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U.S. corn exports: the role of the exchange rate

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Abstract

Time series econometric methods are applied to monthly observational data over the period 1978–1992 on real exchange rates, real corn prices, corn export sales, and corn export shipments for the United States. In-sample fit and out-of-sample forecast results are used to discern whether exchange rates have elicited systematic responses in U.S. corn prices, sales and shipments, and whether the dynamic transmission mechanisms tying these variables together have changed over time. A structural break appears to have occurred in early 1985. No cointegration is found between exchange rates, price, sales, and shipments in either sub-period. Influences are all short-run or between stationary variables. The role of the exchange rate appears to have moderated in the post-1985 period. Implications for policy analysis are discussed.

The role of the exchange rate in the determination of agricultural prices and commodity exports has been a major focus of international agricultural economics research over the past two decades. One camp argues for little or no impact of exchange rates on prices and quantities traded. Justification for this position is found in purchasing power parity theory, wherein exchange rate movements are the result of changes in the money stock, which (theoretically) have an equal and opposite impact on price movements, leaving export prices unchanged. Empirical support for the "exchange rates don't matter" (ERDM) position is provided by Batten and Belongia (1984), who argue that the real stimulus for export demand comes from income enhancements in importing countries, and by Johnson, Grennes, and Thursby, who found that trade distortions had a much greater impact on commodity flows in the early 1970s than exchange rate changes.

The agricultural economics literature has generally supported the alternative position, that exchange rates matter. According to this position, an exchange rate depreciation in an exporting country results in an increase in imports because it lowers the purchase price in the importing country (expressed in the importer's currency), at the same time raising the price received by the exporting country's producers. Proponents explain the commodity price boom of the early 1970s as being a result of the 1971 and 1973

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dollar devaluations and the change in the United States from fixed to flexible exchange rates (Schuh, 1974; Fletcher et al., 1977; Chambers and Just, 1982; Longmire and Morey, 1983). Likewise, they explain the depressed U.S. agricultural economy of the early 1980s as a result of an overvalued dollar in foreign exchange markets (Schuh, 1984; Chambers, 1984; Orden, 1986). Finally, renewed U.S. agricultural export activity in the late 1980s and early 1990s has also been attributed to a decline in the value of the dollar (MacDonald, 1992; Stallings, 1988).

Indeed, the purchasing power parity argument is difficult to defend when we consider specific commodity or product flow responses to exchange rate changes. First, response times vary across products and countries. Although agricultural markets tend to respond relatively quickly to changing economic conditions and market signals, different sets of needs and purchasing institutions in the many importing countries result in very different response sets. Second, response magnitudes vary across commodities and countries. While the same set of elasticity determinants that affect domestic demand for a product also affect export demand, export demand elasticities exceed domestic demand elasticities. In particular, heavily traded products (including many agricultural commodities) tend to have higher response rates. Finally, a monetary shock resulting in an exchange rate change touches on many sets of bilateral exchange rates. Although exchange rate (export demand) elasticities are reported for particular commodities, these calculations (estimations) are compilations of the responses of many importers, each with their own product response sets reacting to differential bilateral exchange rate changes.

In this paper we argue that "conventional wisdom" has overplayed the exchange rate link in explaining agricultural export levels. We argue that there is not an overwhelming amount of evidence to justify the general acceptance of the "exchange rates matter" in agricultural trade position, at least in certain U.S. export markets for major grains. However, purchasing power parity theory alone cannot explain the alternative position that "exchange rates don't matter." We first offer a set of supplemental arguments that go against the conventional wisdom and support a lessened role for the exchange rate in agricultural trade and price determination. We then apply vector autoregression (VAR) methods to monthly data on real exchange rates, real corn prices, corn export sales, and corn export shipments for the United States. In-sample and out-of-sample results from these methods allow us to discern whether exchange rates have elicited systematic responses in U.S. corn prices, sales and shipments, and whether the dynamic transmission mechanisms tying these variables together have changed over time.

1. Arguments in support of a lessened exchange rate role

Despite a relatively small number of empirical studies which support the "exchange rates matter" (ERM) position, this argument seems to have won the day. It may be that the difficulty defending purchasing power parity is responsible for this general acceptance. However, there are at least two additional arguments that either counter ERM or support ERDM. The first argument in favor of ERDM concerns long-term buying relationships between importers and exporters. These relationships do not change with short term price or exchange rate movements. A large portion of U.S. agricultural commodity trade involves sales to countries that have been steady importers for many years. U.S. soybean and grain embargoes in 1973, 1975, and 1980 led to an "unreliable supplier" reputation for U.S. suppliers, a reputation commodity exporters fought hard to shed during the sluggish grain markets of the early 1980s. Likewise, erratic purchasing behavior can lead to unpredictable market movements and a desire for long-term agreements.

