Exchange Rate Sensitivity of Fresh Tomatoes Imports from Mexico to the United States

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EXCHANGE RATE SENSITIVITY OF FRESH TOMATO IMPORTS INTO THE UNITED STATES FROM MEXICO

ABSTRACT

The agri-food trade between Mexico and the United States grew substantially after the implementation of NAFTA in 1994. While some analysts argue that NAFTA has contributed the most to the dramatic expansion of this trade, others have emphasized the role played by the exchange rate in this process. An attempt is made in this paper to address this issue by quantifying the effects of NAFTA, the Mexico-US exchange rate changes and its volatility on the fresh tomato imports into the United States from Mexico using the maximum likelihood cointegration analysis. The results from the cointegration analysis show that while changes in exchange rate have a positive effect on trade flows, volatility of the exchange rate has a significant negative effect on trade flows. The results from the error-correction model show that both NAFTA and the exchange rate have significant positive influence on fresh tomato imports in the United States from Mexico. However, the effects of exchange rate changes outweigh the effects of NAFTA in the short-run.

Key words: Exchange rate, exchange rate volatility, fresh tomatoes trade, NAFTA, cointegration analysis and error-correction modelling,

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INTRODUCTION

Since the initiation of economic and trade liberalization in the mid 1980s, agri-food exports from and imports into Mexico both in terms of volume and value have grown dramatically. Mexican trade sector in general and the agri-food trade in particular, received a significant boost after the successful implementation of the North American Free Trade Agreements (NAFTA) in 1994. In 1995, Mexico switched from the fixed to a floating exchange rate system. It is widely believed that all these macroeconomic developments have shaped the trade performance of Mexican agri-food sector during the last two decades. However, the relative contributions of trade liberalization and other developments in Mexico are unknown.

While some studies of the Unites States-Mexico agricultural trade have highlighted the importance of NAFTA in the increase in agri-food exports and imports (Rosenzweig, 1996; Schwentesius and Gomez, 2001), others have emphasized the role played by exchange rate changes in enhancing Mexican trade performance (Malaga et al., 2001; Mora-Flores et al., 2002). Since the shift from the fixed to a floating exchange rate system in Mexico, the peso/dollar exchange rate is characterized by periods of unexpected calm followed by episodes of high volatility, which may also have influenced agri-food trade flows between these two countries. Therefore, as the border between Mexico and the Unites States became increasingly open due to NAFTA, Mexican agricultural trade may have also been influenced by changes in exchange rate and exchange rate volatility. The available literature on the effects of Mexico-Unites States exchange rate movements on agri-food trade provide some anecdotal evidence that changes in exchange rate have an effect on trade. To the best of our knowledge, no rigorous attempt has been made to determine empirically, the extent to which

NAFTA, the changes in exchange rate and exchange rate volatility have contributed to the expansion of Mexican agri-food trade towards the United States.

An attempt is made in this article to bridge this gap in the existing literature by determining the extent to which NAFTA, the changes in Mexico-U.S. exchange rate and exchange rate volatility have contributed to the growth in fresh tomato exports from Mexico to the United States during the last two decades. Fresh tomato is the most important vegetable crop exported to the United States from Mexico. It represents about 37% of the value of all vegetable exports and 16% of total agricultural exports from Mexico. While fresh tomatoes cross the Mexico-U.S. border in both directions, Mexico is a net exporter of fresh tomatoes to the United States. The market share of fresh tomatoes from Mexico almost doubled from about 16% of total U.S. consumption in 1991-93 to 29% during 2003-2005. While Mexico has a large and growing domestic market for fresh tomatoes, due to climatic conditions favoring year-round production of vine-ripe tomatoes and growing cooperation among growers and shippers regionally and internationally, Mexico will remain a major exporter of fresh tomatoes and other vegetables to the United States (USDA, 2007).

The trade literature related to the effects of exchange rate movements on agricultural trade has focused on whether exchange rate matters to agricultural trade flows. A large number of empirical studies, except Batten and Belongia (1986) and Fuller et al., (1992), have found that exchange rate does matter. They conclude that it is a key variable in explaining trade performance (Anderson and Garcia, 1989; Chambers and Just, 1979; Mora-Flores et al., 2002; and Guzel and Kulshreshtha, 1995; Cho et al., 2002; Orden and Fackler, 1986; Dorfman and Lastrapes, 1996). While most of these studies find a significant positive effect of exchange rate on trade flows, the magnitude varies considerably across countries and commodities

The shift from a fixed to a floating exchange rate system has led to increased exchange rate volatility (Liang, 1998). Despite a general agreement among economists that flexible exchange rate period has been characterized by a high level of volatility, the theoretical and empirical research so far has provided contradictory evidence of the effects of exchange rate volatility on international trade flows (Cho et al., 2002). Studies by Kenen and Rodrick (1986); Cushman (1988), Anderson and Garcia (1989), Lastrapes and Koray (1990), Qian and Varangis (1994), Cho et al. (2002) have found a negative relationship between exchange rate volatility and trade flows. On the other hand, studies by Asseery and Peel (1991 and Kroner and Lastrapes (1993) suggest that a positive relationship exists between exchange rate volatility and trade flows. The controversial findings are attributed to different specification of the volatility measure, different data sets and different estimation methods employed in these studies.

Only a few studies made systematic attempt to determine the effects of exchange rate changes on Mexican agri-food trade. They found either a positive effects or no effect of exchange rate on trade flows (Table 1). Except one, these studies did not investigate the effects of exchange rate volatility on trade flows. Finally, none focused on the effects of exchange rate on trade flows of fresh tomatoes even though it is arguably the most import crop exported from Mexico to the United States. In cognizant of the limitations of previous studies, this article focuses on fresh tomato exports from Mexico to the United States from January 1989 to December 2004. Unit root tests are used to determine the nonstationarity in data while Johansen's maximum likelihood cointegration analysis is employed to determine the long-run effects of all relevant variables on trade flows. Finally, the error-correction model is developed and estimated to determine the short-run effects of all relevant variables on exports of fresh tomatoes from Mexico to the United States.

The rest of the paper is organized as follows. The following section provides a brief introduction to the Mexican tomato industry, its structure and the importance of fresh tomato trade in Mexico. Section three provides a framework for analyzing the effects of exchange rate changes and exchange rate volatilities on agricultural trade flows. Section four provides an exposition of the empirical model and describes data used in this study. Section five deals with the unit root tests, the specification and estimation of the exchange rate volatility and the main results from the cointegration analysis. Section six discusses the results from the error correction models and focuses on some policy implications of the results for the tomato industry in Mexico. The final section summarizes the main results and concludes the paper.

MEXICAN TOMATO INDUSTRY AND FRESH TOMATO TRADE: AN OVERVIEW

The agricultural sector is an important component of the economy in Mexico, not only in terms of employment, income generation and supply of food, but also in terms of its contribution to the national GDP (Bank of Mexico-INEGI, 2005). The most important agricultural sub-sector in terms of value is fresh vegetable which generates almost 20 percent of the total value of agricultural production in Mexico. The United States is the most important agri-food trade partner of Mexico. The United States buys roughly 85 percent of Mexican exports and supplies about 65 percent of Mexican imports in this category (USDA-ERS, 2005). The exports of horticultural products from Mexico accounted for about 55% of total agricultural exports between 1993 and 2005.

