profits (Levy and Markowitz, 1979). The objective function of the processing firm is:

$$E\left[U(\pi)\right] = E\left[\pi\right] - (\alpha/2)\ Var(\pi) \qquad (7)$$

where α can be interpreted as the Arrow-Pratt measure of absolute risk aversion. Optimal capacity is determined by maximizing (7). The comparative static effects on the optimal capacity are summarized in the following proposition.

Proposition 2: *Suppose that the processing firm is risk-averse* $(\alpha > 0)$. *Provided that* $\eta < \theta^e$, $dq^x/d\lambda = 0$. *For all values of* λ *such that* $\theta^e < \theta$, $dq^x/d\lambda < 0$. *Finally, when* $\theta < \theta^e < \eta$, $dq^x/d\lambda \underset{>}{\overset{<}{-}} 0$.

Proof: See the technical appendix.

According to Proposition 1 which was derived under the assumption of risk neutrality, volatility impacts on capacity only if $\theta < \theta^e < \eta$. The first part of proposition 2 is similar in the sense that if exports occur in equilibrium for all possible realizations of the random shock, a change in λ has no effect. More interestingly, Proposition 2 states that volatility has an impact even if this minimum random shock guaranteeing positive exports is lower than the minimum bound of the distribution of the random shock (*i.e.*, if $\theta^e < \theta$). In that case, although an increase in λ leaves unchanged the expected marginal revenue[6], it increases the volatility of the firm's payoff and this has a negative impact on capacity. This negative response of exports to an increase in λ is dubbed the α*-effect*. Under the condition $\theta < \theta^e < \eta$, volatility has an impact whether the firm is risk-averse or not. However, the λ and α effects work in opposite directions under risk aversion. An increase in λ tends to increase planned exports because expected marginal revenue (weakly) increases; but it also tends to reduce exports because of its effect on risk. As a result, the overall impact of an increase in the mean preserving spread of the export price is ambiguous. This ambiguity in the relationship between exports and volatility was previously documented in Franke (1991). The advantage of our model is that it clearly describes the role of risk preferences without having to invoke transaction costs declining with volatility to explain potential non-linearities. Most importantly, the option value of the export market combined with risk aversion implies that volatility *simultaneously triggers* two opposing effects on exports which must be properly accounted for in empirical studies.

In order to shed some light on potential non-linearities that could be encountered empirically when $\theta < \theta^e < \eta$ and risk aversion prevail $(\alpha > 0)$ and the λ*-effect* and the α*-effect* are at work, we performed some simulations. Figure 3 and 4 portray different export responses as a function of the mean preserving spread parameter λ. They illustrate the two competing effects on the capacity choice.[7] Each figure includes two different regions on the horizontal axis for $\theta < \theta^e < \eta$. Region 1 is defined by the condition $\eta < \eta^d$ which implies that domestic sales are observed in this domain. Conversely, region 2 is defined by the condition $\eta^d < \eta$ which implies "exports-only" equilibria. Figure 3 presents exports as a function of λ when the coefficient of risk aversion is small (i.e., $\alpha = 0.25$). For low levels of λ, the α*-effect* offsets the positive impact of λ on marginal revenue and thus exports are decreasing in λ. However, the λ*-effect* offsets the α*-effect* when λ increases past a certain threshold. The latter is positioned at λ=0.727 in Figure 3 (region 1). Figure 4 illustrates the impact of λ on exports when the downstream processing firm has a larger coefficient of risk aversion (*i.e,* $\alpha = 1$). In this instance, the α*-effect* unambiguously dominates the λ*-effect* for all values of λ. Note that when the maximum random shock on the export price that guarantees positive domestic sales in equilibrium is below the upper bound of the distribution $\left(\eta^d < \eta\right)$, there is a structural change in the relationship between exports and λ that may or may not involve a sign reversal. For the chosen parameter values, the inequality $\eta^d \leq \eta$ occurs when $\lambda \geq \left(1 - \overline{p}^x\right) / \eta = 0.8$. However, the conflicting effects of the two forces that condition the effect of λ on exports in region 1 are also present in region 2, but their strength differ due to the aforementioned structural change at the λ value that separates the two regions.

The solution defined by the optimization problem in (7) yields the optimal capacity choice of the processor: $q^T = q^T\left(\gamma\left(\tilde{p}^x\right); \boldsymbol{\beta}\right)$; where $\gamma\left(\tilde{p}^x\right)$ is a function mapping the different moments of the distribution of the export price and $\boldsymbol{\beta}$ is a vector representing all other exogenous variables of the model, such as the risk aversion coefficient. Substituting the optimal capacity choice of producers in the first-order condition defined in (2) yields domestic sales and exports: $q^{d*}\left(p^x, \gamma\left(\tilde{p}^x\right); \boldsymbol{\beta}\right)$ and $q^{x*}\left(p^x, \gamma\left(\tilde{p}^x\right); \boldsymbol{\beta}\right)$ respectively. Consequently, export and domestic sales are both function of the realized export price and the different moments of the export price distribution. Obviously, exports and domestic sales respond to exchange rate changes occurring after the determination of capacity, but such adjustments are offsetting (*i.e.*, $\Delta q^{x*} = -\Delta q^{d*}$). Furthermore, the choice of exports, once capacity is chosen, is linearly impacted by the realized export price p^x. Hence, non-linearities are driven solely by the effect of volatility on the capacity choice.

3 – The Empirical Model

Uncertainty in the model arises because of lags in the production and marketing of the primary commodity. The theoretical framework underlines two key factors conditioning export decisions of processors. First, even though a processor is risk-neutral, its selected level of exports can be influenced by the second moment of the distribution of the export price as stated in Proposition 1. Second, risk aversion introduces non-linearities in the relationship between exports because volatility has two opposite effects on exports. The net effect of volatility on exports is very sensitive to parameter values as shown in figures 3 and 4. In spite of the wide ranging sort of responses to volatility generated by our theoretical model, it must be conceded that our simplistic assumptions about the technologies, market structure, consumers' preferences and risk preferences, tend to limit the possible non-linearities between exports and volatility from a theoretical standpoint.

In what follows, lagged values of the export price and volatility are used as proxies for the expected export price and variance of the export price respectively. Hog production is characterized by production lags of around ten months between the time sows are inseminated and pork meat is marketed. A number of different lag specifications were experimented with, but a ten-month lag performed best.

Past studies provided evidence that a destination-specific volatility measure of the exchange rate plays an important role in determining exports to that market (e.g., Baum *et al.*, 2004). The theoretical model did not explicitly account for multi-market sales and some adjustments to the theory must be made to account for destination-specific volatility. The empirical model segregates the effects of aggregate export price volatility on exports into destination-specific volatility effects on bilateral exports. The lagged aggregate export price is included as an independent variable and is not destination-specific. Finally, the current export price to a specific destination also enters the model's specification because it is a determinant of profitability in a given market.

