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Pricing to Market, (Seasonal) Cointegration and US Agricultural Exports

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Abstract:

In this paper, we examine whether US exporters of agricultural commodities price to market. Specifically, we estimate the fixed-effects model of Knetter (1989; 1995), and alternative specifications based on the use of cost indices, and seasonal and vector error correction models that account for the time-series properties of the data.

Keywords: pricing to market, cointegration, US agricultural exports

JEL classification: F12, L13

1. Introduction

International agricultural trade is often dominated either by state trading agencies with monopoly power or by a few large firms (Patterson, Reca and Abbott 1996). If exporters have some market power and markets are segmented, price discriminating behavior across destination markets may be triggered by exchange rate movements, a phenomenon labeled as ‘pricing to market’ (henceforth PTM) by Krugman (1986). For example, under PTM, a US firm(s) exporting to Japan might find it optimal to lower its markup in the face of a depreciation of the Japanese yen against the US dollar. As a result, Japanese yen prices do not rise one-for-one with the nominal Japanese exchange rate, i.e., the degree of pass-through to yen prices is less than one, implying incomplete exchange rate pass-through. PTM is also a crucial reason for why there may be deviations from the law of one price or purchasing power parity (Goldberg and Knetter, 1997).

A priori, PTM may be expected to exist in agricultural exports, especially for countries with a large volume of agricultural trade, such as the US. The previous empirical literature on PTM has shown different price discrimination patterns of US agricultural exporters across export commodities and destination countries. Both evidence for and against PTM have been found in previous studies, e.g., see Knetter (1989), Pick and Park (1991), Patterson, Reca and Abbott (1996), and Park and Pick (1996). The earliest empirical study by Knetter (1989), whose sample of six 7-digit US export products included dried onions, bourbon, orange juice, and breakfast cereal, showed that US export prices were rather insensitive to exchange rate fluctuations, and that when price adjustment did occur it frequently amplified the effect of exchange rate changes on local currency prices. Pick

and Park (1991) applied Knetter's (1989) model to US cotton, wheat, corn, soybean, and soybean meal exports, and found little evidence of PTM, except in the wheat export sector. Patterson, Reza and Abbott (1996) found discriminatory pricing in US chicken exports to Canada, Mexico, Netherlands Antilles, Columbia, the Netherlands, Hong Kong and Japan. In addition, negative and significant exchange rate coefficients were found for Canada and the Netherlands, providing support for PTM in those markets¹. Following Knetter (1995), Park and Pick (1996) imposed a symmetry restriction in their analysis of US wheat exports, whereby the effects on export prices of marginal cost and exchange rate changes are constrained to be identical. Their results indicated that for six of eight countries in their sample, Egypt, Japan, Korea, Philippines, Taiwan and Venezuela, the exchange rate coefficients were negative and significant.²

A weakness of the Knetter (1989; 1995) PTM model is a failure to fully account for the nature of competition among firms. Adolfson (1999) investigated the transmission of exchange rate changes into Swedish export prices and the importance of market share for price determination. Based on the consideration that marginal costs were likely to change due to movements in the exchange rate because of implied changes in the prices of imported inputs, he allowed for feedback effects between the variables, export price, marginal cost, market share, exchange rates and the price of substitutes. Formal tests using an error correction model indicated consistency with price discrimination. Swift (2004) examined

¹ Brown (2001) analyzed Canadian canola exports, price discrimination being identified for three export destinations, Japan, US and Mexico, and PTM in the case of Japan.

² Carew (2000) concluded that Canadian exporters of wheat, pulse crops, and tobacco were able to price discriminate to Italy, the UK, Japan and Bangladesh. His results for US wheat exports strongly suggested imperfectly competitive and price stabilization behavior by exporters, evidence of price discrimination being found for exports to South Korea, Egypt, Venezuela, and the Philippines.

the effects of exchange rate changes on the price of Australian exports of milk products, cheese, beef, sheep-meat, and hides and skins. A dynamic long-run relationship in all potentially endogenously determined variables, export prices, world price, production cost, and exchange rate, was assumed in the model. The results indicated that dairy exports operate in competitive markets in which pass-through is complete, but no long-run relationship existed between exchange rates and prices for livestock products.

The objective of this paper is to re-examine the evidence for PTM in US agricultural exports using relatively high frequency data and the appropriate cointegration methods. In section 2, we lay out two different empirical models of PTM, followed in section 3 by a discussion of the data and model selection. The estimation results are presented in section 4, and some conclusions are drawn in the final section.

2. A Model of Pricing to Market

Knetter (1989; 1995) and Goldberg and Knetter (1997) proposed the basic model used in many studies of PTM. An exporter is assumed to sell to N foreign destinations, indexed by i . The profit of the firm is given by:

$$\Pi(p_1, \dots, p_n) = \sum_{i=1}^n p_i q_i(e_i p_i; \nu_i) - C\left(\sum_{i=1}^n q_i(e_i p_i; \nu_i), w\right), \quad (1)$$

where p is the export price in the exporter's currency, e is the exchange rate (units of destination market currency per unit of the exporter's currency), q is quantity demanded (a function of the price in the buyer's currency, ep , and a demand shifter ν), w is an index of input prices in units of the exporter's currency, and C is the total cost function. The first order conditions for profit maximization imply that the firm equates the marginal revenue

from sales in each market to the common marginal cost. Alternatively, the export price to each destination is the product of the common marginal cost and a destination-specific markup:

$$p_i = MC \left(\frac{-\eta_i}{-\eta_i + 1} \right) \quad i = 1 \dots n, \quad (2)$$

where the arguments of marginal cost, MC , are suppressed, and η is the absolute value of the elasticity of demand with respect to local currency price in destination market i . A change in the exchange rate *vis-à-vis* the currency of country i can affect the price charged to market i in two ways: by affecting either marginal cost (through changes in quantity or input prices) or the elasticity of import demand. The commodity is assumed to be identical across destination markets, which implies that the marginal cost is independent of the destination. Therefore, the former effect will be common to all destination markets, while the latter, fluctuations in markup, are believed to be destination-specific; they do not reflect the behavior of prices to other export destinations. Both effects determine exchange rate pass-through, while pricing-to-market refers to the second effect only. So an empirical analysis of goods prices and exchange rates must be capable of discriminating between fluctuations in the markup over marginal cost from changes in marginal cost.

Knetter (1989) estimated the following fixed-effects model of export prices across destinations for a particular industry:

$$\ln p_{it} = \theta_t + \lambda_i - \beta_i \ln e_{it} + \varepsilon_{it}, \quad (3)$$

where p is price in units of the exporter's currency measured at the port of export, θ_t is a set of time effects, λ_i is a set of destination country effects, e is the exchange rate in units of the

destination market currency per unit of the exporter's currency, ε is a random disturbance term, i indexes destination, and t time period.