A second argument against the ERM position has to do with structural change. When economic agents form expectations of economic activity, routine changes in the activity level have little or no impact since the outcome has already been built into the actions surrounding the expectations. Conversely, major changes in relevant institutions or programs can lead to substantial impacts on prices and trade. Most of supporting ERM studies use data from the early 1970s to the mid-1980s. Structural change was clearly present in the early 1970s in the move from fixed to flexible exchange rates. Economic agents accustomed to fixed exchange rates found themselves in a very different environment in a flexible exchange rate world. Large price and trade movements are not surprising in this context.

2. Previous work on exchange rates and agricultural exports

Schuh (1974) was among the first to look at the impact of exchange rate changes on U.S. agricultural exports. Schuh argued that the U.S. dollar's 1952–1971 overvaluation had rendered U.S. farm prices uncompetitively high on world markets, causing a decreased world demand for American farm exports. He contended that the two Nixon Administration devaluations and the ensuing world currency realignments decreased the dollar's value relative to other major world currencies, which rendered U.S. agricultural prices more competitive abroad (pp. 10–11). A number of "exchange rate impacts on agricultural trade" studies (both theoretical and empirical) were published in the decade or so following Schuh's article, including Kost (1976); Vellianitis-Fidas (1976); Johnson et al. (1977); Fletcher et al. (1977); Shei and Thompson (1979); Bredahl et al. (1979); Collins et al. (1980); Chambers and Just (1979); Chambers and Just, 1981; Chambers and Just, 1982); Longmire and Morey (1983); Batten and Belongia (1984, Batten and Belongia, 1986); and Orden (1986).

Estimates obtained using both structural econometrics and time series methods have found varying degrees of exchange rate impacts on agricultural prices and quantities traded. Probably the most oft-cited estimates are those obtained by Chambers and Just (1981). Using a simultaneous equation framework, they estimated exchange rate elasticities of corn, wheat, and soybean prices of -1.9, -1.2, and -2.6, respectively, and exchange rate elasticities of corn, wheat, and soybean exports of -5.23, -2.05 and -1.31, respectively. Although these estimates tended to be upper bounds on exchange rate estimates, they became the standard of comparison for other studies. A recent study by Denbaly and Torgerson (1992) uses time series methods that combine long-run trends with short-run dynamics. With all variables treated as endogenous, they find a wheat price elasticity with respect to the exchange rate of -1.27, equal to the level reported by Chambers and Just (1981).

3. Data and empirical methods

The studies cited above suggest that agricultural prices and exports are tied to exchange rates. Here we investigate this relationship empirically. We study the dynamic relationships among real U.S. exchange rates, the real world price of corn, U.S. corn export sales, and U.S. corn export shipments. The real exchange rate here is represented by the real index of corn-import-weighted currencies of major U.S. corn importers relative to the U.S. dollar compiled by the Economic Research Service of the U.S. Department of Agriculture (ERS/USDA), and described in Stallings (1988). The real world price of corn is the deflated U.S. Gulf price of corn compiled by ERS/USDA. We use U.S. corn export sales and shipments as compiled by the Export Sales Reporting Division, Foreign Agricultural Service of the USDA (see also Ruppel, 1984). All data were transformed into natural logarithms. Data are

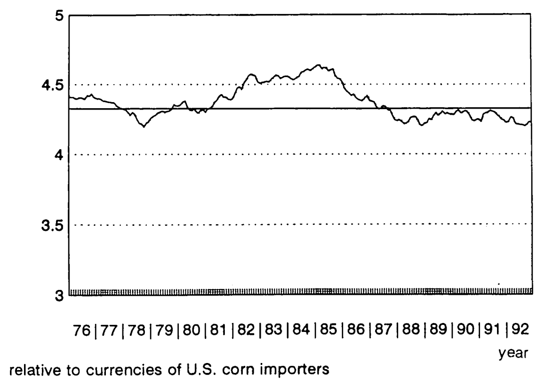

Fig. 1. Log of real exchange rates, 1976–1992.

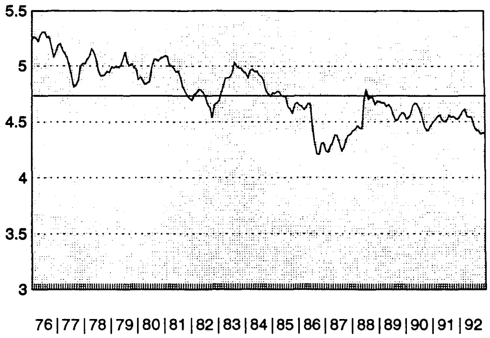

Fig. 2. Log of real corn price, 1976–1992.