Fresh tomatoes represent the most important Mexican agricultural export to the United States. The value of fresh tomato exports rose dramatically from about $200 million US in 1989 to almost $1 billion US in 2006. Exports of fresh tomatoes account for about 16% of all agricultural exports and 32% of total horticultural exports from Mexico (Bank of Mexico-INEGI, 2005).

While tomato production takes place all over Mexico, only six states (Sinaloa, Baja California, San Luis Potosi, Jalisco and Nayarit) contribute more than 70 percent of the value of production. Fresh tomatoes are produced for two distinct market destinations: domestic market which absorbs about 80 percent of total production and the export market consisting of the United States and Canada. The bulk of the supply in the domestic market is marketed through three regional wholesale markets. In the export market, some multinationals act as the grower-shipper. However, several large Mexican growers have expanded their fresh tomato export operations in recent years. While Mexico is normally considered a source of fresh tomatoes, not a destination, the United States is the leading external supplier of tomatoes to Mexico (USDA-ERS, 2003). As the NAFTA economies become more closely integrated, channels of two-way distributions are likely to expand in the near future. The trend in Mexican fresh tomato exports to the United States since 1989 clearly indicates the movement in this direction (Figure 1).

Tomato exports from Sinaloa, Mexico, directly compete with tomatoes produced in South Florida. Mexican shippers in Sinaloa produce mainly the extended-shelf-life tomatoes that are harvested as vine-ripped. It helped them to gain additional market share in the U.S. during the last two decades. On the other hand, the production of Baja California competes with the production of California that goes to the market from May to December. Mature-green tomatoes are produced domestically in both Florida and California. Florida tomatoes are shipped mostly to eastern and mid-western markets, while the western half of the U.S. is served primarily by tomatoes from California and Mexico (Love and Lucier, 1996).

The challenge to supply seasonal, perishable products year-round has favoured Mexican exports of fresh tomatoes and increased horizontal and vertical integration among shippers regionally, nationally and internationally (Cook, 2001). In the grower-shipper combination, the shipper often joint ventures with growers to ensure the needed supply and then markets the produce for a fee. The shippers also advance cartons and sometimes, control the harvest operations and impose desirable grades and standards to generate a consistent quality of fresh tomatoes. The marketing services from production points in Mexico to consumption centres in the US add value to the final consumers of fresh tomatoes. While some multinationals act as the grower-shipper, several large Mexican growers have expanded their operations in recent years. Fewer than a thousand farms dominate fresh tomato production in North America and fewer than 50 shippers control the first marketing stage as tomatoes move into the wholesale, retail and food service sectors. In many instances, these shippers are also the growers (Padilla-Bernal and Thilmany, 2003).

AN ANALYTICAL FRAMEWORK

An expected utility maximization model involving a single commodity, fresh tomatoes, and two countries, Mexico and the United States, is used to shed light on the theoretical relationships between trade flows and a set of explanatory variables. The specification of the economic model is followed by the derivation of a set of comparative statics to determine the direction of the effects of the key explanatory variables on import demand.

In this model, production theory is used to derive the import demand function by treating imports as inputs. It is assumed that import and export decisions are made by profit maximizing firms which operate under perfect competition in commodity and factor markets. These firms use imports with domestic inputs to produce output that can be absorbed at home or exported. Following Appelbaum and Kohli (1997), an attempt is made to model import demand under exchange rate uncertainty.

Let $q = f(X_L, X_M, K)$ be a neoclassical production function. Where X_L is the input labour, X_M is the quantity of the imported input needed to produce output q and K is capital. It is assumed that the only source of risk is the uncertain exchange rate. Therefore the exchange rate is a random variable and by implication, so is the foreign price and profits. The profit maximization problem of the firm can be represented as:

$$Max_{X_L, X_K, X_M} \quad E \;\; U \;\; (\bar{R} \qquad _1)P \;\; f(X_L, X_M, K) - w_L X_L - (\bar{R} + \;\; _M) w_M X_M - w_K K \qquad (1)$$

where $R \quad \bar{R} \qquad _i$, the currency price of imports (foreign currency) and θ is a random variable distributed according to the density function *g*(θ), with *E(θ)* = 0, (so that *E(R)* = $\bar{R}$) and *Var(R)*= *Var(θ)* = σ^2 . It is assumed that $U'(\;)$ is a Von-Newman-Morgestern utility function with $U'(\;) > 0$. The solution to the firm's problem defines the (dual) indirect (expected) utility function *V* which is represented in equation (2) as:

$$Max_{y,x} E\, U\, (\bar{R} \qquad _1)P \;\; y \quad w_L X_L \quad (\bar{R} \qquad _M) w_M X_M \quad w_K K \;: y \quad F(x) \quad V(w_i, P, \bar{R}, \; _i, \;\;) \quad (2)$$

where ρ represents higher moments of *g(θ)* and the random variable *θ* is continuous and convex to the moments (Appelbaum, 1993). The firm's demand and supply function can be obtained from the above indirect expected utility function by applying the envelope theorem. Hence:

$$\frac{V}{w_i} \qquad E\, U'(\;\;) \quad x_i \text{, and} \qquad \frac{V}{\bar{P}} \qquad E\, U'(\;\;) \quad y \,.$$

Therefore, the firm's input demand and output supply functions are given by

$$x_i \quad \frac{V}{w_i} / \frac{V}{w_K} \text{ , and} \qquad (3)$$

$$y \quad \frac{V}{\bar{P}} / \frac{V}{w_K} \;. \qquad (4)$$

By maximizing expected utility of profits the first order conditions are obtained as

$$EU'(\;\;) \;\; (\bar{R} \qquad _1)P \;\; f_L(X_L, X_M, K) - w_L \qquad 0 \text{ , and} \qquad (5)$$

$$EU'(\;\;) \;\; (\bar{R} \qquad _1)P \;\; f_M(X_L, X_M, K) - (\bar{R} + \;\; _M)\, w_M \qquad 0 \,. \qquad (6)$$

The above conditions can be rewritten as:

$$\bar{R} \quad {}_{1} \; P \;\; f_L(\mathrm{X_L}, \mathrm{X_M}, \mathrm{K}) = \mathrm{w_L} \text{, and} \tag{7}$$

$$(\bar{R} \quad {}_{1})P \;\; f_M(\mathrm{X_L}, \mathrm{X_M}, \mathrm{K}) = (\bar{\mathrm{R}} + \quad {}_{M}) \;\; w_M \quad . \tag{8}$$

The term $(\bar{R} \quad {}_{M}) \;\; w_M$ is the full marginal cost of imports and θ_M represents the marginal cost of uncertainty. It follows from (8) that in the presence of uncertainty the value of the marginal product of imports will deviate from the expected marginal cost of imported products. Particularly under risk-aversion, θ is positive so the value of the marginal product of imports will exceed their expected market price. It implies that the quantity of imports will be smaller under uncertainty. The solution to the system (7) and (8) is:

$$X_L^* \quad X_L^*(w_L, w_M, p, \bar{R}, \; {}_{1}), \tag{9}$$

$$X_M^* \quad X_M^*(w_L, w_M, p, \bar{R}, \; {}_{M}). \tag{10}$$

These relations indicate the amount of each factor that will be hired as a function of the factor and product price; they are the choice functions of this model. Assuming that it is possible to solve for equations (9) and (10), it becomes meaningful to perform comparative statics of the profit maximization model to know the changes in factor employment due to given changes in prices and exchange rate.