Figure 5 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 1992 and ending November 2003. 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 near the end of the sample as Japan and the U.S. became relatively less important destinations. Figure 6 presents monthly export unit values in Canadian dollars between January 1992 and November 2003 for each destination. Unit values for Japan are significantly higher than for the U.S. as the product mix of pork meat exports is significantly different between the two destinations due in part to Japan's minimum import price policy (Obara, Dyck, and Stout, 2003).

Our theoretical framework provides the foundation for the specification of our empirical export equations. It should be emphasized that the present analysis focuses on the distribution of the export price defined as the foreign market price received by the firms denominated in Canadian currency. Mackenzie (1999) surveys the various volatility indicators used in the literature. In the current application, volatility is defined as a moving average of the standard deviation of the export price:[8] $V_t = \left[1/m \sum_{i=1}^{m} \left(e_{t+i-1} p_{t+i-1}^{x} - e_{t+i-2} p_{t+i-1}^{x} \right)^2 \right]^{1/2}$. Figure 7 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. export price follows closely the volatility of the aggregate export price. There are however significant differences between the two measures mainly due to sudden surges in the volatility of the export price in Japan. In order to better gauge the robustness of our volatility measure to the choice of the parameter *m*, we computed alternative volatility measures $(m = 3, 6)$, but they generated similar qualitative results, although the measure based on the longer lag generally yielded 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 concern and this is why we began our empirical investigation by analyzing 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 values of the level of the series, a constant, a time trend and, if needed, lagged first differences of the dependent variable to make the residuals white noise:

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

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 export price, export sales and the volatility of the export price. 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 a unit root is not rejected for both volatility measures. All other variables do not seem be integrated of order one.

To assess the reliability of the ADF test, the stationarity test developed by Kwiatkowski *et al.* (1992, hereafter referred to as KPSS) was implemented. The KPSS test complements unit root tests because its null hypothesis is that of stationarity. The KPSS test is implemented by estimating the equation:

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

The null hypothesis of trend stationarity can be ascertained by testing $\sigma_u^2 = 0$. Testing for the null of level stationarity instead of trend stationarity can be done by regressing the series on a constant instead of a 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 reports that the null hypothesis of stationarity is rejected for all variables at the 90% confidence level except for the volatility measure in the U.S. market. Unfortunately, the ADF and KPSS tests yield conflicting evidence; an outcome previously documented in Maddala and Kim (1998). 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 significantly improves 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 the null hypothesis of a unit root is jointly confirmed at the 95% confidence level by the two tests only in the case of the 12-month volatility measure in the Japanese market.

As mentioned previously, the pork export equations are expected to exhibit significant non-linearities in the various moments of the distribution of the export price. To account for these potential non-linearities, the flexible non-linear inference framework developed by Hamilton (2001, 2003) is applied. Hamilton's approach begins with the estimation of a nonlinear regression model of the form: $x_t^* = \mu(\mathbf{z}_t) + \upsilon_t$; where υ_t is a normally distributed random error term with a zero mean and a 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.[9] For a given non-stochastic vector $\mathbf{z}$, the function $\mu(\mathbf{z})$ is assumed to be normally distributed with mean $\gamma_0 + \boldsymbol{\gamma}_1\mathbf{z}$ and variance λ^2. If the variance is zero, the regression equation reduces to $x_t^* = \gamma_0 + \boldsymbol{\gamma}_1\mathbf{z}_t + \upsilon_t$; a standard linear regression framework. However, when λ is large, the export equation 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) assumes that two random realizations, $\mathbf{z}_1$ and $\mathbf{z}_2$, 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 the $\mathbf{z}_j$ vary. When the previous inequality is not satisfied, it can be inferred that the correlation differs from zero and its exact form is described in Hamilton (2001, p. 542).

The regression equation can be rewritten as: $x_t^* = \gamma_0 + \boldsymbol{\gamma}_1\mathbf{z}_t + \lambda m(\mathbf{z}_t) + \upsilon_t$; where $m(\cdot)$ is a stochastic process that characterizes the conditional expectation $\mu(\mathbf{z}_t)$. This process has mean zero and unit variance. The parameters to be estimated are the coefficients $(\gamma_0, \boldsymbol{\gamma}_1)$ of the linear regression, the parameter indicating the presence of a non-linear component (λ), the variance of the error term $\left(\sigma^2\right)$ 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 variances, exports must be a stationary time series. Table 1 confirms that this condition is met for Quebec pork exports sold to the U.S. and Japan.

Finally, Hamilton (2003) suggests reporting the estimation results as:

$$x_t^* = \gamma_0 + \boldsymbol{\gamma}_1 \mathbf{z}_t + \sigma \cdot \omega \cdot m(\mathbf{z}_t) + \sigma\varepsilon_t \tag{10}$$

where the innovation υ_t is replaced by the product of σ and ε_t (which follows a standard normal distribution), and the parameter λ is re-parameterized as $\lambda = \sigma \cdot \omega$. We begin our investigation with bilateral exports to the U.S.[10] As mentioned previously, we assume that exports $\left(x_t^{US}\right)$ are function of the current realized export price in the U.S. market $\left(p_t^{US}\right)$, the lagged average export price across all destinations $\left(p_{t-10}^{X}\right)$, and lagged volatility in the U.S. and Japanese markets $\left(vol_{t-10}^{US}, vol_{t-10}^{Jap}\right)$. The purpose of these volatility variables is to decompose the effect of the aggregate volatility of the export price on the 1st stage capacity choice of processors into destination-specific volatility effects.[11] The ten-period lag length captures the biological constraints in adjusting capacity in the hog industry. The superscript *x* identifies variables that pertain to aggregate exports and are not destination-specific.

The maximum likelihood coefficient estimates of (10) and their standard error (between parentheses) for the bilateral Canada-U.S. export equation are:

$$\begin{aligned} x_t^{US} &= \underset{(0.23)}{8.91} - \underset{(0.14)}{0.39}\, p_t^{US} - \underset{(0.15)}{0.17}\, p_{t-10}^{X} - \underset{(0.08)}{0.04}\, vol_{t-10}^{US} - \underset{(0.04)}{0.09}\, vol_{t-10}^{Jap} + \underset{(0.0004)}{0.008}\, Trend \\ &\quad + \underset{(0.01)}{0.07}\left[\underset{(0.43)}{1.13}\, m\left(\underset{(3.81)}{11.22}\, p_t^{US} + \underset{(3.60)}{9.84}\, p_{t-10}^{X} + \underset{(2.66)}{6.77}\, vol_{t-10}^{US} + \underset{(1.74)}{3.79}\, vol_{t-10}^{Jap}\right)\right] \end{aligned} \tag{11}$$

A time trend is also included in (11) because exports are assumed to be stationary around a deterministic trend. A non-negligible advantage of the flexible non-linear framework is that it allows for a direct test of the null hypothesis that the true relation in (10) is linear. This amounts to testing whether λ^2 is different from zero with a Lagrange Multiplier (LM) test. The null hypothesis of a linear model is soundly rejected in light of the *p-value* of the LM test of 0.001. Of the outmost interest is the fact that all coefficients in the linear part of (11) have a relatively large standard error except for the constant and the time trend. Conversely, all the parameters in the non-linear component of (11) are positive and significantly different from zero.