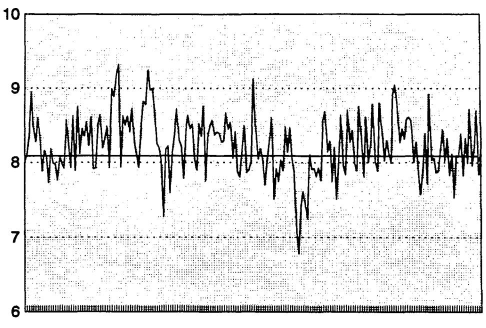

Fig. 3. Log of U.S. corn export sales, 1976–1992.

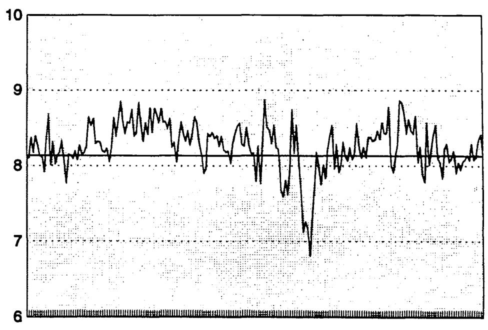

Fig. 4. Log of U.S. corn export shipments, 1976–1992.

monthly from the 1976:1 to 1992:12 period (here and throughout, months are stated numerically with 1 denoting January and 12 denoting December).

Plots of these data are provided in Figs. 1–4. Also plotted in each figure is the historical mean of the series (the horizontal line). Notice from Figs. 1 and 2 that exchange rates and price appear to be mean nonstationary; they wander, for long periods of time, away from their historical means. In fact, over the 204 data points, price and exchange rates have less than five mean crossings each. On the other hand, sales and shipments have numerous mean crossings (each crosses its historical mean over fifty times), suggesting that these series are mean stationary. We study this issue in more detail below.

4. Unit roots, cointegration, and choice of method

Our plan is to study both within sample fit and out-of-sample forecasts. We used the period 1978:2–1989:12 for estimation, saving data points 1976:2–1977:1 for the lag search and the 1990:1–1992:12 period for the out-of-sample forecasts.

Augmented Dickey-Fuller tests conducted on the data levels over the period 1978:2–1989:12 confirmed that the sales and shipment series are stationary in (logged) levels. The procedures for the τ_μ and τ_τ Dickey-Fuller (DF) tests are outlined in Dickey and Fuller (1979). We performed the augmented form of these tests (ADF tests) on logged levels of corn export sales and shipments as described in Hamilton (1994). Corn export sales generated pseudo-t values on the lagged nondifferenced levels regressor of -8.8 in both the τ_μ and τ_τ ADF tests. Corn export shipments generated pseudo-t values of -4.1 in the τ_μ and τ_τ ADF tests. Given the critical values of -2.89 (τ_μ test) and -3.5 (τ_τ test), evidence is sufficient at the 5% significance level to reject the hypotheses that corn sales and shipments are nonstationary (Fuller, 1976, p. 373).

Augmented Dickey-Fuller tests conducted on data over the period 1978:2–1989:12 suggest that the exchange rate and price series are I(1). The pseudo-t values on the lagged, nondifferenced exchange rate regressor ranged at about -1.0 for both τ_μ and τ_τ ADF tests. The pseudo-t values on the lagged, nondifferenced price levels were

-2.3 and -3.1 for the τ_μ and τ_τ ADF tests. The same ADF tests performed on the second differences suggest that the first differences are stationary.

Two of the four variables are I(0) (sales and shipments) and two are I(1) (exchange rates and real prices). Eigenvalue tests of cointegration (Johansen, 1988) were applied to levels of the four variables. These results suggest an absence of cointegration. Over the period 1978:2–1989:12 we have calculated trace statistics of 436.6 (for $r=0$), 31.3 (for $r \leq 1$), 8.21 ($r \leq 2$), and 2.24 ($r \leq 3$).[1] Using Table A2 from Johansen and Juselius (1990), we reject hypotheses (at the 5% level) that there are zero and one cointegrating vectors. We fail to reject the hypothesis that the number of cointegrating vectors is less than or equal to 2. However, as we note from the augmented Dickey-Fuller tests (above), shipments and sales are I(0); so including these in a system should result in two long-run relations (sales by itself represents one and shipments by itself represents the other) with no cointegration (see Johansen and Juselius, 1990). We considered the possibility that the two I(1) variables (exchange rate and price) were cointegrated; we rejected this hypothesis using the same tests.