Three forms of market behavior can be inferred from the estimated parameters of (3). First, there is the null hypothesis of perfect competition where prices equal marginal cost and export prices are the same across destinations, i.e., λ_i and all β_i are zero. In addition, changes in marginal cost over time are measured by the time dummy variable θ_t . Second, there may be price discrimination across markets, but the price elasticity of the residual demand curve is assumed constant for each export market, i.e., the λ_i can differ across markets, but all β_i are zero. Third, there is price discrimination, and the price elasticity of the residual demand curve is assumed to be non-constant, generating non-zero values for β_i , i.e., the change in the importer's local price causes a change in the price elasticity of demand, resulting in a change in the mark-up. The expected sign of β_i , negative (positive), will depend on whether demand is less convex (more convex) than a constant elasticity demand function. A negative coefficient is consistent with Krugman's (1987) original idea of PTM, while a positive coefficient implies that the exporter is amplifying the exchange rate effect.

Due to the popularity of Knetter's (1989; 1995) model, in our subsequent empirical analysis, we treat estimation of (3) as a benchmark. However, there are some key weaknesses to using (3) that need to be recognized. First, as noted by Knetter (1989) and emphasized in Goldberg and Knetter (1997), export unit values are not sufficiently detailed to ensure that product qualities are identical across all export destinations. Consequently, using only one time-specific effect as a measurement of marginal costs for various export

destinations is questionable, impairing the explanatory power of λ . Second, while the use of a time-specific constant to capture changes in marginal cost avoids the potential biases caused by using cost indices as a proxy for marginal cost, excluding such cost indices may also generate biased results. For example, if costs are non-stationary, then different results can be obtained for the significance of β_i when using a cost index compared to using the time-specific effect which is deterministic.

Therefore, taking logarithms of (2) and using the first-order Taylor series approximation of the function $\ln(\eta_i/\eta_i-1)$, and collecting terms, yields the following estimable relationship:

$$\ln p_{it} = \mu_i + \gamma_i \ln MC_t + \beta_i \ln e_{it} + \varepsilon_{it}, \quad (4)$$

where μ_i is a destination-specific intercept that captures the constant terms in the Taylor series, MC_t is a cost index, designed to capture common changes in cost over time, and all other terms have the same definition as in (3). In terms of interpreting the parameter γ , if the cost index MC_t is an exact measure of marginal cost and the estimated value of γ is significantly different from 1, there is evidence for imperfect competition. However, if MC_t is a cost index then the levels, and hence γ are arbitrary and no inferences can be drawn about the extent of competition.

In estimating (4), three cases can be identified where β_i will not be significantly different from zero. First, with a single competitive market for exports, MC_t and γ measure the effect on export prices of common cost movements over time, and μ_i captures the effect on export prices of any product differentiation by destination, assumed constant over time. Second, if the market is integrated, but firms act imperfectly competitively, MC_t and γ measure the effect on export prices of common cost movements over time and a

common mark-up, while μ_i captures the effect of any product differentiation by destination and any country-specific asymmetric effects of cost movements and common mark-ups on export prices, again constant over time. Third, if the price elasticity of the residual demand curve is constant, MC_t and γ measure the effect on export prices of common cost movements over time, while μ_i captures the effects of both product differentiation by destination and any country-specific mark-ups, both effects being constant over time.

A final case can be noted where the coefficient β_i is significantly different from zero, implying behavior consistent with pricing to market across destination markets. As with the model presented in (3), the expected sign of β_i , negative (positive), will depend on whether demand is less convex (more convex) than a constant elasticity demand function. In addition, MC_t and γ measure the effect on export prices of common cost movements over time, and μ_i captures the effect of product differentiation by destination, assumed constant over time.

3. Data Tests and Model Selection

Recent studies have pointed out that contemporaneous effects are not enough to capture the actual relationship among the three key variables in (4) and that future or delayed impacts should be accounted for. In addition, given that export prices and exchange rates are typically non-stationary time series, then assuming stationarity in these data series is likely to generate improper results for the relationship between them (Adolfson 1999; Swift 2004). The method used by the latter researchers, vector autoregressive error-correction (VECM) for first differenced series with constant seasonal dummy parameters, requires that the data

should be integrated of degree 1, $I(1)$, which assumes the roots of interest are precisely one, and there are no other unit roots in the system, thus differencing them will result in stationary data. If a variable such as the export price, is $I(2)$, its inclusion is likely to generate an explosive data process instead of converging to a stationary long-run equilibrium.

In addition, if non-stationary stochastic seasonality, which can be characterized by seasonal unit roots corresponding to peaks at seasonal frequencies in the spectrum, is an important source of variation in the system, simple seasonal adjustment might lose a significant part of valuable information about seasonal fluctuations, and result in mistaken inference about economic relationships between time series data. Problems of this kind do suggest extending the concept of cointegration to consider the possibility that common roots exist at seasonal frequencies as well as the zero frequency, which results in the idea of seasonal cointegration (SECM) as discussed by Hylleberg, Engle, Granger, and Yoo (1990) (HEGY), Lee (1992), Engle, Granger, Hylleberg and Lee (1993) (EGHL), Johansen and Schaumburg (1999).

In terms of data, quarterly data for US export prices and exchange rates from 1989:1 to 2003:4 are used in the estimation. Export prices are measured by the unit export values for six U.S. agricultural products, beef, pork, wheat, corn, soybean and lettuce.³ These data are taken from USDA's *Foreign Agricultural Trade of the United States* (FATUS), which guarantees that the product divisions of exports are consistent over time. Quarterly nominal

³ Beef is defined as fresh/frozen beef and veal, pork is fresh/frozen pork, wheat is un-milled wheat, corn is corn, soybeans are soybeans, although according to FATUS, soybeans ex seed (soybeans, whether or not broken, except seeds for sowing) is the subcategory of soybean, but these two have the exact same numbers over the period of interest, and lettuce is fresh lettuce.

and real exchange rates, measured as the local currency per US dollar, are constructed by taking the quarterly average of nominal monthly average exchange rates and real monthly country exchange rates, the latter being the nominal rates deflated by the consumer price index in the relevant destination markets. The exchange rate data are derived from the *International Financial Statistics* of the International Monetary Fund and the *Financial Statistics* of the Federal Reserve Board, as summarized by USDA. Producer price indices are taken from the Bureau of Labor Statistics work, and are used as a proxy for marginal cost.

Given these three data series, stationarity and trend stationarity testing of the variables is performed using Augmented Dickey-Fuller (ADF), Phillips-Perron (PP), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests. Even though these three test statistics do not show perfectly consistent results for all data series, strong non-stationarity is observed in the exchange rate series, both nominal and real. On the other hand, the export unit values series for beef and lettuce exhibit strong stationarity, while the other export unit values series and the producer price series show mixed evidence. We are unable to reject the null hypotheses of non-stationarity based on the ADF and PP test, and we also fail to reject the null hypothesis of stationarity using the KPSS test. As noted earlier, an assumption underlying the VECM model is that all variables are integrated of the same order, i.e., integrated of order 1. To rule out the possibility that our data are I(2), we perform a stationarity test of the first difference of all the data series. Use of the ADF test rules out the possibility of I(2) in any of our data series.