In order to assess the advantage of Hamilton's flexible framework over the usual empirical applications, the OLS estimates of the linear component in equation (10) were computed. The coefficient estimates and their standard errors are:

$$\begin{array}{l} x_t^{US} = \underset{(0.77)}{8.91} - \underset{(0.11)}{0.45} p_t^{US} - \underset{(0.11)}{0.11} p_{t-10}^{X} - \underset{(0.06)}{0.04} vol_{t-10}^{US} - \underset{(0.03)}{0.10} vol_{t-10}^{Jap} + \underset{(0.0002)}{0.008} Trend \end{array} \qquad (12)$$

The results in (12) would be quite disheartening if we had to rely on them because the lagged U.S. volatility variable is not significant and the U.S. and lagged aggregate export prices have the wrong algebraic sign. This confirms that ignoring the potential non-linearity of the export equation can result in severe misspecification biases. Although the results in (11) provide evidence that the relationship between exports and volatility is non-linear, it is difficult to infer what this non-linear relationship looks like. Hamilton (2003) suggests fixing all but one of the independent variables to their sample mean and to examine the impact of variations in one variable on the conditional mean of $\mu(\mathbf{z})$ in the export equation.

Figure 8 plots the response of Quebec exports to the U.S. market to changes in the lagged volatility of the U.S. export price holding all other independent variables fixed at their mean. In other words, it shows how $\mu\left(\overline{p}_t^{US}, \overline{p}_{t-10}^{X}, \overline{vol}_{t-10}^{Jap}, vol_{t-10}^{US}, \overline{Trend}\right)$ changes as the export price volatility varies from $\pm$ two times its standard deviation around its mean. There are significant non-linearities in the lagged volatility of the U.S. export price. The plot of point estimates of $\mu(\mathbf{z})$ confirms that the conditional mean is not monotonic in vol^{US}. Starting at low levels of volatility, increases in volatility decrease exports, but at higher levels, further increases trigger increases in exports. This occurs when volatility approaches its mean value. The evidence suggests that although increases in volatility can potentially increase the expected payoff from export activities, there are levels of volatility for which export activities are less attractive. In any case, there are substantial differences in the predictions between the linear and non-linear models. It should be noted that Figure 8 mimics quite well the numerical simulations from our theoretical model displayed in Figure 3. An increase in volatility increases the expected payoff of the firm, but it also increases risk. As demonstrated earlier, the α-effect and the λ-effect have orthogonal impacts on exports under risk aversion and this is why the relationship between exports and volatility can experience a sign reversal.

The maximum likelihood estimates for Canadian pork exports to Japan, along with their standard errors, are shown below:

$$\begin{aligned}
x_t^{Jap} &= \underset{(0.95)}{3.78} + \underset{(0.37)}{0.56}\, p_t^{Jap} + \underset{(0.37)}{0.37}\, p_{t-10}^{X} - \underset{(0.13)}{0.23} vol_{t-10}^{Jap} - \underset{(0.23)}{0.51} vol_{t-10}^{US} + \underset{(0.002)}{0.02} Trend \\
&\quad + \underset{(0.05)}{0.15}\left[\underset{(0.91)}{2.28} m\left(\underset{(2.29)}{10.59}\, p_t^{Jap} + \underset{(0.36)}{0.00}\, p_{t-10}^{X} + \underset{(3.04)}{10.87} vol_{t-10}^{Jap} + \underset{(1.62)}{5.25} vol_{t-10}^{US}\right)\right]
\end{aligned} \tag{13}$$

The Lagrange multiplier test did not reject the null hypothesis of non-linearity. The OLS estimates of the linear component in (10) with their standard errors are:

$$x_t^{Jap} = \underset{(0.82)}{4.00} + \underset{(0.32)}{0.32}\, p_t^{Jap} + \underset{(0.38)}{0.53}\, p_{t-10}^{X} - \underset{(0.10)}{0.29} vol_{t-10}^{Jap} - \underset{(0.18)}{0.51} vol_{t-10}^{US} + \underset{(0.002)}{0.02} Trend \tag{14}$$

The coefficients of the volatility variables in (14) are statistically different from zero and negative. Hence, one would conclude from this linear model that exports to Japan are negatively correlated with volatility. Moreover, the coefficients of the export price (both country specific and aggregate) are not significant. The non-linear specification in (13) indicates that there are significant non-linearities in the lagged volatility measures since the coefficients for these variables are significant and the estimate of σ and ω are quite large compared to their standard error. The coefficients of the linear component in (13) are quite similar to the coefficients in (14).

Figure 9 presents the marginal impacts on Quebec exports to Japan of changes in the lagged volatility of the export price in the Japanese market. It illustrates the importance of destination-specific volatility. Again, starting at low levels of volatility, increases in the volatility measure have a negative impact on exports. This trend looks quite linear at the beginning. However, there is a threshold in volatility above which exports increase rapidly with volatility. Beyond the domain over which there is a rapid increase in exports, the impact of destination-specific volatility is not clear. This behavior seems consistent with kinks in the expected marginal revenue function induced by the ability of the processor to price discriminate.

4 - Concluding Remarks

The literature on the impact of exchange rate volatility on exports is voluminous yet puzzling in light of the conflicting empirical evidence reported in studies conducted on aggregated data as well as in studies relying on disaggregated data. Intuitively, volatility should matter most in sectors in which firms face severe constraints limiting their ability to respond to changes in exchange rates/export prices. Most agricultural sectors are characterized by long periods of time between production and marketing decisions. As such, individual and aggregate supplies are very inelastic once production decisions are made. In this context, exchange rate volatility can bring about large differences between expected and realized profits. The theoretical and empirical models developed in this paper are motivated by the Quebec hog/pork industry, but they could be applied in numerous other settings for which capacity decisions must be made before marketing decisions by imperfectly competitive firms.