Given these results, the four variables were modeled as a VAR: sales in levels, shipment in levels, differenced exchange rates, and differenced price. This gives consistent regressions (in the sense of Granger (1981)), as all variables are I(0). Likelihood ratio tests over the period 1978:2–1989:12, conducted at the 1% significance level, indicate a lag order of two on each of the four equations.[2] The resulting model suggests that levels of price and levels of exchange rates do not move the real quantities, sales, and shipments; rather it is changes in price and exchange rates which move the real quantities. Further there is no long-run "equilibrium" relationship between price levels and exchange rates or between price levels and sales and shipments.

5. Structural change

We use the above model to discern whether there has been structural change over the sample period. The VAR is estimated over the entire period with a two-period lag on each of the four endogenous variables. Recursive residuals were generated and subjected to the CUSUM and CUSUMSQ plot tests summarized in Harvey (1990). The plots were not presented here because of the large space requirement of plotting 8 different series (CUSUM and CUSUMSQ plots for four variables) over the lengthy time periods. Evidence from these plots suggest that structural change may have occurred at 1985:2. Accordingly, the sample was "split" into period 1 or the "early" period (1978:2–1985:1) and period 2 or the "recent" period (1985:2–1992:12). A Chow test (as described in Sims, 1980) was conducted on each equation to test the null hypothesis that there has

[1] We modeled the system as a vector error correction model of order $P=4$. Trace tests applied were on the rank ($r \leq P=4$) of the coefficient matrix (π) which relates the nondifferenced variable ($x(t-k)$) to current changes in each variable ($\Delta x(t)$):

$$\Delta x(t) = \mu + \Gamma(1)\Delta x(t-1) + \ldots + \Gamma(k-1)\Delta x(t-k+1) + \pi x(t-k) + \epsilon(t)$$

Here μ is a constant, the Γ matrices are short-run parameters, the π matrix a long-run parameter matrix and $\epsilon(t)$ a white noise residual. The trace test considers the hypothesis that the rank of π is less than or equal to r. The trace test with the time trend is given by:

$$-2ln(Q) = -T \sum_{i=r+1}^{p} \ln(1-\lambda_i)$$

Here λ_i are ordered eigenvalues (ordered from largest to smallest) of the matrix π and T is the number of observations. Note, there is an added cointegrating vector for each stationary variable in the system. Hence, with two $I(0)$ variables in levels, then r must be 2 without cointegration in the system. The results indicate $r=2$, suggesting that the two nonstationary variables, price and exchange rate, are not cointegrated.

[2] To account for seasonal effects, we included eleven seasonal indicator variables. Another indicator variable, valued at 1.0 in March 1980 and zero otherwise, was included in order to account for the effects of unusually large negative sales due to the Carter Administration grain embargo. The likelihood ratio test for lag selection is of the type described in Tiao and Box (1981).

been no structural change near the 1985:2 point. Test values of 1.95 for the exchange rate relation and 2.25 for the price relation exceed the critical $F(8,\infty)$ of 1.94 at the 5% significance level. No change is noted in the corn export sales and shipments equations.

Candidates for the cause of the structural change are many. Any event or set of events which occurred in early 1985 are possibilities. From the tests we note that the change appears to be in the price and exchange rate series and not in the sales and shipments series. One possible event which may have changed the price series is the debate on the 1985 U.S. Farm Bill and its subsequent adoption. Lin et al. (1995) write:

The FSA (Food Security Act) of 1985 was developed under agricultural economic conditions that demanded a change in direction for U.S. farm programs.... The goal was 'market orientation', (and) for the first time, legislation provided for future, planned reductions in annual target price minimums.

The act was adopted by Congress in December of 1985 (Ackerman and Smith, 1990), so our finding of structural change in early 1985 appears to be too early, unless of course actors in the corn market anticipated the adoption of the act. With observational data we are not able to say for sure. Candidates for change in structure in the exchange rate equation are even more difficult to identify. But the strength of the structural change in the exchange rate equation appears to be weaker than the break in the corn price equation (below this point is illustrated by similar forecast error decomposition between the two periods in the exchange rate equation and more dissimilar decompositions in the corn price equation).