The comparative static results showed that while the import demand (traded volume) responds negatively to exchange rate uncertainty, it responds positively to changes in exchange rate; the price of imports responds positively to changes in exchange rate; demand of domestic input responds positively to changes in output price; demand of labor responds negatively to changes in its price, and demand of imported inputs responds negatively to changes in its price. These results are used to guide the empirical analysis in this paper.

EMPIRICAL ISSUES, DATA AND ESTIMATION

While both the structural approach and the reduced form approach have been used in the literature to determine the effects of exchange rate on trade flows of agri-food commodities, a reduced form model was used in this study for the simplicity of accommodating data nonstationarity and testing meaningful economic relationships. As for the appropriate specification of an import demand function, economic theory does not provide any specific direction on the best functional form to be used in the analysis. Thursby and Thursby (1984) examined nine most commonly used specifications of aggregate import demand functions for the United States and found the log-linear form to be better than others. Thus, for fresh tomatoes imported in the United States from Mexico, the following equation is estimated in log-linear form.

$$\boldsymbol{Q_t^* = \beta_0 + \beta_1 USY + \beta_2 Pt + \beta_3 Pus + \beta_4 W + \beta_5 ER + \beta_6 V + \beta_7 D_i + \varepsilon_i} \qquad (11)$$

Where, the following relationships are expected, *a priori*:

$$\frac{Q^*}{USY} \; 0 \qquad \frac{Q^*}{Pt} \; 0 \qquad \frac{Q^*}{Pus} \; 0 \qquad \frac{Q^*}{W} \; 0 \qquad \frac{Q^*}{ER} \; 0 \qquad \frac{Q^*}{V} \; 0 \qquad \frac{Q^*}{D_i} \; 0$$

Where Q* is quantity of fresh tomatoes imported, USY is the US per-capita income, W is the ratio of United States-Mexico farm wage rates, Pt is the border price of fresh tomatoes, Pus is the price of a substitute (the tomato price in the Unites States), D_i is NAFTA and seasonal dummy variables to indicate the effects of liberalization, ER is the Mexico-Unites States exchange rate, V is a measure of exchange rate volatility. The expected signs of all relevant variables are given immediate after the specified import demand function.

The econometric estimation conducted in this article consists of four steps. The first step is to determine whether the relevant data are level or difference-stationary. It is followed by the estimation of the volatility measure of the exchange rate, which is incorporated as an explanatory variable. Then, the cointegration analysis is performed to estimate the long-run relationship between fresh tomatoes imported and a set of explanatory variables. Finally, a VEC model is specified and estimated to determine the short-run effects of exchange rate movements, exchange rate variability and NAFTA on fresh tomatoes imported in the United States from Mexico. Since the maximum likelihood cointegration analysis relies on a VAR model and the results from VAR formulations are sensitive to the lag-length choices, (Haffer and Sheenan, 1991), special emphasis is given in this research to determine the optimal lag structure of the VAR model.

To formally evaluate the relevance of difference stationarity in data, the Augmented Dickey and Fuller (1979) test (ADF) is used in this research. To test for the presence of a unit root in each data series, the ADF test was computed by running the following regression:

$$(1-L)y_t = \alpha_0 + a_1 y_{t-1} + \sum_{i=1}^{p} \gamma_i (1-L) y_{t-i} + \varepsilon_t . \qquad (12)$$

A negative and significant estimate of a_1 is inconsistent with the null hypothesis of a unit root in y_t. The t-ratio, however, does not have a standards t-distribution in this framework. Therefore, the critical values provided by Dickey and Fuller (1979) are used.

While it is generally recognized in the trade literature that exchange rate volatility induces additional uncertainty in trade, no consensus exists on how to measure it. Economic theory does not provide any clear guidance on the features of the most appropriate volatility measure (Clark, et al. 2004). As indicted by Bollerslev et al. (1992), financial time series are typically heteroskedastic, leptokurtic and exhibit volatility clustering. He suggests that these features could be handled more successfully by modeling the volatility of the time series as conditional on its past behaviour. The ARCH model introduced by Engle (1982) and its generalization, the GARCH, have proved to be very useful for modeling exchange rate volatility. These models have been used in a number of recent studies to measure exchange rate volatility with monthly data in a few recent studies (McKenzie, 1998; Lastrapes and Koray, 1990; and Qian and Varangis, 1994). Engle (2001) provides an excellent survey of this literature. Owing to the growing popularity of this approach, a volatility measure based on a GARCH model is used to measure exchange rate volatility in this research. The general form of the GARCH model used to estimate the volatility of the exchange rate is as follow:

$$ {}_{t}^{2} \qquad \overset{q}{\underset{i\ 1}{}} \ {}_{i}\ {}_{t\ i}^{2} \qquad \overset{p}{\underset{i\ 1}{}} \ {}_{i}\ {}_{t\ 1}^{2} \qquad (L)\ {}_{t}^{2} \qquad (L)\ {}_{t}^{2}\ , \qquad (13) $$

Where, ${}_{t}^{2}$ is the conditional variance, ε_{t-i} is the moving average term, and e_t is the error term.

If there is a unit root in each series, a precondition for the existence of a stable steady-state relationship is cointegration among the variables. A vector of variables is said to be cointegrated if each variable in the vector has a unit in its univariate representation, but some linear combination of these variables is stationary (Engle and Granger 1987). At least five alternative approaches for testing cointegration have been advanced in the literature, but the approach developed by Johansen (1988, 1991) possess some interesting features. For example, it derives maximum likelihood estimators of the cointegrating vectors for a VAR system. It extends the Engle-Granger procedure to a multivariate context where there may exist more than one cointegrating relationship among a set of n variables. Moreover, it allows for testing meaningful economic hypothesis. Gonzalo (1994) and Hubrich et al., (2001) used Monte Carlo Simulations to demonstrate that the maximum likelihood procedure outperforms all other cointegration methods. Because of these attractive features, Johansen's approach is used in this research. What follows next is a brief exposition of the maximum likelihood cointegration procedure.