Our theoretical model demonstrates that export price volatility can decrease, leave unchanged or increase the chosen capacity and hence exports depending on the distribution assumption about the export price. The positive effect of volatility on capacity is due to price discrimination between domestic and export markets and the option of selling only domestically when the realized export price is below its expected value. In this instance, increased volatility translates into higher expected average returns. However, increases in volatility also increase risk which encourages a risk-averse firm to reduce its capacity *ex-ante* and thus expected exports. The theoretical model suggests that the relationship between export price volatility and exports is non-linear, yet sensitive to changes in the parameters embodying risk preferences, the degree of volatility and the mean export price.

For empirical purposes, we define export price volatility as the variability of the foreign market price for exports converted back to Canadian currency. Our empirical model must account for the potential non-linearities identified in our theoretical model and this is why we relied on Hamilton's (2001) flexible estimation approach. The empirical results strongly reject the hypothesis of linearity in the relationship between destination-specific volatility measures and exports from Quebec to the U.S. and Japan. We also estimated linear specifications to show to what extent they can be misleading. In particular, the linear models suggest that volatility of the export price in the U.S. market has no effect on exports while the export equation for the Japanese market suggests that volatility is negatively linked to exports. In contrast, Hamilton's flexible approach uncovered significant non-linearities in the relationship between exports and volatility, confirming that the sign of the derivative depends on the degree of volatility at which it is evaluated. As such the empirical results for the two destination markets are consistent with the existence of the two conflicting volatility effects identified in our theoretical model.

The theoretical and empirical frameworks could be extended in two important ways. The positive response of exports to exchange rate volatility is primary due to the existence of a lower bound for the expected marginal revenue of the firm. It would be interesting to extend further the portfolio analogy to export activities by introducing a second export market. It is likely that if correlation exists between the two exchange rates, it will be positive. Hence, it would be unlikely that a second export market could be used to diversify risk in the usual sense that assets' payoff (returns in the export market) move in opposite directions. However, if volatility in one market increases while leaving unchanged volatility in the second destination's market, capacity could increase *ex-ante* because of the upside risk. From an empirical standpoint, it would require modeling the dependency between the two export markets. Second, it would have been instructive to decompose the measure of volatility into a pure exchange rate volatility measure and a price variability measure. The question of whether commodity price or exchange rate volatility is a more important determinant of exports remains open for future research.

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6 - Technical appendix

Proof of Proposition 1

We use the fact that total capacity and exports are positively linked. It is easy to demonstrate that if $\eta < \theta^e$, total capacity is $q^{T*} = 0.5/(1+c)$ because there are no planned exports. Conversely, if $\theta^e < \theta < \eta < \eta^d$, $\theta^e < \theta < \eta^d < \eta$ or $\theta^e < \eta^d < \theta$, total capacity is: $q^{T*} = \overline{p}^x/2c$ because an equilibrium in which there are only domestic sales cannot occur. Hence, for all values of λ, and when $\eta < \theta^e$ or $\theta^e < \theta$, we have that λ has no impact on capacity $\left(dq^T/d\lambda = 0\right)$ and thus no impact on expected exports $\left(dq^x/d\lambda = 0\right)$. This proves the first part of Proposition 1.

The second part of proposition 1 can be proved using the first order condition defined in (5). Multiply (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, *i.e.* $\partial E[\pi]/\partial q^T = MR - MC = 0$; where $MC \equiv \left(q^T \dfrac{\partial r^d}{\partial q^T} + r^d\right) > 0$. Comparative static on the previous equation yields: $\dfrac{dq^T}{d\lambda} = \dfrac{\partial MR}{\partial \lambda} \Big/ \left(\dfrac{\partial MC}{\partial q^T} - \dfrac{\partial MR}{\partial q^T}\right)$. The second order condition for a maximum requires that $\partial MC/\partial q^T - \partial MR/\partial q^T > 0$. It follows that $sign\left(dq^T/d\lambda\right) = sign\left(\partial MR/\partial \lambda\right)$. When $\theta < \theta^e < \eta$, the derivative of marginal revenue with respect to λ, $\partial MR/\partial \lambda = \dfrac{\eta^2\lambda^2 - \left(\overline{p}^x - 1 + 2q^T\right)^2}{4\eta\lambda^2}$, is greater than zero if $\left(\eta - \left(\dfrac{\overline{p}^x - 1 + 2q^T}{\lambda}\right)\right)\left(\eta + \left(\dfrac{\overline{p}^x - 1 + 2q^T}{\lambda}\right)\right) > 0$. This result is true for both cases $\theta < \theta^e < \eta < \eta^d$ and $\theta < \theta^e < \eta^d < \eta$). Since $\eta > \left(1 - \overline{p}^x - 2q^T\right)/\lambda \equiv \theta^e$ and

$\eta = -\theta > -\left(1 - \overline{p}^x - 2q^T\right)\Big/\lambda \equiv -\theta^e$, it follows that $\partial MR/\partial\lambda > 0$. Hence, when $\theta < \theta^e < \eta$, we have that $dq^T/d\lambda > 0$ and $dq^x/d\lambda > 0$ because exports are positively correlated with capacity; thus proving the second part of Proposition 1.

Proof of Proposition 2

As in the previous proof, we use the fact that total capacity and exports are positively linked. Proving that $dq^x/d\lambda = 0$ when $\eta < \theta^e$ does not rely on specific assumptions regarding risk preferences because the condition $\eta < \theta^e$ implies that there will be no export sales.

The proof of the second part of the proposition relies on showing that exports are decreasing in λ $\left(dq^x/d\lambda < 0\right)$ for the three possible cases: 1) $\theta^e < \theta < \eta < \eta^d$; 2) $\theta^e < \eta^d < \theta$; and 3) $\theta^e < \theta < \eta^d < \eta$. 1) The previous inequalities imply that the expected marginal revenue is unaffected by changes in λ. The mean preserving spread parameter only impacts the second moment of the distribution of profits. Under the assumption that $\theta^e < \theta < \eta < \eta^d$, we have that $dq^T/d\lambda = -\left[6\left(\overline{p}^x\left(1+c\right)-c\right)\alpha\,\eta^2\lambda\right]\Big/\left(6c+\alpha\eta^2\lambda^2\right)^2$, and thus $dq^T/d\lambda < (>) 0$ if $\overline{p}^x > (<) c/\left(1+c\right)$. Since marginal revenue is $c/\left(1+c\right)$ if there are no exports*, if $\overline{p}^x < c/\left(1+c\right)$ then for some values of ε, it is more profitable for the processing firm to sell exclusively the local market, it follows that $\overline{p}^x < c/\left(1+c\right)$ is possible only when $\theta < \theta^e$. Then $\forall\lambda$, such as $\theta^e < \theta < \eta < \eta^d$, $\overline{p}^x > c/\left(1+c\right)$, which implies that $dq^T/d\lambda < 0$ and $dq^x/d\lambda < 0$. 2) In the case