We break the sample into the early and recent subsamples, and the various subperiods of Table 1 were formulated according to two considerations. First, limited degrees of freedom led to our decisions on 12-month lengths for the periods set aside for lag searches and out-of-sample validation. The 1978:2–1979:2 and 1985:2–1986:2 periods were set aside for start-up lag searches. Procedures similar to those conducted above on the full-period model prescribe specifications of one lag on the endogenous variables, a constant, and seasonal variables for both models. [3] Table 2

Table 2
Tiao and Box Likelihood ratio test values for lag structures on the early and recent subperiod models

Monthly horizon	Early model	Recent model
1	588.5	561.5
2	22.5	27.5
3	18.4	20.8
4	17.1	22.3
5	13.4	16.9
6	8.6	12.9
7	14.2	9.9
8	8.9	13.7
9	22.2	18.4
10	16.7	14.6
11	2.8	13.7

Note: The critical chi-square value with 16 d.f. is 32.0 (1%).

[3] We included an indicator variable, DUM, valued at unity for 1980:3 and zero otherwise, in the models estimated over the entire period before structural change was detected. This was to account for the extraordinary effects that generated the negative sales during March 1980. Since subperiod 2, 1985:2–1992:12 begins after 1980:3, we then deleted DUM from the recent model.

Table 1
Delineation of subperiods

	Subperiod 1: First (early) model	Subperiod 2: Second (recent) model
Complete period	1978:2–1985:1	1985:2–1992:12
Subperiods saved for Tiao-Box search	1978:2–1979:2	1985:2–1986:2
Estimation (sub)periods	1979:2–1984:1	1986:2–1991:12
Forecast period	1984:2–1985:1	1992:1–1992:12

suggests that evidence strongly suggests VARs in both periods having no more than one lag.

6. Moving average representations

One in-sample aspect of the early and recent VARs of interest here is the relative strength of influence that one variable has on itself and on other modeled variables in each system. These strengths of relationships are summarized through analysis of decompositions of forecast error variance (FEV decompositions). Critical to such decompositions is the method for treating contemporaneous innovation correlation. We follow the factorization commonly referred to as the "Bernanke ordering" (see Doan, 1990). We write the innovation vector (u_t) from the vector autoregression as: $Au_t = v_t$, where A is a 4×4 matrix and v_t is a 4×1 vector of orthogonal shocks. We considered several alternative A matrices. A factorization is "identified" (see Doan, 1990, pp. 8–10), if there is no combination of i and j ($i \neq j$) for which both A_{ij} and A_{ji} are non-zero. Several overidentifying factorizations are considered below. These are additional zero restrictions on the A matrix. Likelihood ratio tests of these overidentifying restrictions can be made to help judge the "plausibility" of a particular set. A general description of the models considered is given in equation

$$\begin{bmatrix} 1 & a & b & c \\ d & 1 & e & f \\ g & h & 1 & i \\ j & k & l & 1 \end{bmatrix} \begin{bmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \\ u_{4t} \end{bmatrix} = \begin{bmatrix} v_{1t} \\ v_{2t} \\ v_{3t} \\ v_{4t} \end{bmatrix} \tag{1}$$

Below we do not allow exchange rates to be affected by other variables in the system in contemporaneous time, so that we set $a = b = c = 0$. Further, contemporaneous shocks in shipments cause none of the other variables (at $t = 0$), so that $c = f = i = 0$. Sales may be determined by exchange rates ($d \neq 0$) and price ($e \neq 0$) in contemporaneous time. Price may be determined by exchange rates ($g \neq 0$) and sales ($h \neq 0$) in contemporaneous time, and shipments may be determined by sales in contemporaneous time ($e \neq 0$). Exchange rates and price do not affect shipments in contemporaneous time ($j = 0$ and $l = 0$). The nonzero elements of the A matrix will be investigated using the data.

Twelve models from the general specification are considered for each time period in Table 3.

Table 3
Marginal significance level (α) for rejecting over-identifying restrictions on alternative factorizations of covariance of contemporaneous innovations, for VARs from period I and II

Model	Pattern [a]	α (I,II) [b]	Model	Pattern [a]	α (I,II) [b]
1 row 2	d100	0.12 0.30	7 row 2	d1e0	0.12 0.30
row 3	gh10		row 3	g010	
2 row 2	0100	0.15 0.11	8 row 2	d100	0.00 0.00
row 3	gh10		row 3	g010	
3 row 2	d100	0.16 0.22	9 row 2	01e0	0.20 0.28
row 3	0h10		row 3	g010	
4 row 2	0100	0.19 0.09	10 row 2	0100	0.00 0.00
row 3	0h10		row 3	g010	
5 row 2	d1e0	0.12 0.08	11 row 2	d100	0.01 0.00
row 3	0010		row 3	0010	
6 row 2	01e0	0.19 0.09	12 row 2	0100	0.09 0.00
row 3	0010		row 3	0010	

[a] Pattern refers to the pattern of nonzero coefficients in the Bernanke factorization of contemporaneous covariance of innovations from the VAR. That is $Au_t = v_t$. Here we give only rows 2 and 3 from the A matrix, as row 1 = (1000) and row 4 = (0k01) are the same under all possibilities considered. See Doan for details on the general problem. [b] The α(I,II) levels are the marginal significance levels at which we reject the over-identifying restrictions implied by the A-matrix fit to the innovations from the estimated VAR model from period I and II respectively.