Due to the fact that there is typically strong seasonality in the prices of agricultural products, it is also natural to check for the presence of unit roots at some seasonal

frequencies as well as at the zero frequency, for which the so-called HEGY test is applied. The inclusion of lags was determined by first estimating the equation with three years of lags and then excluding those lags that failed to enter significantly at the 15 percent level. This approach trades off the loss of power, which results from including unnecessary lags, against the bias that results from excluding necessary lags (Beaulieu and Miron 1993). We have also combined the Akaike Information Criterion (AIC), the Corrected Akaike Information Criterion (AICC), and the Schwarz Bayesian Criterion (SBC) to determine the number of lags. Results of seasonal unit roots test are very sensitive to the selection of lagged terms. More lagged terms are included when using the AIC and AICC criteria and seasonal unit roots are more likely to exist in the data. Generally, the seasonal error correction model (SECM) fits our data requirements, in which cointegrating relationships, not only at zero frequency but at seasonal frequencies, are what we need to derive.

By using the SECM method, we can distinguish between the long-run equilibrium relationship between export unit values, exchange rates and production costs, that is, the manner in which the three variables drift upward together, and the short-run dynamics, that is, the relationship between deviations of export unit values from their long-run trend and deviations of exchange rates and production costs from their long-run trends, not only at the zero frequency but also at seasonal frequencies. While differencing the data obscures the long-run relationship, cointegration and error correction preserves information about both forms of covariation.

We estimate the following SECM version (Lee, *op. cit.*) of (4):

$$\Delta_4 X_t = \Pi_1 Y_{1,t-1} + \Pi_2 Y_{2,t-1} + \Pi_3 Y_{3,t-1} + \Pi_4 Y_{3,t-1} + \sum_{j=1}^p \Phi_j^* \Delta_4 X_{t-j} + \varepsilon_t, \quad (5)$$

where:

$$\begin{aligned}
X_t &= (\ln p_{it} \ln e_{it} \ln pp_i)' \\
Y_{1t} &= S_1(B)X_t = (1 + B + B^2 + B^3)X_t \\
Y_{2t} &= S_2(B)X_t = (1 - B + B^2 - B^3)X_t \\
Y_{3t} &= S_3(B)X_t = (B - B^3)X_t = B(1 - B^2)X_t
\end{aligned}$$

where $S_l(B)$ is a seasonal filter that eliminates unit roots at all seasonal frequencies ($\omega = 1/2$ and $1/4$), while $S_2(B)$ removes unit roots at frequency $\omega = 0$ and seasonal frequency $\omega = 1/4$; $S_3(B)$ eliminating frequencies $\omega = 0$ and $\omega = 1/2$. Therefore Y_{1t} , Y_{2t} and Y_{3t} have unit roots only at frequency $\omega = 0$, $\omega = 1/2$ and $\omega = 1/4$, respectively. Because the coefficient matrices Π_1, \dots, Π_4 reveal information about long-run economic relationships of the series at each corresponding frequency, we need to derive properties of these matrices in order to determine whether or not the components of X_t are seasonally cointegrated in the presence of unit roots at other frequencies. If the matrix Π_k ($k = 1, \dots, 4$) has a full-rank (in our case $r = 3$), all the components of Y_t theoretically do not contain unit roots. On the other hand, a cointegration relationship does not exist if the rank of matrix $\Pi_k = 0$ even though there may exist (seasonal) unit roots in the data series. In the intermediate case, where the rank of the matrix Π_k is $0 < r < 3$, there are r stationary cointegrating relationships at the corresponding (seasonal) frequency, when $r = 1$, $\alpha_k = (\alpha_{1k} \alpha_{2k} \alpha_{3k})'$ and $\beta_k = (\beta_{1k} \beta_{2k} \beta_{3k})'$; when $r = 2$, $\alpha_k = \begin{pmatrix} \alpha_{11k} & \alpha_{21k} & \alpha_{31k} \\ \alpha_{12k} & \alpha_{22k} & \alpha_{32k} \end{pmatrix}'$ and $\beta_k = \begin{pmatrix} \beta_{11k} & \beta_{21k} & \beta_{31k} \\ \beta_{12k} & \beta_{22k} & \beta_{32k} \end{pmatrix}'$, which satisfies $\Pi_k = \alpha_k \beta_k'$ for suitable matrices α_k and β_k such that $\beta_k' Y_{k,t-1}$ is stationary, defined as the underlying or long-run equilibrium economic relationship even though $Y_{k,t-1}$ itself is non-stationary, and α_k are the adjustment coefficients that result in agents reacting to the disequilibrium error, returning variables to their equilibrium path.

The annual restriction of contemporaneous cointegration at complex frequencies is also imposed, i.e., $\Pi_4 = 0$. Lee has argued that the absence of non-synchronous seasonal cycles should have little effect on the test for seasonal cointegration at frequency $\frac{1}{4}$, although this restriction is criticized by Johansen and Schaumburg (*op. cit.*), who state that the test for cointegration rank at complex frequencies is only partially correct. Lof and Lyhagen (2002) show that the version of the seasonal cointegration model with the annual restriction and unrestricted seasonal dummies included is better than the SECM proposed by Johansen and Schaumburg with restricted seasonal intercepts if one step ahead forecasts are considered. So Lee's model has a strong advantage in terms of a seasonal cointegration test and estimation.

As mentioned above, a full-rank matrix Π_k implies no (seasonal) unit root at the corresponding frequency. If the cointegration rank test shows no seasonal unit roots at all seasonal frequencies, then the SECM reduces to a simple VECM. Therefore, if the test results confirm no existence of a cointegration relationship at frequency zero and no seasonal unit roots, we try a VECM version including a constant in the error correction term, implying that the components are stationary around constants. Due to the fact that restriction on the vector of intercepts will lead to different null distributions of the likelihood ratio tests of the number of cointegrating vectors, this modification may generate more evidence of cointegrating relationships among the relevant variables.

We estimate the following VECM version of (4):

$$\Delta \begin{pmatrix} \ln p_{it} \\ \ln e_{it} \\ \ln mc_t \end{pmatrix} = \Pi (\ln p_{it}, \ln e_{it}, \ln mc_t, 1)' + \sum \Phi_j^* \Delta \begin{pmatrix} \ln p_{it-j} \\ \ln e_{it-j} \\ \ln mc_{t-j} \end{pmatrix} + \varepsilon_{it} \quad (6)$$

where $\Pi = \alpha\beta'$ and $\beta' = (\beta_1, \beta_2, \beta_3, \beta_0)$. $\beta' (\ln p_{it}, \ln e_{it}, \ln mc_t, 1)'$ is again defined as the underlying or long-run equilibrium economic relationship, and α are the adjustment coefficients which have the same explanation as in the SECM model. If the matrix Π has a full-rank ($r=3$), all the components of Y_t theoretically do not contain unit roots. On the other hand, a cointegration relationship does not exist if the rank of matrix $\Pi = 0$. In the intermediate case, where the rank of the matrix Π is $r < 3$, there are r stationary cointegrating relations, which mean the linearly independent vector $Z_{it} = (\beta_1, \beta_2, \beta_3, \beta_0)(\ln p_{it}, \ln e_{it}, \ln mc_t, 1)'$ is stationary. A constant is entered via the error correction term also based on the earlier discussion that an intercept term should be included in the relationship among export unit values, producer prices and exchange rates in order to capture country-specific effects. Therefore in terms of the VECM model, our first hypothesis test is a constant restriction on the error term, i.e., the constant exists or not. Based on the theoretical concerns, this restriction is preferred, but it is not necessary based on the characterization of the data series since use of producer cost indices makes the level of correlation between the variables arbitrary.