Following Johansen (1988; 1991) and Johansen and Juselius (1990), this approach starts with a k_{th} order unrestricted VAR representation of X_t such that:

$$ X_t = c + \pi_1 X_{t-1} + \pi_2 X_{t-2} + \ldots . \pi_k X_{t-k} + \quad + \quad TD + \Phi D_t + \varepsilon_t \quad (t = 1, \ldots, T) \qquad (14) $$

Where: X_t = a vector of p I(1) variables; D_t = eleven seasonal dummies; π_i = a (p x p) matrices of parameters; c = a (p x 1) vector of constant terms; TD = trade dummy variables; and ε ~ NID (0, Ω). In general, economic time series are non-stationary processes, and the VAR system like (14) has been expressed in first differenced form. Using =1 – L, where L is the lag operator, the model in equation (14) can be reparameterized as:

$$X_t = C + \Gamma_1 \quad X_{t-1} + \Gamma_2 \quad X_{t-2} + \ldots + \Gamma_{K-1} \quad X_{t-k+1} - \Pi X_{t-k} + \Phi D_t + \varepsilon_t \qquad (15)$$

where:

$\Gamma_i = - I + \pi_1 + \ldots\ldots.. + \pi i$, and $- \Pi = I - \pi_1 - \pi_2 - \ldots. - \pi_k$ with $i = 1,2,\ldots.. , k-1$

It is interesting to note that the reparameterized model is a traditional first-differenced VAR model except for the term ΠX_{t-k}. The coefficient matrix of X_{t-k}, Π, contains information about the long-run relationships among variables in the model. If Π has a full rank, then X is a stationary process. In this case, a non-differenced VAR model is appropriate. If Π has a zero rank, then Π is a null matrix and X_t is an integrated process. Only in this case, a traditional first-difference VAR model is appropriate (Hamilton, 1994). If, however, 0 < (rank (Π) = r) < p, cointegration holds and Π can be represented as the product of two p x r matrices and , such that Π = . The s are the cointegrating vectors and s are the weights. In this case, X_t is stationary. The long-run equilibrium is unique when r =1.

The maximum likelihood estimation of Π consists of two sets of regressions. One set generates the residuals R_{0t} from the regression of X_t on X_{t-1}, …., X_{t-k+1}, and the other set generates R_{kt} from the regression of X_{t-r} on $X_{t-1,\ldots.,}$ X_{t-k+1} (Johansen, 1995). The concentrated likelihood function can be expressed as:

$$L(\ ) = |\ \hat{}(\ )|^{T/2} \ |S_{00} \quad S_{ok} \ (\ S_{kk} \ )^{1} \quad S_{k0}|^{T/2} \ \hat{\Omega}(\beta)|^{-T/2} \qquad (16)$$

where $S_{00,}$ $S_{0k,}$ S_{k0} and S_{kk} are the product moment matrices of the residuals defined as:

$$S_{ij} = \quad ^{1} \sum_{t\ 1}^{T} R_{it} R_{jt} \qquad _{i,j} = 0, k$$

It is clear from equation (16) that maximizing the concentrated likelihood function is equivalent to minimizing $|\hat{\Omega}(\ )|$. This minimization amounts to solving the following eigenvalue problem:

$$|\ \ I \quad C^{\ 1} S_{k0} S_{00}^{\ 1} S_{ok} C^{\ \ 1}\ | = 0 \tag{17}$$

where C is a (p x p) matrix such that $S_{kk} = C\ C$. The vector of eigenvalues is given by λ while the corresponding eigenvectors can be derived as $v_i = C^{\ \ 1} ei$, where *ei's* are the eigenvectors from equation (17). The estimates of and Ω can be obtained by using the estimated value of *β*. The null hypothesis that there are r cointegrating vectors is tested using two likelihood ratio tests called the trace test and the maximum eigenvalue test. The critical values for these tests are generated through simulation and are reported in Johansen and Juselius (1990), Osterwald-Lenum (1992) and MacKinnon (1999).

Since the introduction of the error correction approach by Sargan (1964) to the economic literature, the use of error correction models has proved to be a useful tool in linking the long-run results to the short-run dynamics of an economic process. In the maximum likelihood cointegration framework, the long-run equilibrium relationships are used to impose constraints on the short-run dynamics in the ECM in economically meaningful ways. After estimating the long-run cointegration relations using the Johansen approach, it is possible to reformulate equation (15) and estimate the vector error-correction (VEC) model with the error-correction term(s) included in it. The unconstrained VEC model is:

$$Z_t \quad \sum_{j\ 1}^{k\ 1} \ _j\ Z_{t\ j} \quad Z_{t\ k} \quad D_t \qquad _t \tag{18}$$

where Π = , with *β* being a matrix of long-run coefficients and represent the speed of adjustment to disequilibrium. For the estimation of the VEC model, the vector of error correction terms obtained from the cointegrated relationship is incorporated in equation (18) to obtain an empirical version of the model as follows:

$$Z_t \quad \sum_{j\ 1}^{k\ 1} {}_j\, Z_{t\ j} \quad ({}_1\tilde{Z}_{t\ 1}) \quad D_t \qquad {}_t \tag{19}$$

Where the terms $\sum_{j\ 1}^{k\ 1} {}_j\, Z_{t\ j}$ and $({}_1\tilde{Z}_{t\ 1})$ are the vector autoregressive (VAR) component in first differences and error-correction components, respectively. In this formulation, Z_t is a p x 1 vector of I(1) variables and ${}_j$ is a p x p matrix that represents short-term adjustments among variables across p equations at the *j*th lag. The parameter ${}_1$ represents a (p x r) cointegrating vector present in the system and Δ denotes first differences. The parameters in represent the speed at which equilibrium error is corrected in this system, D_t is a set of monthly dummy variables, is a p x 1 vector of constant; and ${}_t$ represents a px1 vector of white noise error terms. The Error Correction Model as specified in equation (19) is estimated to obtain the short-run relationships among the variables in the model.

The estimation of the error correction model follows Hendry’s (1987) "general-to-specific" methodology. This process helps to develop a parsimonious specification of the VEC model unique for the selected commodity (Brooks, 2002). A negative and statistically significant coefficient of the ECT re-reconfirms cointegration relationships obtained earlier.

This article focuses on the 1989-2004 period. Relevant data were obtained based on their codes under the Harmonized Commodity Description and Coding System (Harmonized System). Monthly data from January 1989 to December 2004 have been used to estimate the fresh tomato import demand function for the United States. This period was selected based on the availability of data. Table 2 describes the main characteristics of the data used in this analysis and their sources.

RESULTS AND DISCUSSION:

Since the cointegration analysis is meaningful when the relevant data are characterized by unit root nonstationarity, it is important to determine first if the data contain unit root. We employed the ADF test and the results are presented in Table 3. Akaike's final prediction error (FPE) criterion is used to determine the optimal lag-length for each series. For all the variables, the null hypothesis of a unit root cannot be rejected when the data are in their level form. However, the null hypothesis is soundly rejected for each of the variables when the series is first-differenced. Thus, all variables in the model for fresh tomatoes are I(1).