These twelve are alternative specifications of rows two and three of Eqn. (1); the only restriction is elements a_{ij} and a_{ji} cannot both be nonzero. There we list rows two and three from Eqn. (1) and the associated α-level at which we reject the zero restrictions for each model in each period. Model (1), for example, puts element 2,3 equal to zero ($e = 0$ in Eqn. (1)), element 2,4 equal to zero, and element 3,4 equal to zero (in addition to zero restriction on rows one and four, which are the same for all models considered in Table 3). We reject the zero restrictions in model (1) at

Table 4
Proportions of forecast error variance k months ahead allocated to innovations in respective series: early and recent models

Percentage explanation by					
Step	se	xrts	Sales	Price	Shipments
			(XRTS)		
Early period 1979:2–1984:2					
0	.015	100.00/100.00	00.00/00.00	00.00/00.00	00.00/00.00
1	.015	99.11/99.12	.07/.07	.47/.46	.35/.35
2	.015	98.86/98.87	.10/.10	.50/.49	.54/.54
6	.015	98.77/98.77	.11/.11	.50/.49	.62/.62
12	.015	98.76/98.77	.11/.11	.50/.49	.63/.62
Late period 1986:2–1992:12					
0	.016	100.00/100.00	00.00/00.00	00.00/00.00	00.00/00.00
1	.016	99.02/98.99	.20/.21	.57/.59	.21/.22
2	.016	98.26/98.20	.57/.58	.79/.83	.38/.38
6	.017	97.61/97.54	1.04/1.07	.84/.89	.50/.50
12	.017	97.60/97.53	1.05/1.08	.84/.89	.50/.50
			(SALES)		
Early period 1979:2–1984:2					
0	.153	01.35/00.26	87.61/88.05	11.04/11.69	00.00/00.00
1	.160	2.68/1.65	79.42/79.84	17.90/18.51	.00/.00
2	.161	3.47/2.46	78.53/78.95	17.97/18.58	.02/.02
6	.162	3.60/2.59	78.39/78.80	17.96/18.56	.05/.05
12	.162	3.60/2.59	78.39/78.80	17.96/18.56	.05/.05
Late period 1986:2–1992:12					
0	.146	04.37/00.85	81.60/83.04	14.02/16.11	00.00/00.00
1	.154	5.95/1.90	81.36/83.41	12.61/14.62	.07/.07
2	.156	6.41/2.24	81.06/83.24	12.42/14.40	.11/.12
6	.156	6.55/2.35	80.95/83.16	12.37/14.34	.14/.14
12	.156	6.55/2.35	80.94/83.16	12.37/14.34	.14/.14
			(PRICE)		
Early period 1979:2–1984:2					
0	.040	02.14/02.14	.00/.00	97.86/97.86	00.00/00.00
1	.042	7.47/7.45	.01/.01	92.42/92.45	.10/.10
2	.042	8.23/8.22	.02/.02	91.55/91.56	.20/.20
6	.042	8.33/8.32	.02/.02	91.39/91.40	.26/.26
12	.042	8.33/8.32	.02/.02	91.39/91.40	.26/.26
Late period 1986:2–1992:12					
0	.048	05.03/05.03	.00/.00	94.97/97.86	00.00/00.00
1	.050	4.54/4.52	.44/.45	94.61/94.63	.40/.40
2	.051	4.45/4.44	1.08/1.10	93.81/93.81	.65/.65
6	.051	4.44/4.41	1.75/1.77	93.01/93.02	.80/.80
12	.051	4.45/4.41	1.76/1.79	92.99/93.01	.80/.80

Table 4 (continued)

Percentage explanation by					
Step	se	xrts	Sales	Price	Shipments
			(SHIPMENTS)		
Early period 1979:2–1984:2					
0	.175	00.17/00.03	11.02/11.07	1.39/1.47	87.42/87.42
1	.194	.31/.12	11.20/11.25	2.93/3.04	85.56/85.59
2	.197	.48/.26	11.20/11.25	3.39/3.51	84.93/84.97
6	.198	.59/.36	11.19/11.25	3.54/3.66	84.68/84.73
12	.198	.59/.36	11.19/11.25	3.54/3.66	84.68/84.73
Late period 1986:2–1992:12					
0	.218	01.00/00.20	18.67/19.00	3.21/3.68	77.12/77.12
1	.294	1.60/.21	39.45/40.11	4.06/4.83	54.90/54.85
2	.325	2.36/.48	45.91/46.86	3.63/4.41	48.09/48.24
6	.343	3.26/.94	48.91/50.14	3.28/4.03	44.55/44.88
12	.344	3.28/.96	48.94/50.18	3.28/4.03	44.50/44.84