4. Empirical Results

(i) Basic Evidence

We first present some basic empirical evidence of PTM by comparing the movements of US export prices to Japan, Canada, Mexico, and South Korea respectively with US export prices to all other trading partners from 1995 to 1998 (Figure 1). Export unit values of the US to Japan (Canada, Mexico, and South Korea) are used as export prices for both 1995 and 1998.

Export prices to all other trading partners are constructed using the difference between the US total export value and export value to Japan (Canada, Mexico, and South Korea), divided by the difference between total US export quantity and the export quantity to Japan (Canada, Mexico, and South Korea). So each export market, Japan, Canada, Mexico and South Korea has their own export price and correspondingly there are different export prices for all the other trading partners.

The reason for choosing this time period is that according to the nominal exchange rate data, the US dollar kept appreciating with respect to these four countries from 1995 to 1998. Since the US dollar exchange rate overall showed the same tendency over this period, the starting and ending export prices capture the overall changing patterns over this period. If the US did not engage in any PTM, the movement of export prices to Japan and export prices to all other trading partners would be identical (the same reasoning for the other three countries). We assume that opposite patterns of change (positive or negative) for movements in export prices to a particular country and those to all other countries are too significant to be explained by changes in other economic variables, and are presented as basic evidence for PTM by US exporters, although PTM may also exist when export prices exhibit the same pattern, export prices changed to differing degrees.

From figure 1, one can see the extreme cases where opposite movements in export prices occurred. The data indicate that US exporters apply PTM for exports of pork and lettuce to Japan, exports of pork to Mexico, and exports of lettuce to South Korea, indicating concentration of PTM in two product categories, pork and lettuce. Other cases deserving of attention are those where there is a relatively large gap between movements in export prices.

These occur in beef, pork and lettuce for Canada, beef for Mexico, and pork for South Korea, where the smallest divergence between export prices exceeds 225 percent. Therefore, roughly speaking, there seems to be evidence of PTM for all four export markets. However, PTM is not universal across product categories and destination markets. The evidence seems to show that PTM is limited to pork and lettuce, and perhaps beef and wheat. The sensitivity of the occurrence of PTM is what we would like the empirical estimation to confirm.

An interesting product category is soybeans which contrasts with the export price patterns of the other five products. Given the strong US dollar from 1995 to 1998, US exporters chose to reduce export prices in their own currency in order to absorb part of the appreciation for the other five products, while export prices of soybeans increased, albeit by a relatively small amount, which may imply that US exporters amplified the effects of exchange rate changes in their export prices.

(ii) *Empirical Results*

We present results based on estimation of the fixed-effects model given in (3) and also estimation of the SECM and VECM models in (5) and (6) respectively. Since all exports to the various destination markets are from the same exporting country, the US, it is natural to assume the existence of contemporaneous correlation of errors across equations (importing countries), consequently, seemingly unrelated regression (SUR) is applied to the fixed-effects model.

The cointegration rank is tested for by applying Lee's trace test, and the number of cointegrating vectors at zero frequency are calculated as well as at seasonal frequencies. The annual restriction simplifies trace values as partial canonical correlations at each

frequency. Table 1 shows the results of the trace test statistics for seasonal cointegration in the case of US wheat exports. When nominal exchange rates are included in the estimated system, existence of cointegration at zero frequency is confirmed for wheat exports to Egypt, Japan and Mexico. Cointegration at frequency $1/4$ has been found for wheat exports to China, Philippine, Jamaica and Thailand, in which two long-run cointegrating vectors exist in all four cases with the first cointegrating vector as the main equilibrium relationship in the system. No cointegration relationship exists at frequency $1/2$.

Cointegration ranks are less likely to be verified when real exchange rates are used. Cointegration relationships at zero frequency exist in US wheat exports to China, Egypt and Japan, but only with ten percent significant level for the first two countries. Two cointegrating vectors at frequency $1/4$ are substantiated in wheat exports to Jamaica and Thailand. Again cointegration at frequency $1/2$ is significantly rejected in all cases. A lack of cointegrating relationships could be caused by either no unit roots at corresponding frequencies or no long-run relationships even though (seasonal) unit roots exist. The former causes no cointegrating rank at zero frequency for wheat exports to South Korea and Philippines; the latter leads to no cointegration at zero frequency for Mexico in the case of real exchange rates.

The coefficients of the cointegrating vectors for US wheat exports, based on the first cointegrating vectors, are reported in table 2a. For exports to South Korea, since the trace test indicates no cointegration at zero frequency and no seasonal unit roots in the three time series, irrespective of whether nominal or real exchange rates are used, restriction of a constant in the error term in VECM is applied and one underlying relationship among

export unit values, the exchange rate and the producer price index is derived, however, when we use the same method for wheat exports to the Philippines with the real exchange rate, no long-run equilibrium relationship can be found. For the case of wheat exports to Mexico and the real exchange rate, (seasonal) unit roots are not supported by the data.

When nominal exchange rates are used, for exports to Egypt and the Philippines, positive coefficients on the exchange rate show that US exporters tend to amplify the effect of nominal exchange rate fluctuations, so that when the US dollar appreciates, exporters increase the local price in their destination markets. On the other hand, for exports to China, Japan, South Korea, Mexico, Jamaica and Thailand, negative coefficients for nominal exchange rates reveal that the effect of nominal exchange rate changes is proportionately offset, which implies that if the US dollar appreciates, US exporters decrease US dollar denominated export prices in order to absorb part of the increase in prices in destination country currencies. Exports to China and Philippines show a particularly strong effect of the nominal exchange rate on export unit values, the size of the coefficient seeming to be much larger than for other trading partners, which suggests other factors may be important in affecting exporter behavior, such as import quotas, and a state trading agency. In all cases, the coefficients for the producer price index show consistency with the economic expectation that cost increases will result in an increase in prices, especially for China, Mexico, Jamaica and Thailand, a larger percentage rise in export prices follows an increase in producer prices.

Similar evidence for the effect of real exchange rates on US wheat export prices has been discovered as long as cointegration exists, except in the case of South Korea, in which

the magnitude of the effect of exchange rates is almost negligible. There is one counter-intuitive result in the case of the impact of producer prices on export prices for Egypt; however, the evidence for cointegration is not strongly significant.