* When there are no planned exports, we have that $dRT/d\,q^T = 1 - 2q^T$ and $dCT/dq^T = 2cq^T$, with $RT = \left(1-q^T\right)q^T$ and $CT = rq^T = cq^{T^2}$. Since under the distribution assumptions $q^T = 0.5/\left(1+c\right)$, marginal revenue evaluated at the optimal solution is $c/\left(1+c\right)$; it must also equal marginal cost from the profit maximization first-order condition.

for which $\theta^e < \eta^d < \theta$, we have that $q^T = 3\overline{p}^x \big/ \left(6c + \alpha\,\eta^2 \lambda^2\right)$, it follows that $dq^T \big/ d\lambda < 0$ and $dq^x \big/ d\lambda < 0$. 3) The closed-form solution in the case for which $\theta^e < \theta < \eta^d < \eta$ does not yield sufficient conditions on the parameters of the model to establish that $dq^T \big/ d\lambda < 0$. However, a whole set of numerical solutions suggest that exports are decreasing in λ. For example, Figure 2 illustrates the impact of changes in λ for $\lambda \in \left(0.6, 0.75\right)$ given the parameters in the model have been set at: $\eta = 0.5$, $c = 1$, $\overline{p}^x = 0.7$, and $\alpha = 0.25$. The interval of admissible values for λ guarantees that the inequalities $\theta^e < \theta < \eta^d < \eta$ are verified in the numerical example.

To prove the third part of the proposition, we must show that for $\theta < \theta^e < \eta < \eta^d$ and $\theta < \theta^e < \eta^d < \eta$, the impact of an increase in the mean preserving spread is ambiguous because two effects work in opposite directions. This is done using a numerical example that is illustrated in Figure 3. It illustrates the impact of λ on output capacity when $\eta = 0.5$, $c = 1$, $\overline{p}^x = 0.6$, $\alpha = 0.25$, and $\lambda \in (0.503, 0.8)$ or $\lambda \in \left(0.8, 1.2\right)$. The former interval guarantees that the conditions $\theta < \theta^e < \eta < \eta^d$ hold. In this case (region 1), exports are not a monotonic function of λ. Exports initially decrease as λ increases and follow an upward trend once the value of λ reaches 0.727. When $\theta < \theta^e < \eta^d < \eta$ (region 2), the impact λ on exports is also ambiguous. Note that when λ is higher than 1.2, we have that $\lambda\theta = -\overline{p}^x$ and thus can no longer be interpreted as a mean preserving spread. This case is not considered. Q.E.D.

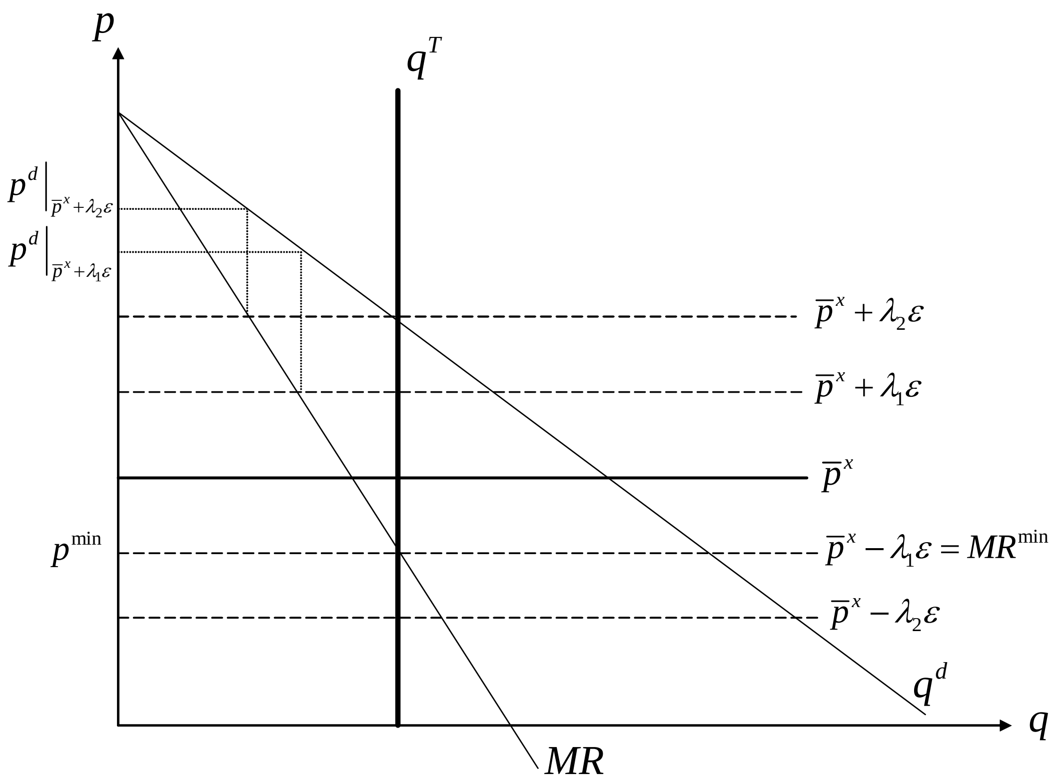

Figure 1. The impact of volatility given a capacity choice

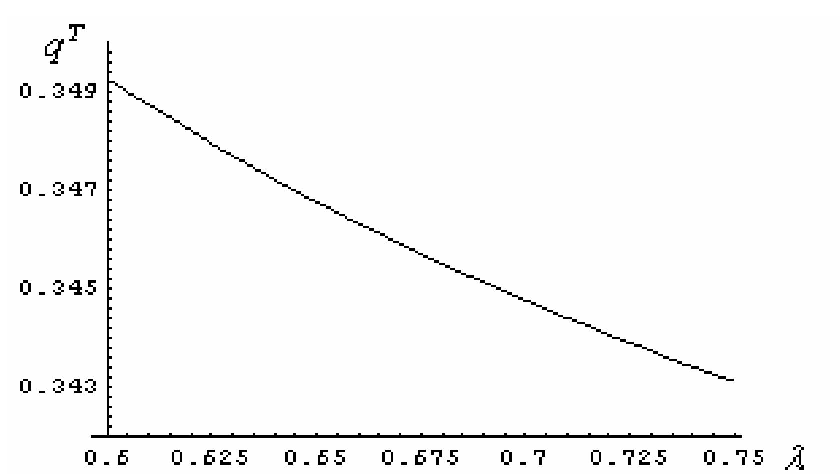

Figure 2. The impact of λ on q^T, when $\eta = 0.5$, $c = 1$, $\overline{p}^x = 0.7$, $\alpha = 0.25$, and $\lambda \in (0.6, 0.75)$.