Decomposition of uncertainty (standard, se) at each step are given for two factorizations of contemporaneous covariance. The numerator in each column gives the partition associated with model 7 from Table 3; the denominator gives the partitions with model 9 from Table 3. Partitions associated with the other 10 models can be obtained from the authors.

the 0.12 level for the early model and 0.30 for the late model using a likelihood ratio test (as given in Doan, 1990 pp. 8–11).

Marginal α-levels for each period (I and II in Table 3) are given for eleven alternative orderings of the sales and price equations of Eqn. (1). Notice that models putting elements 2,3 and elements 3,2 equal to zero are rejected at fairly low levels of significance ($\alpha < 0.10$) in both periods, while those models not having both of the elements equal to zero are not rejected at 0.10 or lower. It is clear that contemporaneous correlation between sales and price ought not be treated as zero (from the strength of the rejection probabilities). Further, it is clear that contemporaneous shocks in exchange rates affect contemporaneous sales and/or price much more in the second period than in the first period. To see this, contrast models 3, 4, and 9. In models 3 and 4 the only difference is that exchange rates are allowed to affect sales in contemporaneous time in the former (3) and not in the latter (4). The marginal significance level at which we reject the overidentifying restrictions actually goes up when exchange rates are deleted from the sales equation in the early period ($0.19 > 0.16$). This same result is observed when we contrast models 9 and 10, only now exchange rates are allowed to affect price (not sales) in contemporaneous time in model 9. Here the drop in marginal significance is quite dramatic (0.20 to 0.00 in the early period and 0.28 to 0.00 in the later period).

Obviously, selecting a particular "true" ordering for contemporaneous correlation from observational data is a hopeless task; however the strength of the marginal significance levels given in Table 3 suggest that ordering from model 9 is a strong candidate for both periods (as the marginal significance level on this model is the highest in period one and third highest in period two). As a second alternative candidate one might seriously consider model 7, as it shows rather high α levels in both periods and differs from model 9 by allowing exchange rates and price to affect sales in contemporaneous time.

The forecast error decompositions from models 7 and 9 for both periods are given in Table 4. The numerator in each column gives the proportion of the uncertainty (standard error at each horizon) in each series (written in parentheses) attributed to the series in the column as calculated from model 7. The denominator gives a similar number; however, it is associated with model 9. A number of relationships emerge. Exchange rates have been largely exogenous to the system, with no less than about 98% of the uncertainty in exchange rates being attributed to its own innovations, under both models. Corn export sales and shipments were not greatly influenced by exchange rate in both models and in both periods. Exchange rate variation accounted for no more than 7% of sales and shipments over both periods and under both factorizations of contemporaneous correlation. Interestingly, the link between sales and shipments in the second period appears to be stronger than the link between sales and shipments in the first period. Sales accounts for 11-14% of the variation in shipments in the first period; while sales accounts for about half of the variation in shipments in the second period under both factorizations of contemporaneous correlation.

While exchange rates appeared not to (greatly) influence sales and shipments, exchange rates did have modest influence on price. In period 1, variation in exchange rates accounted for 7-8% of price variation, under both factorizations. This falls off to about 4 or 5% in the second period. Price, in turn has a more considerable influence on sales, from 10 to almost 20% in period 1 and about 15% in period 2. The results for corn coincide with the wheat-related findings of Bessler and Babula (1987): "Under the unrestricted VAR, exchange rates account for, at most, 8% of the error variance in wheat sales... exchange rates do have a considerable impact (18%) on shipments...." (p. 405).