In table 2b, the relevant coefficients from the fixed-effects model are reported for US wheat exports. Comparing the results between the SECM and fixed-effects models, a different story emerges for US wheat exporter pricing behavior. In most cases, the fixed-effects model supports a positive effect of exchange rates on unit export values, especially in the case of real exchange rates, although not all of the estimated coefficients are significant.

Due to the fact that the data series for US beef, pork, corn, soybean and lettuce exports do not exhibit cointegration at seasonal frequencies, we now report the results for the VECM model. As mentioned in section 3, if the relevant hypothesis test fails to reject the constant restriction on the error correction term in the VECM model, it is treated as the correct model, implying that the components are stationary around constants. If this restriction is rejected, the error correction model without restriction should be considered. However, in our empirical results, we found no big difference in coefficients between the restricted and unrestricted model in cases where the restriction was rejected, which suggests that in these cases including the constant in the error correction term does not substantively change the basic relationship among the variables.

Two test statistics proposed by Johansen (1988), the trace test and the maximum eigen-value test, are used to test the cointegration rank and get the number of cointegrating vectors. Greater weight is placed on the trace test results, since the trace test is more robust

in terms of testing for the possible non-normality of the residuals. In addition, the SAS 9.1 Program only performs the restriction test using the trace value. If the restriction is rejected, the unrestricted model is estimated and then a significance test on the coefficient of the exchange rate in the long-run relation is applied to check that this effect indeed exists.

Table 3a shows the coefficients of the cointegrating vectors for US beef exports. The results indicate there are no long-run relationships between the variables for exports to Japan, Canada, Mexico and Switzerland, irrespective of whether nominal or real exchange rates are used. For exports to Hong Kong, there exists an underlying relationship between export unit values and the real exchange rate, but not with the nominal exchange rate. For the Philippines, when the real exchange rate is included, one cointegrating vector is found in the restricted model, but for the nominal exchange rate, the restriction was rejected, so the unrestricted model is estimated, however, the significance test shows that the effect of exchange rates is not significantly different from zero.

In the cases where cointegrating vectors are found, only one rank was confirmed. For exports to the Philippines and Indonesia, US exporters tend to amplify the effect of exchange rate fluctuation, so that when the US dollar appreciates, exporters increase the dollar denominated export price. On the other hand, for exports to South Korea, Hong Kong, and Singapore, if the US dollar appreciates, the increase in foreign currency prices received by importers is less proportionate than the rise in the dollar value. In most cases, the coefficients for the producer price indices show the expected positive signs..

In table 3b, the results from using the fixed-effects model indicate a different story about US beef exporters' pricing behavior. For exports to Japan, Canada and Switzerland,

significant negative effects of nominal exchange rates on export unit values are found, but since no equilibrium relationships were found using the VECM model, it is hard to compare results. In other cases, the fixed-effects model indicates significant positive effects on export unit values for exchange rates.

Table 4a shows the results for US pork exports, with stronger cointegration evidence being found for real exchange rates: cointegrating evidence is found for five export destinations using real exchange rates, compared to three export destinations using nominal exchange rates. In addition, the restricted VECM model supports a cointegrating relation in the system because full rank happens occasionally in the unrestricted model, while reduced rank is often found in the restricted model. In the case of Japan, restriction of a constant in the error term was not rejected for both nominal and real exchange rates, however, no cointegrating vector was found in the restricted model, implying that there is no long-run economic relationship between the variables for US pork exports to Japan. Opposite patterns were found for Mexico in the case of nominal exchange rates, full rank for Mexico implying the series has no unit roots. No long-run relationship exists for Canada. For the rest of the cases, negative coefficients on the nominal exchange rate again imply US exporters absorb part of the increase in prices in destination country currencies. When real exchange rates are involved in the system, positive effects are found in two cases, and there are also some counter-intuitive results in the case of the impact of producer prices on export prices for Mexico and Singapore.

In table 4b, the results for the fixed-effects model are reported. Using nominal exchange rates, the results show that in the case of Hong Kong and South Korea, the opposite

results are obtained compared to the VECM model, while for Singapore, the same effect appears. With real exchange rates, the results also show inconsistency with the VECM model for Canada and South Korea. Overall, changes in real exchange rates tend to have a positive impact on export unit values.

In table 5a, real exchange rates give us more clear-cut results about the number of cointegrating vectors in the sense that restrictions on the error term fail to be rejected and only one cointegrating relationship is found in most cases except for the Dominican Republic in which there is no equilibrium relationship. All the effects for real exchange rates are negative meaning U.S. corn exporters partially absorb the effect of fluctuations in real exchange rates, although the magnitude of some of the coefficients is close to zero, e.g., -0.01 for Japan, and -0.02 for Egypt. In the case of nominal exchange rates, the restricted model fails to be rejected only at the 10 percent significant level for Mexico, for which a positive effect of the nominal exchange rate is found. Two cointegrating relationships are found for Algeria, although the results of the two models do not show significant difference from zero with a p-value equal to 0.11. The outlier is Jamaica, where a trend variable probably ought to be included in the estimation.

The fixed-effects model (table 5b) in the case of corn exports supports many of the results from the VECM model for the case of nominal exchange rates, except for the opposite effects for Japan and Algeria. In the case of real exchange rates, only the coefficient for Japan is consistent with that from the VECM model; for the other countries, there is a sharp contrast between the VECM and the fixed-effect models in terms of the effects of real

exchange rates on export prices. The results show that the fixed-effects model tends to generate significant positive effects on export unit values just as in the wheat case.

Clearly, as shown in table 6a, restriction on the error term is supported by real exchange rates for soybean exports, and in addition, the effects of real exchange rates on export unit values are all positive, suggesting that US exporters believe local consumers in importing countries have a special preference for US soybean and their demand will not decrease even when soybean prices in local currencies increase. Unrestricted models were estimated for Mexico and Costa Rica in the case of nominal exchange rates, but the impact of the nominal exchange rate turned out to have no significant effect on export unit values for Mexico. Generally, the results for nominal exchange rates suggest that country-specific concerns are important. For the Netherlands, South Korea and Canada, it may be supposed that local consumers in importing countries are more likely to choose US products, and more unwilling to switch to local or other countries' exports, while for Mexico, Costa Rica and Japan, local consumers are assumed to be relatively more willing to change their source of soybeans. Again, as can be seen in table 6b, the transparent pattern of results for the fixed-effects model is positive coefficients for both exchange rates, especially real exchange rates.

Finally the results for lettuce shown in table 7a are more complicated, with more than one cointegrating relationships being found in most cases for both nominal and real exchange rates. In terms of the first cointegrating relationship, for Hong Kong, Japan and Singapore, US exporters tend to absorb fluctuation in exchange rates, the notable exceptions being Canada and Mexico. Quite similar results are found for each importing country in table 6b. For Canada and Mexico, the effects of exchange rates on export unit values are positive,

while for Hong Kong, Japan and Singapore, the effects of exchange rates on export unit values are negative.

5. Conclusions

In this paper, we revisit the question of whether US exporters of agricultural commodities price to market. Specifically, we estimate the commonly used fixed-effects model of Knetter (1989; 1995), and an alternative specification based on the use of cost indices and seasonal and vector error correction models that account for the time-series properties of the data.