The exchange rate data used to estimate the GARCH model consist of the difference in the log of the real exchange rate with respect of the previous period. For this model, a test for non-normality and a test for autocorrelation were conducted. Looking at the normality test, the model shows non-normal errors since the Jarque Bera test is significant at 5 percent level. Thus the residuals are highly leptokurtic. Furthermore, the high values of kurtosis (greater than 3) and negative skewness suggest the presence of ARCH residual in the data. Additionally, since the standardized residuals have a mean value close to zero (0.0085) and variance close to one (0.9914), it is an indication that the GARCH (1,1) model describes the data and the residuals are independently and identically distributed (Bollerslev, 1986). The resulting conditional variance function is specified in the following equation:

$$\hat{h}_t = 0.00059 + 0.0031\hat{\varepsilon}^2_{t-1} + 0.967\hat{h}_{t-1} \qquad (20)$$

The predicted values from this equation closely resemble the movement exchange rate and are used as the measure of Mexico-US exchange rate volatility in this article (see Figure 2).

Since all relevant variables in fresh tomato import demand function are integrated of order one, we employed Johansen's maximum likelihood cointegration analysis to determine if there is any long-run relationships among these variables. As this approach relies on an unrestricted VAR which is sensitive to the number of lags of each variable in the system, it is important to determine the appropriate lag-length for the VAR system. We employed Sims' modified likelihood ratio test to determine the appropriate lag-length in this article. The results presented in Table 4 suggest an appropriate lag to be 9. Table 5 reports the results of cointegration analysis. Both the trace (0.95) and the maximum eigenvalue (0.90) statistics reject the null hypothesis of no cointegration. While the trace statistic suggests the existence of two cointegrating vectors, the maximum eigenvalue statistic indicates the presence of only one. Consequently, one cointegrating vector was chosen in this article. This vector represents the long-run relationship among the variables and is given by:

$$Qt = -0.285Pt + 0.844Pus + 0.246USY + 0.589ER - 0.460VG + 0.505W1. \qquad (21)$$

Equation (21) shows that all variables have coefficients with theoretically expected signs. Since the coefficient of own price variable is less than one, it implies that the import demand for Mexican fresh tomatoes in the U.S. is inelastic. Based on the positive sign of the related price, tomatoes produced in the United States are substitute for tomatoes imported from Mexico. Also the income elasticity is less than one.

The positive value of the exchange rate variable indicates that a devaluation of the Mexican peso exerts a positive long-run effect on the United States imports of fresh tomatoes from Mexico. Since the exchange rate volatility has a negative sign, it implies that higher volatility discourages fresh tomato imports in the U.S. from Mexico in the long-run. Finally, as the difference between rural-wage rates in the U.S. and in Mexico increases, Mexican fresh tomato imports into the United States increases. The estimated weights ($'s$) for imports of fresh tomatoes suggest that if an external shock upsets the long-run equilibrium, trade volume would respond faster than any other variables to bring this system back to the long-run equilibrium.

The short-run results of the U.S. import demand for fresh tomatoes from Mexico along with a set of diagnostic statistics are presented in Table 6. The goodness of fit measured by the adjusted R^2 is 0.33 and the F-statistic is significant at 5% level of error probability. As the VEC model includes the lagged dependant variable, the conventional Durbin-Watson test can not be used. Instead, one is required to use the Durbin "h" statistic to test for autocorrelation. While the Jarque-Bera test can't reject the null hypothesis of normality of the residuals, the Lagrange Multiplier (LM) statistic indicates that there is no serial autocorrelation. Therefore, the results of the VEC model for fresh tomatoes are satisfactory.

The own price coefficient exhibits the expected negative sign and is statistically significant at the five percent level. These results are consistent with the results obtained from the long-run cintegration analysis. The own price elasticity is 0.10, suggesting a more inelastic import demand for Mexican fresh tomatoes in the U.S. in the short-run than in the long-run. The price of the related commodity (price of tomatoes produced in the Unites States) also exhibits the expected positive sign (0.195) and it is significant at the five percent level. The coefficient of the income variable is positive (0.11) and significant suggesting a direct relationship between income in the Unites States and quantity demanded of fresh tomatoes from Mexico. The short-run import demand elasticities are smaller than the corresponding long-run elasticities, and hence, are consistent with the *Le Chatelier* principle.

With regard to the exchange rate variable, the model yields a coefficient with a positive sign as expected and the magnitude is 0.66. This result supports the hypothesis that a depreciation of the Mexican currency relative to the Unites States dollar makes imports into the Unites States less expensive resulting in an increase of fresh tomato imports from Mexico. With regard to the volatility of the exchange rate variable, in the short-run, it has a negative and statistically significant influence on the Unites States imports of tomato from Mexico. This result lends empirical support to the comparative statics results performed in this research and suggests that risk-averse importers are discouraged by higher volatility episodes of the Mexico-Unites States exchange rate. To the best of our knowledge, this is the first set of results demonstrating the effects of exchange rate volatility on exports of fresh tomatoes from Mexico to the Unite States.

Regarding the ratio of the Unites States-Mexico farm wage rate (W1), defined as farm wage rate in the Unites States over farm wage rate in Mexico, the VEC model yields results in line with those obtained from the long-run analysis as the coefficient is positive and statistically significant. These results imply that an increase in farm wage rate in the United States relative to that in Mexico will encourage fresh tomato imports from Mexico.

The error correction term yields a negative and significant coefficient. This result reconfirms the presence of a cointegrating relationship in the Unites States import demand function for fresh tomatoes from Mexico. The negative sign of this coefficient indicates that the direction of correction is towards the long-run equilibrium while the size indicates the speed of adjustment towards the long-run equilibrium (Harris and Sollis, 2003).

The coefficient of the free trade dummy variable is positive and significant suggesting that NAFTA had a positive effect on the Unites States imports of fresh tomatoes from Mexico. This is contrary to the findings of previous researchers who argued that NAFTA had no significant effect on the U.S. fresh tomato imports from Mexico. Note, however, while both NAFTA and peso devaluation have statistically significant effects on fresh tomato imports into the United States, the peso devaluation effect dominates that on NAFTA in the short-run. Finally, it appears that the fresh tomato imports in the U.S. from Mexico are not significantly influenced by any seasonal pattern.

Traditionally, exchange rate has been incorporated into trade models by expressing all prices in common currency units or using a composite relative price variable (Carone, 1996). This implies that both foreign price and exchange rate have equal coefficients. This specification is very restrictive and has the potential to bias the analysis (Chambers and Just, 1979). In this article, we used exchange rate as a separate variable in the fresh tomato import demand function. To determine if it was appropriate to do so, an equality restriction on coefficients was formulated. The null hypothesis of the equality of foreign price and exchange rate coefficients is rejected at the five percent level of significance both in the short-run and in the long-run (Table 7). Therefore, the approach followed in this study to incorporate exchange rate as a separate variable into the trade model is appropriate.