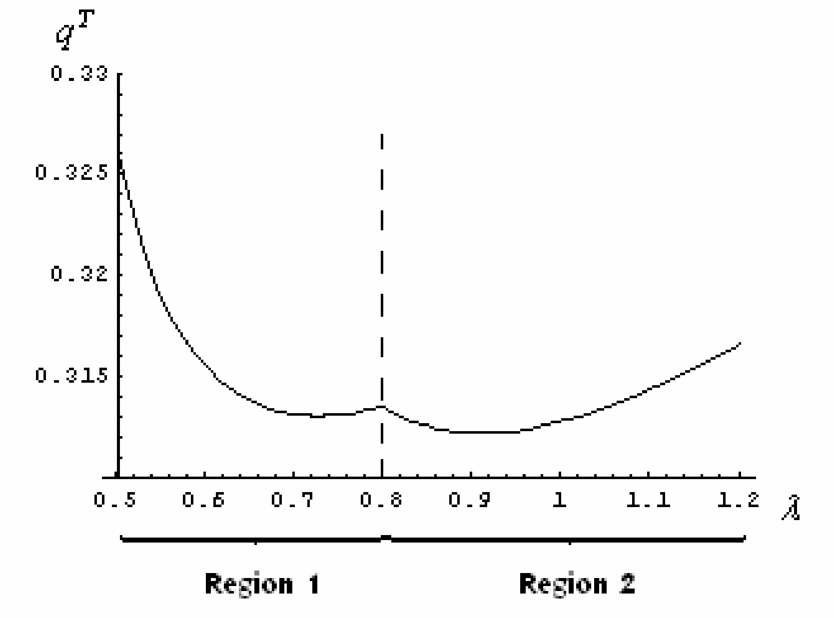

Figure 3. The impact of λ on q^T, when $\eta = 0.5$, $c = 1$, $\overline{p}^x = 0.6$, $\alpha = 0.25$, and $\lambda \in (0.503, 1.2)$.

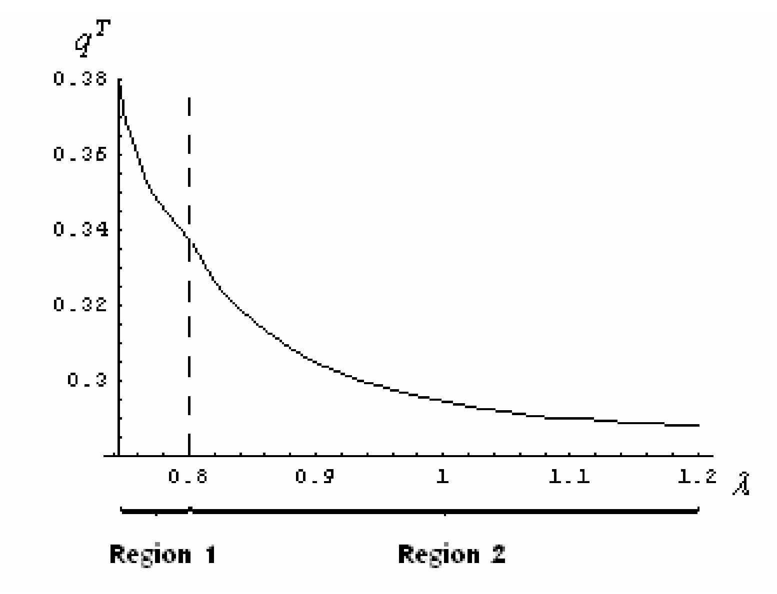

Figure 4. The impact of λ on q^T, when $\eta = 0.5$, $c = 1$, $\overline{p}^x = 0.6$, $\alpha = 1$, and $\lambda \in (0.744, 1.2)$.

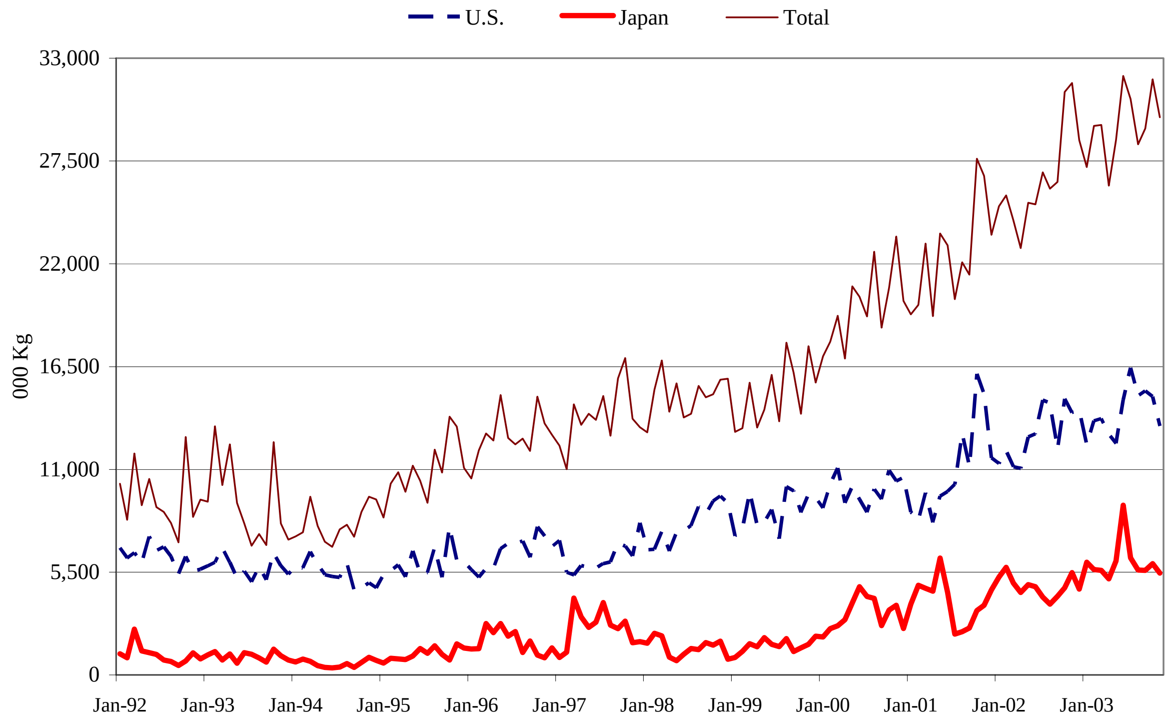

Figure 5. Total monthly pork exports from Quebec and bilateral exports to the U.S. and Japan from January 1992 to November 2003