In the case of the SECM model for US wheat exports, nominal exchange rates are more likely to form cointegrating relationships when zero frequency as well as seasonal frequencies is considered, which gives support to the view that nominal exchange rates are what exporters really consider in their pricing decisions. This is consistent with intuition since exporters easily observe changes in nominal exchange rates, but it is both difficult and costly to get information on consumer prices in each importing country. A negative effect of exchange rates on export unit values is the main result in the equilibrium relationships at both zero frequency and seasonal frequencies.

In all cases under nominal exchange rates, long-run equilibrium cointegrating relationships are substantiated for US wheat export markets. For six importing countries, including China exporters lower the dollar price in order to stabilize prices in the importer's currency. These results support the notion of a local-currency stabilization pricing strategy consistent with Krugman's original idea of PTM, and also that wheat is a fundamental

consumption product with a less convex demand structure. Considering that China's nominal exchange rate is fixed by the Chinese government it is somewhat surprising to find existence of cointegration. However, over the time period covered in the data, four jumps in nominal exchange rates occurred, although during the four sub-periods there was no change. Consequently, we can think of the coefficient on the nominal exchange rate as the effect of exporters' expected changes in China's nominal exchange rate.

In the case of Egypt and the Philippines, wheat exporters try to amplify US dollar denominated export prices in the face of an appreciation of the US dollar. In order to explain the positive effects of exchange rates in these two countries, we could either include more variables in the system, for example market share of imported US wheat in local consumption, price of substitution etc, to capture the underlying economic reasons, or consider the possible distorting effects of trade policy.

Two main conclusions can be drawn from the VECM model: real exchange rates are more likely to form cointegrating relationships, and a constant restriction on the error term in the model is supported in more cases when real exchange rates are used. The equilibrium relationships are more clear-cut in the case of real exchange rates compared to nominal exchange rates since nominal exchange rates tend to generate more than one cointegrating vector and provide mixed results about the effects of exchange rates.

In most cases, long-run equilibrium cointegrating relationships are confirmed, which hints at the existence of price discrimination, PTM and market power in US export markets. The estimation results for five US agricultural exports show product-specific and country-specific characterization in terms of PTM. In the case of corn and soybean

exports there are consistent results across importing destinations, real exchange rates having a negative effect on corn export unit values, and a positive effect on soybean export unit values, suggesting soybean exporters try to amplify US dollar denominated export prices in the face of an appreciation of the US dollar, while corn exporters will lower dollar prices to stabilize prices in the importer's currency. When comparison is made not across products, but across destinations, we find that for Japan, Hong Kong, Singapore and South Korea, negative effects of exchange rate changes on export unit values are often found, suggesting that when the US dollar depreciates, these countries probably play a less important role in expanding export volumes and improving the current account of the US balance of payments. A positive impact has been found for the Philippines. In the case of Mexico and Canada, two of main trading partners of the US, the effects of exchange rate changes on export unit values are very mixed depending on product categories.

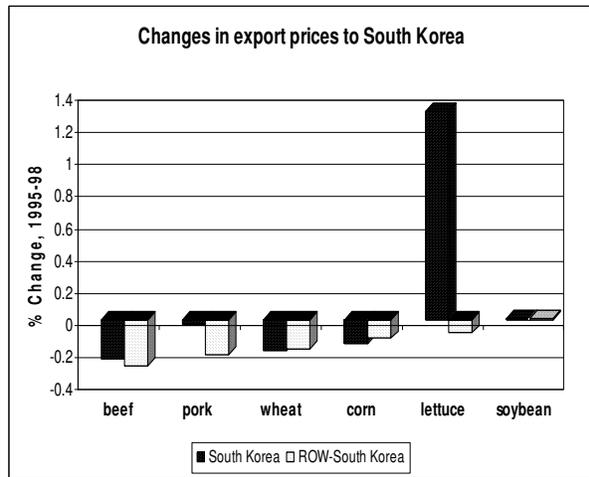
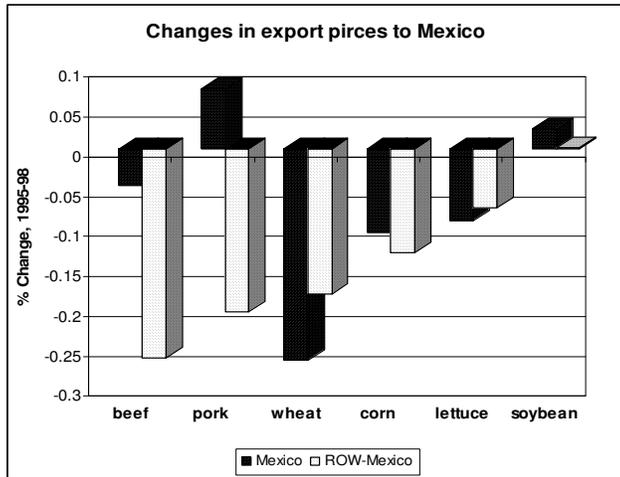
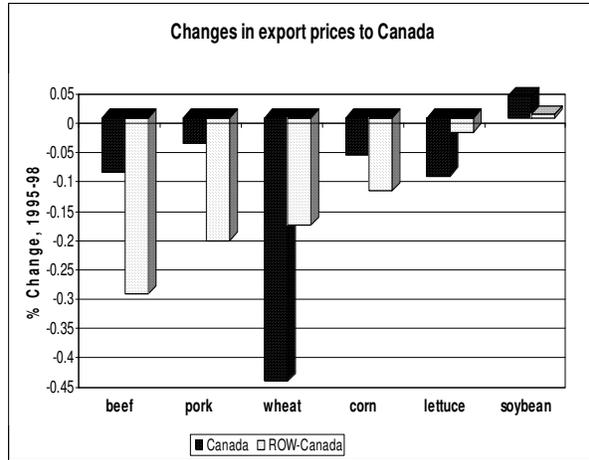
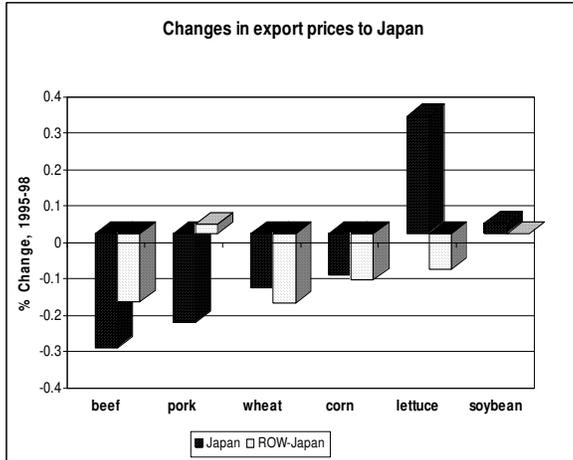
Finally, compared to the SECM and VECM models, the fixed-effects model shows that exchange rates typically have a positive effect on export unit values, and with a relatively larger magnitude. Therefore, evidence of PTM is very sensitive to model and data selection, which suggests that we need to treat the results with a certain amount of caution.