Formal hypothesis tests were also performed to determine if exchange rate and the volatility of exchange rate have significant effects on trade flows of fresh tomatoes between Mexico and the United States. For the long-run results, the above hypotheses are implemented using the general model H [8] which reflects relevant linear restrictions imposed on the cointegrating vectors (Johansen, 1995). The dimension of matrix H is a *pxs* where *s* equal to the number of variables minus the number of restrictions and φ *(sxr)* is a matrix of unknown parameters. In each case, a separate H matrix is used to impose relevant restrictions. For the short-run results, the hypotheses are tested using the standard t- and F-statistics. For both the short- and long-run models, the null hypotheses of zero effects the exchange rate and the volatility of the exchange rate on the U. S. imports of fresh tomatoes from Mexico are rejected at the five percent level of significance (Table 7). With regard to the joint test on the regression coefficients of the exchange rate and exchange rate volatility, the null hypothesis is also rejected at the five percent level of significance. These results imply that individual as well as the combined effects of exchange rate and its volatility on the Mexico-United States trade flows of fresh tomatoes are statistically significant both in the long-run and in the short-run. Thus, if either variable is ignored in an empirical analysis, bias due to mis-specification could result.

CONCLUDING REMARKS:

Mexico is an important NAFTA partner and the United States has been the most important agri-food trade partner of Mexico. While the border between Mexico and the United States became increasingly open due to NAFTA, Mexican agri-food trade with the United States has also been shaped by changes in peso-dollar exchange rate and exchange rate volatility. To the best of our knowledge, no formal attempt has been made in recent years to measure the effects of NAFTA, changes in exchange rate, exchange rate variability and other relevant factors on the growth of the U. S.-Mexico agri-food trade. An attempt is made in this article to bridge this gap by estimating the effects of changes in exchange rate, exchange rate volatility, NAFTA and other relevant factors on the U.S. imports of fresh tomatoes from Mexico.

Fresh tomato is the most important export crop in Mexico. The exports of fresh tomatoes from Mexico to the United States have grown significantly since 1989 and Mexico is the largest exporter of vine-ripped fresh tomatoes to the United States. Johansen's maximum likelihood cointegration analysis and vector error correction models are used to determine long run and short run effects of changes in exchange rate and its volatility on fresh tomato imports in the United States from Mexico. The results show that changes in exchange rate, income and the ratio of U.S. and Mexican real farm wages all have significant positive impacts on the U.S. imports fresh tomatoes from Mexico in the long-run. The results from the VEC model suggest that while NAFTA and exchange rate both have significant positive effects on Mexican fresh tomato exports to the United States, the effects of exchange rate outweigh the effects of NAFTA. The exchange rate variability, however, has reduced the volume of fresh tomatoes imported into the United States from Mexico. The results also suggest that when commodity specific disaggregated data are used in the analysis, it is appropriate to incorporate exchange rate as a separate variable in a trade model.

The economic importance of this commodity makes this study important from a policy perspective. For first time, this research provides a set of estimates on how changes in exchange rate and the volatility of exchange rate affect the Mexico-Unites States trade flows of fresh tomatoes. The results suggest that Mexican exporters have the opportunity to make additional profits by responding quickly to changing market conditions when Mexican peso devaluates against the U.S. currency. Since the fresh tomato trade is also sensitive to exchange rate uncertainty, assisting the traders to access forward markets could be helpful.

While this article makes valuable empirical and policy contributions, two of its limitations need to be brought out. Even though the results from the VEC model are consistent with economic theory and are statistically satisfactory, it can only explain 33% of the month-to-month variations in fresh tomato imports in the U.S. from Mexico. While the explanatory power of a VEC model is typically low, a large portion of the total variations remains unexplained. Future research should explore avenues for improving the explanatory powers of such error-correction models. Finally, an alternative econometric approach such as impulse-response and variance decomposition could be employed to contrast the short-run results of this study. Future research along this line will enhance our understanding of the short-run dynamic effects of changes in exchange rate, volatility and other relevant factors on trade flows fresh tomatoes between Mexico and the United States.

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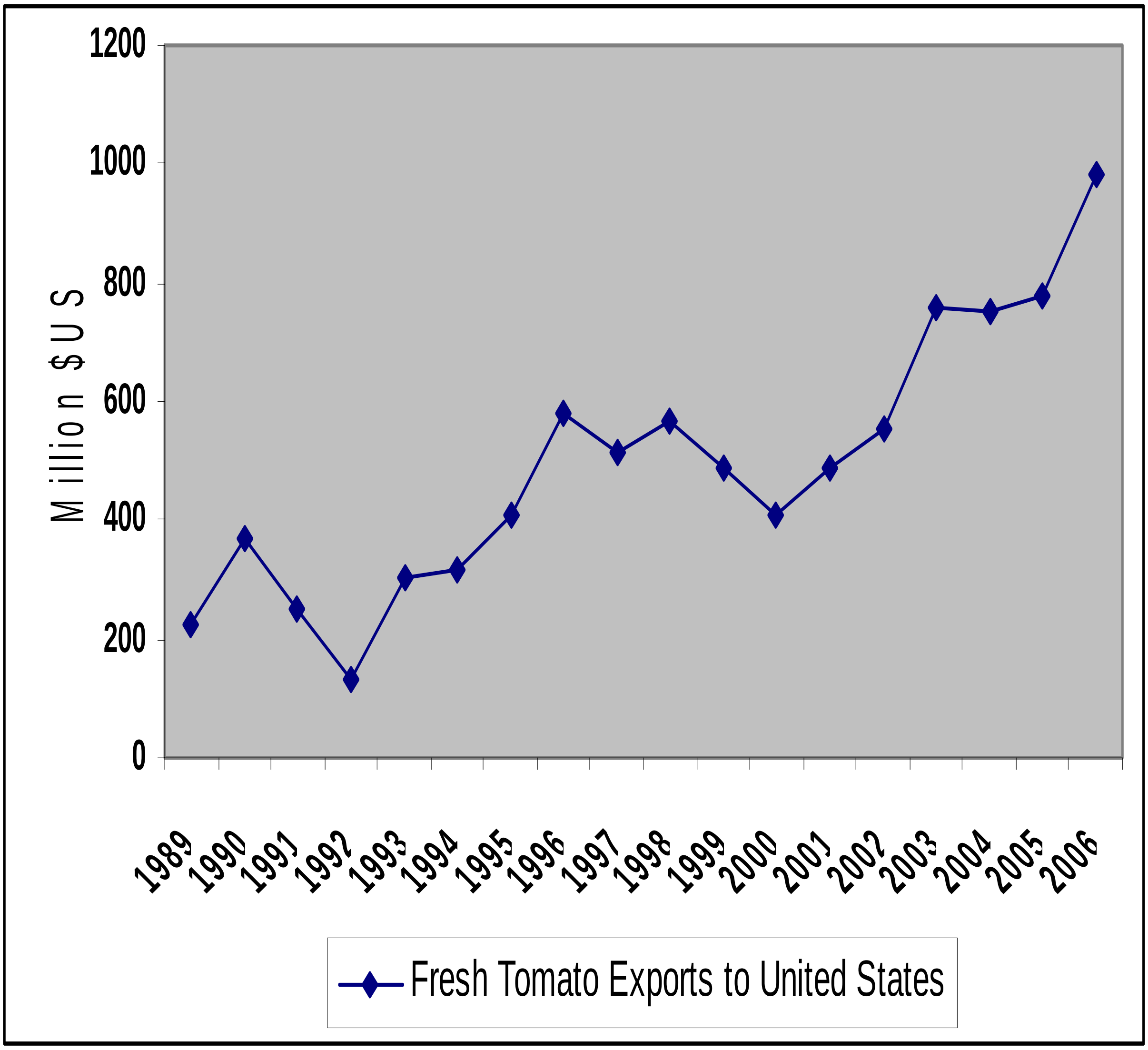
Figure 1. Mexican Fresh Tomato Exports to the Unites States