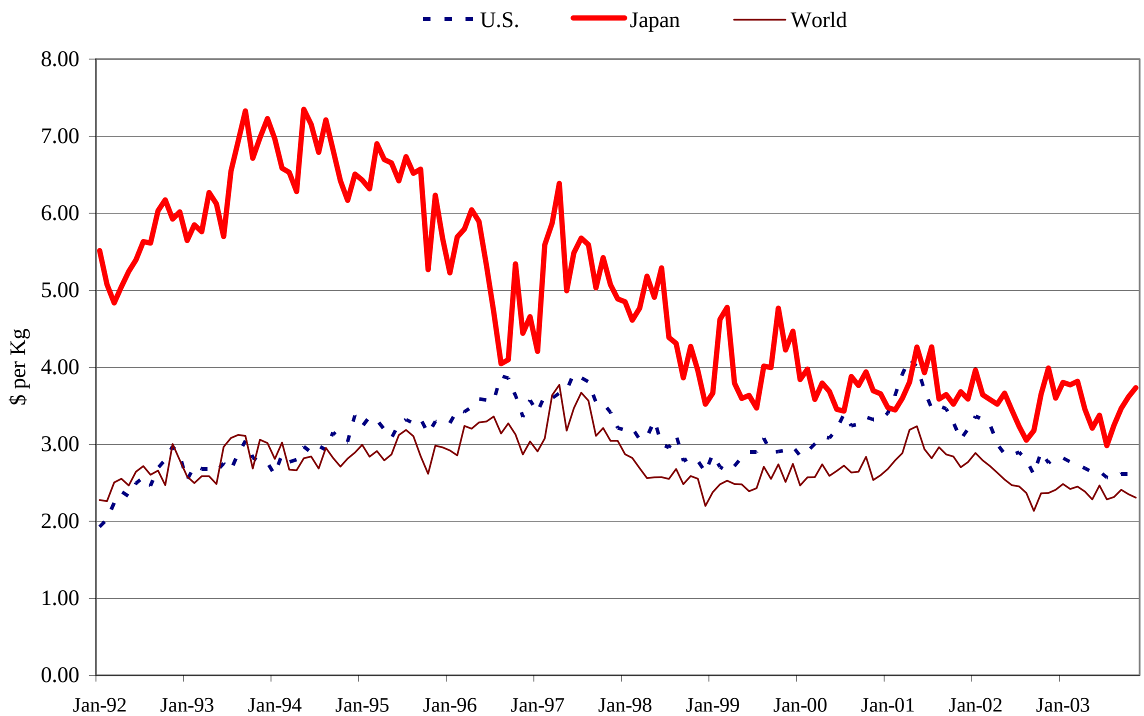

Figure 6. Monthly unit value (in $Can) of Quebec total pork exports and bilateral pork exports to the U.S. and Japan from January 1992 to November 2003

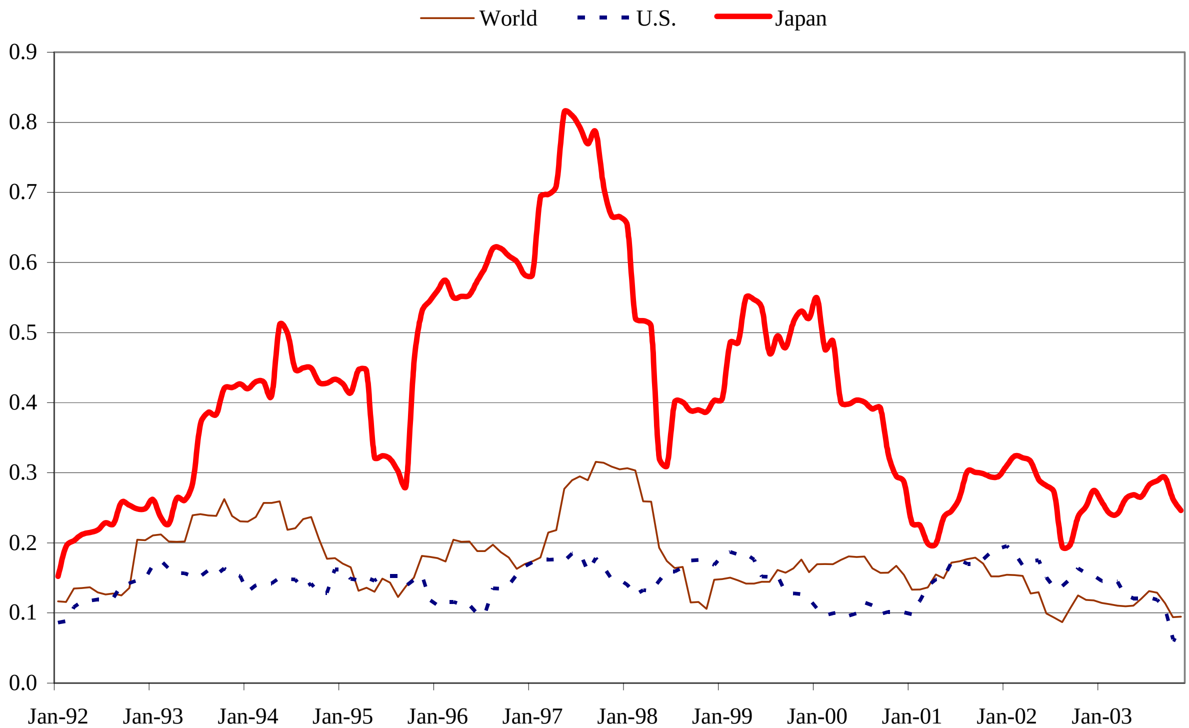

Figure 7. Monthly volatility measure of export unit values for total exports and in the U.S. and Japanese markets from January 1992 to November 2003.

Table 1. Unit root testing

Variables	ADF test		KPSS test	Joint confirmation of a unit root
	Lag	Statistic		
U.S. exports (T)	1	-3.84*	0.44*	No
Japan exports (T)	0	-5.11*	0.14**	No
U.S. real exchange rate (NT)	0	-3.37*	0.30*	No
Japan real exchange rate (T)	0	-4.61*	0.22*	No
U.S. 12-month Vol (T)	1	-2.12	0.09	No
Japan 12-month Vol (T)	0	-2.62	0.56*	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).