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Figure 1: Changes in US export prices between 1995 and 1998



Note: ROW-Japan (ROW-Canada, ROW-Mexico, ROW-South Korea) represent rest of the world except Japan (Canada, Mexico, South Korea). All data are annual from FATUS/ERS/USDA

Table 1: Trace test statistic for seasonal cointegration

	Rank	Nominal exchange rate			Real exchange rate		
		$\omega = 0$	$\omega = \frac{1}{2}$	$\omega = \frac{1}{4}$	$\omega = 0$	$\omega = \frac{1}{2}$	$\omega = \frac{1}{4}$
China	r = 0	29.75*	65.2*	41.77*	28.64 [†]	68.1*	47.75*
	r = 1	14.18*	32.34*	19.95*	11.29	32.47*	22.26*
	r = 2	5.9*	12.54*	2.94	2.39	11.89*	5.49*
Egypt	r = 0	29.85*	64.18*	97.67*	28.25 [†]	58.34*	95.81*
	r = 1	9.33	27.53*	44.18*	12.03 [†]	27.77*	45.6*
	r = 2	1.62	10.22*	14.72*	2.89	10.08*	14.89*
Japan	r = 0	56.02*	58.51*	57.18*	52.65*	58.24*	58.37*
	r = 1	19.23*	20.05*	34.87*	14.94*	19.12*	36.31*
	r = 2	3.15	8.98*	14.12*	3.16	8.1*	16.03*
South Korea	r = 0	19.67	64.77*	47.95*	20.2	63.09*	51.4*
	r = 1	9.15	33.28*	24.31*	9.76	31.52*	27.23*
	r = 2	1.52	10.99*	8.6*	2.15	10.72*	10.48*
The Philippines	r = 0	25.54	59.3*	44.2*	24.26	60.89*	50.91*
	r = 1	6.49	24.46*	21.37*	7.26	26.6*	27.3*
	r = 2	0.14	9.04*	4.43	0.44	9.15*	9.78*
Mexico	r = 0	33.86*	74.8*	79.93*	36.45*	74.53*	80.43*
	r = 1	10.96	34.94*	35.77*	11.98	34.09*	39.22*
	r = 2	2.26	12.2*	16.9*	5.53*	12.5*	19.15*
Jamaica	r = 0	45.23*	58.15*	38.14*	39.73*	60.05*	40.75*
	r = 1	24.43*	32.56*	12.93	17.91*	30.93*	14.95*
	r = 2	6.89*	13.04*	1.06	4.02*	12.73*	0.14
Thailand	r = 0	17.26	59.43*	38.68*	17.64	59.84*	39.13*
	r = 1	7.24	31.65*	19.52*	7.54	33.53*	19.71*
	r = 2	0.73	11.86*	3.6	1.0	11.91*	3.82

Note: no constant, deterministic seasonal dummies and a trend are included in the estimation. No lagged terms are in the model according to the information criterion (AIC, AAIC, BIC). * and [†] represent the 5 percent and 10 percent significant levels respectively. Finite-sample (50) critical values are from Lee (1992).

Table 2a: Coefficients of the cointegrating vectors for US wheat exports (normalized on export price)

Importing countries	Nominal exchange rate			Real Exchange rate		
	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)
China	1 ^a	-5.35	5.83	1	-16.09	4.08
Egypt	1	0.53	0.59	1 ^b	1.43	-0.15
Japan	1 ^c	-0.08	0.79	1 ^d	-0.08	0.77
South Korea*	1	-0.004	0.93	1	0.01	0.95
The Philippines	1 ^e	1.51	0.48	No equilibrium relationship		
Mexico	1	-0.04	1.08	No (seasonal) unit roots		
Jamaica	1 ^f	-0.65	1.55	1 ^g	-0.61	1.52
Thailand	1 ^h	-0.67	2.73	1 ⁱ	-0.95	2.97

Note: Coefficients of the second cointegration relationship are ^a 1, 0.73, -0.65; ^b 1, -0.004, 1.81; ^c 1, -1.3, 1.7; ^d 1, 1.69, 2.63; ^e 1, -5.17, 5.97; ^f 1, 0.43, -0.68; ^g 1, 0.38, -0.65; ^h 1, 0.17, 0.18; ⁱ 1, 0.16, 0.21. * Coefficients for South Korea are from VECM model with intercept restriction, where the intercept terms ($-\beta_0$) are 0.77 for nominal exchange rates and 0.58 for real exchange rates

Table 2b: Coefficients in fixed-effect model for US wheat exports

Importing countries	Nominal exchange rate		Real exchange rate	
	country-specific(λ)	ex(β)	country-specific(λ)	ex(β)
China	2.65	-0.54**	2.03	0.15
Egypt	2.2	-0.11**	2.1	0.02
Japan	2.22	0.006	2.7	-0.23**
South Korea	2.21	-0.003	1.59	0.2**
The Philippines	1.86	0.21**	1.78	0.26**
Mexico	2.15	0.05	2.12	0.07**
Jamaica	2.1	0.05	2.08	0.07**
Thailand ¹	--	1.47**	--	1.48**

Note: 1 means intercept term in this equation (country) is dropped to avoid singularity problem in estimation

** represents 5 percent significant level and * represents 10 percent significant level

Table 3a: Coefficients of the cointegrating vectors for US beef exports (normalized on export price)

Importing countries	Nominal exchange rate					Real exchange rate				
	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)
Japan	r = 0	No equilibrium relationship				r = 0	No equilibrium relationship			
Canada	r = 0	No equilibrium relationship				r = 0	No equilibrium relationship			
Mexico	r = 0	No equilibrium relationship				r = 0	No equilibrium relationship			
South Korea	r = 1	1	-0.43	0.38	9.3	r = 1	1	-0.65	0.48	10.4
Hong Kong	r = 0	No equilibrium relationship				r = 1	1	-1.05	1.81	1.94
Singapore	r = 1	1	-0.6	0.66	6.0	r = 1	1	-0.73	0.98	4.51
Switzerland	r = 0	No equilibrium relationship				r = 0	No equilibrium relationship			
The Philippines	r = 1	1	0.49	0.07		r = 1	1	0.61	-0.30	7.92
Indonesia	r = 1	1	0.08	0.11	7.28	r = 1	1	0.29	0.15	5.16

Table 3b: Coefficients in fixed-effects model for US beef exports

Importing countries	Nominal exchange rate		Real exchange rate	
	country-specific(λ)	ex(β)	country-specific(λ)	ex(β)
Japan	4.74	-0.54**	4.11	-0.27**
Canada	3.62	-0.45**	3.49	0.05
Mexico	3.59	-0.12	3.24	0.2**
South Korea	3.87	-0.11	2.88	0.19**
Hong Kong	3.27	0.37**	-1.03	5.13
Singapore	3.81	0.09	3.73	0.14
Switzerland	4.1	-0.69*	4.04	-0.62
The Philippines	2.77	0.65**	2.91	0.54**
Indonesia ¹	--	0.97**	--	1.0**