Source: ERS, USDA

Figure 2. Mexico-US Exchange Rate Volatility from a GARCH (1,1) Model

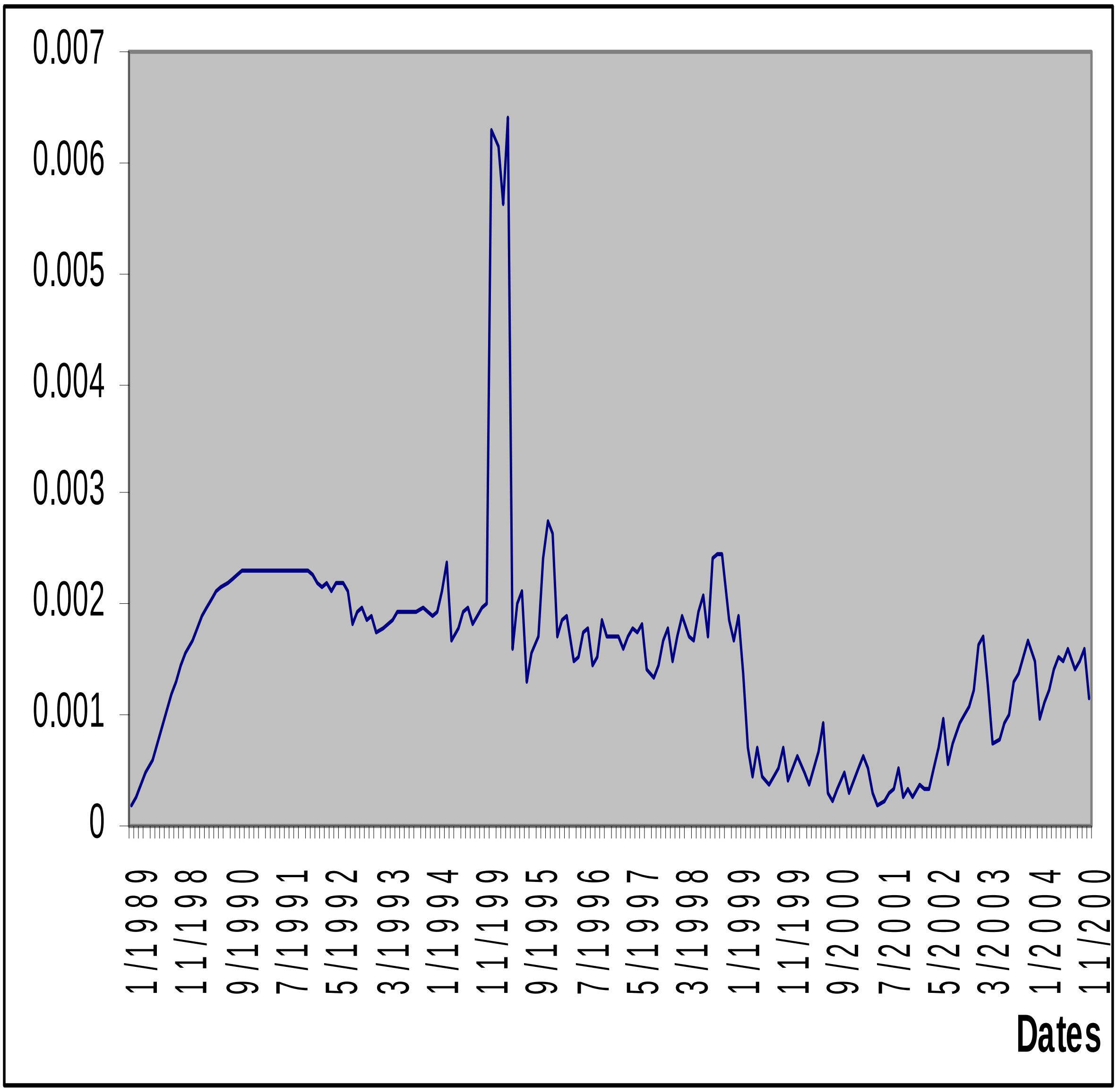


Table 1. A Synoptic View of Selected Empirical Studies on Mexican Agricultural Trade

Study	Commodities	Data/ Country	Volatility	Method	Results	Remarks
Diaz-Graces (2002)	Agg. Trade flows of consumption & capital goods	Yearly, 1980-89 and 1990-2000; Mexico	None	ECMs	Stable Export & import demand functions for 1990-2000.	-Did not test for unit roots; -equations estimated separately.
Espinoza-Arellano et al., (1998)	Winter Melon	Seasonal data from 1970 to 1994; Mexico-U.S.	None	3-SLS	Devaluation increased Melon exports to the U.S.	-Did not consider nonstationarity in data.
Konno and Fukushige (2002)	Aggregate	Quarterly data, 1981 to 1994; Mexico-U.S.	Long-term measure	Dynamic OLS	NAFTA has no significant impact	- Did consider nonstationarity in data but not cointegration.
Malaga et al., (2001)	Fresh Vegetables	Winter season data: 1974-93; Mexico-U.S.	None	3-SLS	Devaluation was effective only in the short-run.	- Did not consider either nonstationarity in data or cointegration.
Mora-Flores et al., (2002)	Crops and livestock	Yearly, 1988-96; Mexico	None	2-SLS	Real prices influence exports.	-Did not consider nonstationarity in data.

Table 2. Data description and Sources

Raw data	Description	Sources
Exchange Rate (ER)	Nominal bilateral monthly Mexico-US exchange rate, end of period exchange rate. 1989;1-2004;12.	USDA-ERS. http://www.ers.usda.gov/Data/ExchangeRate.
US CPI	Monthly general US CPI. All items. 1989;1-2004;12. Base Period: 1982-84=100.	US department of labor http://www.bls.gov/cpi/cpifaq.htm#Question%2015.
US Income ((USY)	US Personal Disposable Income1989;1-2004;12. ($US/person).	U.S. Department of Commerce http://www.bea.doc.gov/bea/an/nipaguid.
US Wage Rate (W)	US Farm Labor, Monthly, Dollars/hour. 1989;1-2004;12	US National Agricultural Statistics Service. (NASS). http://www.usda.gov/nass/pubs/reportname.htm#Farm_Labor.
Tomato Volume(Qt)	US Monthly exports (MT), HS 4-DIG. 1989-2004.	USDA-ERS. FAS, U.S. Trade Internet System
Imported Tomato Price (Pt)	Pt is the quotient of the total value of Mexican exports to US and the volume of these exports ($US/Ton). 89-2004.	USDA-ERS. FAS, U.S. Trade Internet System
Unites States Tomato Price (Pus).	Pus is the wholesale price in the Unites States. ($US/Ton). 1989-2004.	USDA-ERS. FAS, U.S. Trade Internet System