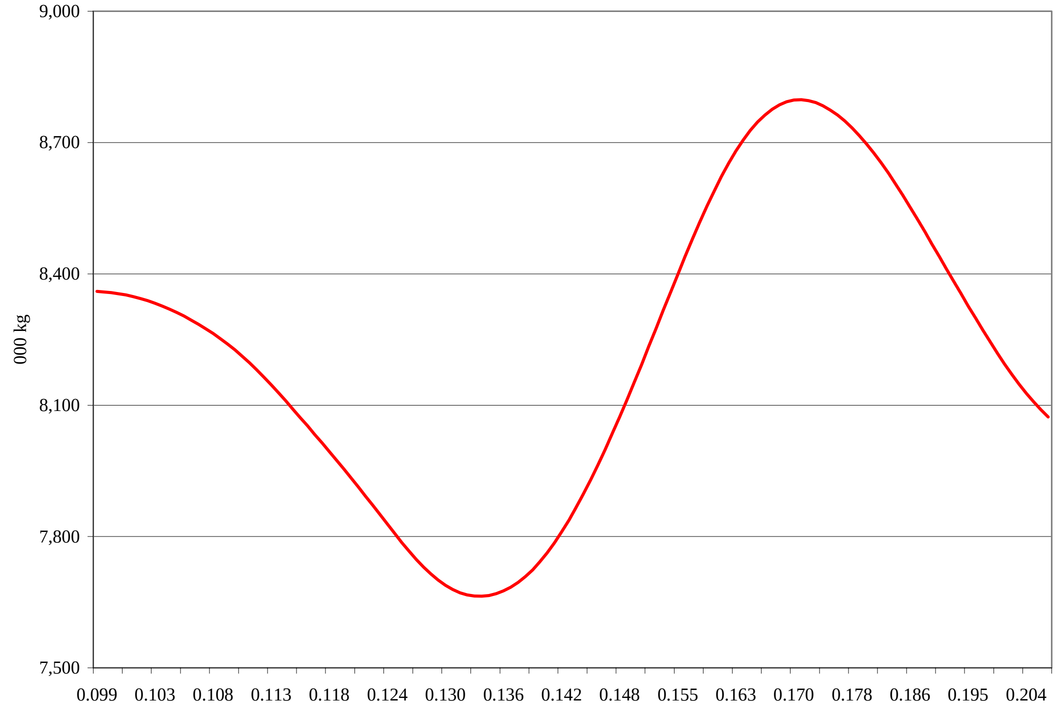

Figure 8. Impact of the lagged export price volatility on exports to the U.S. holding all other independent variables at their sample mean.

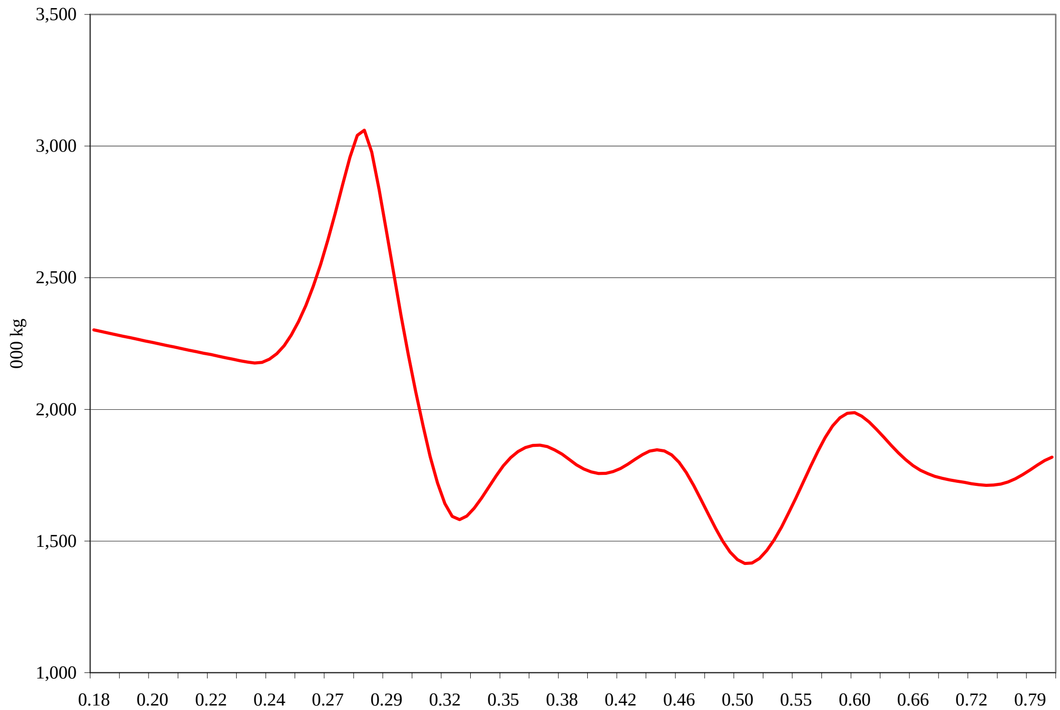

Figure 9. Impact of the lagged export price volatility on exports to Japan holding all other independent variables at their sample mean.

Endnotes

[1] The two largest pork processors have recently announced their intention to merge. The new firm will have a market share of about 70%, assuming that the merger is approved by the Competition Bureau of Canada.

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

[3] At this stage, the choice variable of the processor 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, introducing oligopsonistic behavior would unduly clutter the analytical model because it would involve equilibria in mixed strategies. Franke (1991) has assumed away the issue of imperfect competition by assuming that competition (or rival firms' output) is invariant to volatility.

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

[5] The price difference in the domestic and foreign markets leads to arbitrage opportunities but we implicitly assume that the existence of transaction costs prevents this price difference from dying out.

[6] This result is due to the linearity assumption about domestic demand. Because the slope of the domestic marginal revenue is constant, an increase in the mean preserving spread parameter does not change the expected marginal revenue of the firm given $\theta^e < \theta$. Adding some convexity in the domestic marginal revenue function would imply that an increase in the mean preserving spread parameter would change the expected marginal revenue of the firm. The λ*-effect* would reappear into the equation and thus an increase in volatility would have an ambiguous effect. This ambiguous effect is identified in proposition 2 when it is assumed that $\theta < \theta^e < \eta$.

[7] Our simulations are based on the assumptions of a linear demand and marginal cost, exogenous terms of trade and no change in the exchange rate from the time capacity is determined and exports are realized.

[8] McKenzie (1999) terms our volatility estimate a measure of "changeableness" in the 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 suffer from one serious flaw in that they are usually estimated over the whole sample and thus include information that is not available to agents.

[9] 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 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.

[10] Hamilton's flexible framework is a single equation framework. Ideally, a capacity choice equation would be estimated along with export equations in a multiple-equations framework. Unfortunately, as for threshold cointegration estimators, Hamilton's single-equation estimation problem is highly non-linear and it cannot easily be extended to account for contemporaneous correlation between equations. However, given that the bulk of the literature relies on single-equation models, comparisons between linear and non-linear models can be made to size up the importance of allowing for non-linearities. We report on such comparisons later on.

[11] A covariance variable was also included in equation (13) to measure the correlation between the volatility of export prices in the Japanese and U.S. markets. The Pearson correlation coefficient was not significant and thus was dropped from the export equation.