Note: 1 means intercept term in this equation (country) is dropped to avoid singularity problem in estimation

** represents 5 percent significant level and * represents 10 percent significant level

Table 4a: Coefficients of the cointegrating vectors for US pork exports (normalized on export price)

Importing countries	Nominal exchange rate					Real Exchange rate				
	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)
Japan	r = 0	No equilibrium relationship				r = 0	No equilibrium relationship			
Mexico	r = 3					r = 1	1	1.86	-0.91	7.49
Canada	r = 0	No equilibrium relationship				r = 1	1	-1.2	0.39	6.31
Hong Kong	r = 1	1	-6.9	1.53	14.37	r = 1	1	0.45	1.31	0.26
South Korea	r = 1	1	-0.45	0.62	7.85	r = 1	1	-0.52	0.64	8.32
Singapore	r = 1	1	-2.03	-0.005	9.0	r = 1	1	-1.24	-1.00	13.21

Table 4b: Coefficients in fixed-effect model for US pork exports

Importing countries	Nominal exchange rate		Real exchange rate	
	country-specific(λ)	ex(β)	country-specific(λ)	ex(β)
Japan	4.74	-0.58**	5.0	0.7**
Mexico	3.19	0.07	3.17	0.1**
Canada	3.37	-0.22	3.28	0.31
Hong Kong	2.45	0.79**	-9.92	14.68
South Korea ¹	--	1.1**	--	1.1**
Singapore	3.6	-1.05*	3.87	-2.23**

Note: 1 means intercept term in this equation (country) is dropped to avoid singularity problem in estimation

** represents 5 percent significant level and * represents 10 percent significant level

Table 5a: Coefficients of the cointegrating vectors for US corn exports (normalized on export price)

Importing countries	Nominal exchange rate					Real Exchange rate				
	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)
Japan	r = 1	1	-0.04	0.78	1.34	r = 1	1	-0.01	0.78	1.2
South Korea	r = 1	1	-0.12	0.77	2.01	r = 1	1	-0.18	0.75	2.53
Mexico	r = 1	1	0.02	0.84	0.83	r = 1	1	-0.23	0.93	1.01
Egypt	r = 1	1	-0.03	0.89	0.64	r = 1	1	-0.02	0.9	0.6
Algeria	r = 2	1	-0.02	0.81		r = 1	1	-0.05	0.78	1.35
		1	0.2	0.77						
Dominican Republic	r = 0	No equilibrium relationship				r = 0	No equilibrium relationship			
Canada	r = 1	1	-0.005	0.79	1.04	r = 1	1	-0.001	0.79	1.04
Jamaica	r = 1	1	5.11	111.62	-411.96	r = 1	1	-33.4	101.13	-218.02

Table 5b: Coefficients in fixed-effect model for US corn exports

Importing countries	Nominal exchange rate		Real exchange rate	
	country-specific(λ)	ex(β)	country-specific(λ)	ex(β)
Japan	2.02	0.002	2.29	-0.42**
South Korea	2.54	-0.16**	-0.17	0.52**
Mexico ¹	--	1.99**	--	1.63**
Egypt	2.04	-0.02	1.23	0.3**
Algeria	2.09	0.03	0.88	0.31**
Dominican Republic	1.98	0.06	0.55	0.74**
Canada	2.01	-0.05	1.14	1.51**
Jamaica	2.33	0.17**	0.96	0.31**

Note: 1 means intercept term in this equation (country) is dropped to avoid singularity problem in estimation

** represents 5 percent significant level and * represents 10 percent significant level

Table 6a: Coefficients of the cointegrating vectors for US soybean exports (normalized on export price)

Importing countries	Nominal exchange rate					Real Exchange rate				
	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)	rank	Price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)
Netherlands	r = 1	1	0.18	0.83	1.51	r = 1	1	0.19	0.83	1.52
Japan	r = 1	1	-0.07	0.78	2.23	r = 1	1	0.07	0.81	1.44
South Korea	r = 1	1	0.006	0.92	1.19	r = 1	1	0.02	0.92	1.09
Mexico	r = 1	1	-0.004	0.88		r = 1	1	0.055	0.87	1.35
Canada	r = 1	1	0.09	0.54	2.96	r = 0	1	0.08	0.53	3.01
Costa Rica	r = 1	1	-0.06**	0.95		r = 1	1	0.28	1.03	0.86

Note: ** represents 5 percent significant level for case of no constant restriction on the error term

Table 6b: Coefficients in fixed-effect model for US soybean exports

Importing countries	Nominal exchange rate		Real exchange rate	
	country-specific(λ)	ex(β)	country-specific(λ)	ex(β)
Netherlands	2.4	0.2*	2.29	0.28**
Japan	2.31	0.08*	2.08	0.15
South Korea	2.47	-0.003	1.2	0.39**
Mexico	2.39	0.07**	2.21	0.21**
Canada	2.44	0.19**	2.22	1.18**
Costa Rica	--	1.0**	--	1.03**

Note: 1 means intercept term in this equation (country) is dropped to avoid singularity problem in estimation.

** represents 5 percent significant level and * represents 10 percent significant level

Table 7a: Coefficients of the cointegrating vectors for US lettuce exports (normalized on export price)

Importing countries	Nominal exchange rate					Real Exchange rate				
	rank	price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)	rank	Price (β_1)	ex ($-\beta_2$)	ppi ($-\beta_3$)	intercept ($-\beta_0$)
Canada	r = 2	1	0.55	0.42	4.0	r = 2	1	0.41	0.42	4.05
		1	-5.15	4.27	-13.01		1	-6.4	5.6	-19.6
Hong Kong	r = 2	1	-0.14	-0.01	6.68	r = 1	1	-0.2	0.08	6.34
		1	-10.69	0.55	25.53					
Japan	r = 2	1	-19.57	-23.91	216.75	r = 1	1	-1.14	2.87	-1.76
		1	-1.23	0.08	12.33					
Singapore	r = 2	1	-0.67	0.18	5.9	r = 2	1	-0.64	0.34	5.06
		1	-1.0	3.1	-8.18		1	0.13	-1.36	13.0
Mexico	r = 2	1	0.5**	-0.92		r = 1	1	0.07	1.04	0.66
		1	0.17**	0.22						

Note: ** represents 5 percent significant level for case of no constant restriction on the error term

Table 7b: Coefficients in fixed-effect model for US lettuce exports

Importing countries	Nominal exchange rate		Real exchange rate	
	country-specific(λ)	ex(β)	country-specific(λ)	ex(β)
Canada	2.73	0.49**	1.41	2.04**
Hong Kong	2.88	-0.05	2.29	-0.6
Japan	4.63	-0.76**	3.95	-0.94**
Singapore	2.99	-0.46**	1.86	-0.22
Mexico	--	2.6**	--	1.94**

Note: 1 means intercept term in this equation (country) is dropped to avoid singularity problem in estimation

** represents 5 percent significant level and * represents 10 percent significant level