Table 3. Optimal Lag Length Selection for Model for Tomatoes using VG

Model VG Lags	NLAGS	Log-Likelihood	Number Variables	N*	ML-Test	LRm ${}^{2}_{d}$ (df=7) 5%	Test Results
Two to Three	2	-37.585		155	47.917	14.02	Reject Lag 2
	3	-37.991	37				
Three to Four	3	-37.991		147	45.669	14.02	Reject Lag 3
	4	-38.439	45				
Four to Five	4	-38.439		139	53.860	14.02	Reject Lag 4
	5	-39.065	53				
Five to Six	5	-39.065		131	63.517	14.02	Reject Lag 5
	6	-39.972	61				
Six to Seven	6	-39.972		123	24.330	14.02	Reject Lag 6
	7	-40.423	69				
Seven to Eight	7	-40.423		115	18.375	14.02	Reject Lag 7
	8	-40.907	77				
Eight to Nine	8	-40.907		107	20.403	14.02	Reject Lag 8
	9	-41.834	85				
Nine to Ten	9	-41.834		99	3.599	14.02	Accept Lag 9
	10	-42.434	93				
Ten to Eleven	10	-42.434		91	-332.696	14.02	**
	11	-9.164	101				
Eleven to Twelve	11	-9.164		83	-34.096	14.02	
	12	-10.476	109				

N* Net Number of Observations
**** Denotes that once an optimal lag is achieved, no further testing is conducted.**

Table 4. Unit Root Test Results for US and Mexican Macroeconomic Variables

Variable in Level Form	Estimated Coefficient	Lag Length (Months)	ADF Statistic*
Real US-Mexico ER	- 0.014912	5	- 1. 092
Trade Weighted Exchange Rate**	-0.018528	5	-0.9442
US Personal Disposable Income	- 0.026235	4	- 1. 076
US Rural Wage Rate	- 0.192870	12	- 2. 772
MX Real Per-capita GDP	- 0.326130	12	- 2.710
MX Rural Wage Rate	- 0.184490	3	- 1.315
Volatility of the Exchange Rate (VG)	- 0.502867	5	- 2.028
Variables in First Differenced Form			
Real US-México ER (Dls/Peso)	- 0.97241	5	- 7. 162
Trade Weighted Exchange Rate	-1.0096	5	-8.253
US Personal Disposable Income	- 1.59250	4	- 9. 156
US Rural Wage Rate	- 4.48070	12	- 8. 573
MX Real Per-capita GDP	- 4.53180	12	- 5.860
MX Rural Wage Rate	- 1.01620	3	- 29.14
Volatility of the Exchange Rate (VG)	- 1.23845	5	- 8.089

Table 5. Maximum Likelihood Cointegration Results: United States Import Demand for Tomatoes

A Seven-Variable Nine Lag System using VG

Eigenvalues	0.2400	0.1805	0.1340	0.1060	0.1013	0.0574	0.0210
Eigenvectors							
USY	0.4115	-0.2463	-0.6911	1.8000	1.2848	-4.0166	-0.4092
ER	1.0963	-0.5888	4.4093	0.3389	0.5607	1.6184	-0.6003
VG	-0.4492	0.4604	-3.1409	1.2229	1.0981	3.1982	0.2617
W1	0.4909	-0.5053	0.3450	-0.8676	-0.1159	-0.9232	2.6691
Pt	-2.4009	0.2850	2.2059	3.1834	-0.1577	-0.0887	0.5835
Pca	-1.6991	-0.8440	1.0792	-3.8290	1.1606	-0.3269	-0.0205
Qt	1.0000	1.0000	1.0000	1.0000	1.0000	1.0000	1.0000
Weights							
USY	0.00048	-0.00058	-0.00066	0.00031	-0.00001	-0.00070	-0.00021
ER	-0.00502	-0.00501	0.00418	-0.00147	0.00428	-0.00355	0.00126
VG	-0.00044	-0.00920	0.00009	0.00080	0.00243	0.00057	-0.00063
W1	-0.00995	-0.00791	0.00829	0.00263	0.00072	-0.00416	0.00000
Pt	0.04810	-0.01693	0.02731	-0.03221	-0.00068	0.00373	-0.00767
Pca	0.10529	-0.02008	0.03870	0.02992	0.01274	0.00086	0.01151
Qt	0.02564	0.04707	0.04380	-0.00188	0.01092	-0.00894	0.00101

Testing the Number of Cointegrating Vectors

Null hypothesis	**Trace Statistic**	**Trace (0.95)**	**MAX (λ)**	**MAX(λ) (0.95)**		
r = 0	166.80*	150.40	49.95*	50.51		
r ≤ 1	116.85	117.49	36.22	44.37		
r ≤ 2	80.63	88.59	26.19	38.22		
r ≤ 3	54.44	63.66	20.40	31.99		
r ≤ 4	34.04	42.70	19.43	25.68		
r ≤ 5	14.61	25.64	10.75	19.21		
r ≤ 6	3.85	12.34	3.85	12.34		

* Significance at 5 percent level (MacKinnon, 1999

Table 6. Error Correction Model results for the US Tomato Import Demand

Variable	Estimated Coefficient	T-ratio		Statistic	Critical Value
QT_{t-12}	0.198	3.274	R-Square	0.364	
USY_{t-10}	0.110	3.750	R^2 Adjusted	0.332	
ER_{t-1}	0.663	3.269	F-Value	9.198	
VG_{t-2}	-0.043	-2.671	DW-h Statistic	1.677	
PT_{t-10}	-0.099	-3.360	Skewness	0.041	0.000
PCA_{t-11}	0.195	1.508	Kurtosis	2.221	3.000
$W1_{t-6}$	0.130	2.936	J-B Normal	8.209	9.210 (X^2 (0.95))
ECT_{t-1}	-0.030	-3.453	Instability Test		
NAFTA	0.427	2.081	Variance	0.704	0.748
M3	0.011	0.809	Joint	2.777	3.150
M11	0.070	0.809	LM Statistic	26.816	35.172 (X^2 (0.95))
CONSTANT	0.309	2.365			

* Denotes the critical value of chi-squared with 23 degrees of freedom.
** Denotes the critical value of chi-squared with two degrees of freedom.

Table 7. Long and Short-run Results of the Hypothesis Testing for some key Variables

	Long-run*		**Short-Run**	
Null Hypothesis	LR-Statistic	Critical Value X^2 (0.95)	t/F Statistics	Critical Value (0..95)
Individual Coefficients				
$\beta_{er} = 0$	18.60*	3.84	3.27*	1.96
$\beta_{v} = 0$	13.44*	3.84	2.67*	1.96
Joint Coefficients				
$\beta_{er} = \beta_{v} = 0$	22.03*	5.99	6.97*	3.00
Equalityof Coefficients				
$\beta_{3j} = \beta_{5j}$	8.74*	3.84	6.22*	3.84

* indicates significance at 5 Percent level of error probability with *r*(*p*-*s*) degrees of freedom for the long-run,. and 180 degrees of freedom for the short-run.