Historical decompositions on changes in sales attributable to movements in exchange rates and price are presented in Figs. 4 and 5. The former represents the early period model (decomposition are given for the period January 1981 through January 1985); the latter represents the late period model (decompositions are given for the period June 1988 through December 1992). Each figure presents the portion of the sales series at each date that is attributable to movements in exchange rates and price. The general form of the decomposition derives from the moving average representation:

$$y_{t+j} = \sum_{s=j}^{j-1} B_s V_{t+j-s} + \sum_{s=0}^{\infty} B_s v_{t+j-s}$$

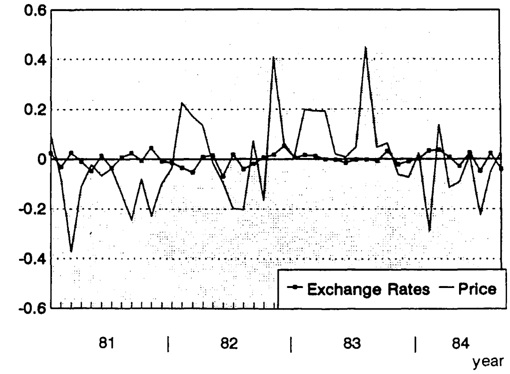

Fig. 5. Historical decompositions of sales due to exchange rates and price.

here y_t is the 4×1 vector changes in exchange rates, levels of sales, changes in price, and levels of shipments, B_s is the 4×4 moving average parameter matrix at lag s, and v_t are orthogonalized innovations. At each date, the series $y_{i,t+j}$ (series i at date $t+j$ from the y vector) can be decomposed into shocks in its own past $v_{i,t+j-s}$ as well as shocks in each of the other series of the VAR, $v_{k,t+j-s}$, $k \neq i$ (see Doan, 1990, pp. 8–16). The innovations are orthogonalized using the Bernanke ordering described above and given as model 9 in Table 3.

Notice from Fig. 5 that corn price shocks account for a much larger proportion of sales than do exchange rates over the entire 1981:1 through 1984:9 period. From Fig. 6, notice that in general the movement in sales attributable to price are much smoother, relative to those in Fig. 5. And in

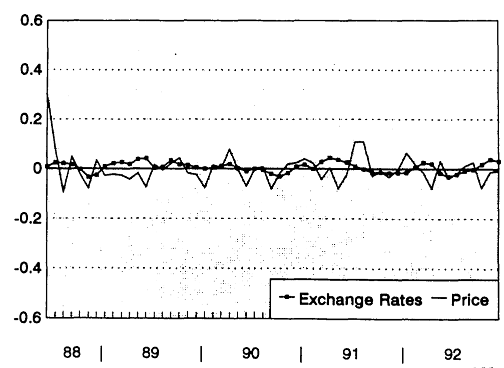

Fig. 6. Historical decompositions of sales due to exchange rates and price.

:he second period exchange rates do show a more sustained influence on export sales; both consistent with the forecast error decompositions from Table 4. Apparently, if the structure break in 1985 was due to the "1985 Farm Bill", its implication for corn export sales was to dampen the large swing in sales attributable to price. Recall from above (Table 4) that price accounts for about 5% less of export sales in the 1986–1992 period than in the 1979–1984 period. The swings are much less pronounced in the 1986–1992 period. Recall as well (from Table 4), that exchange rates account for about 3-4% more of export sales in the 1986–1992 period (relative to the

Table 5
Root mean square errors (RMSEs) on out-of-sample forecasts of series based on VAR and univariate representations

		Root mean squared error			
		Early Model		Recent Model	
Variable	k [a]	VAR	Univar. [b]	VAR	Univar. [b]
Price	1	0.051828	0.04874 *	0.035357	0.033998 *
	2	0.043202	0.04267 *	0.034974	0.033727 *
	3	0.040250	0.03786 *	0.034304	0.033771 *
	4	0.040300	0.03995 *	0.035203 *	0.036234
	5	0.041969	0.04163 *	0.026331 *	0.026686
	6	0.044575	.04453 *	0.027114 *	0.027188
	7	0.038111 *	0.03827	0.025516	0.025487 *
	8	0.041656 *	0.04179	0.025691	0.025552 *
	9	0.046640 *	0.04669	0.027437 *	0.027462
	10	0.039810	0.03979 *	0.027743 *	0.027796
	11	0.038680	0.03867 *	0.013927 *	0.014024
	12	0.052610	0.05260	0.018899 *	0.018928
Sales	1	0.205750	0.19056 *	0.147870	0.139410 *
	2	0.206780	0.19860 *	0.123200	0.117200 *
	3	0.205350	0.20050 *	0.121140 *	0.125460
	4	0.189670	0.18679 *	0.134240	0.133800 *
	5	0.195950	0.18488 *	0.136290	0.136190 *
	6	0.184900 *	0.18527	0.140990	0.140760 *
	7	0.174950	0.17480	0.150640	0.150500 *
	8	0.190780 *	0.19109	0.115860 *	0.115870
	9	0.145590 *	0.14579	0.129160	0.129040 *
	10	0.161780 *	0.16192	0.147860 *	0.147890
	11	0.153000 *	0.15305	0.112670	0.112670 *
	12	0.133080 *	0.13312	0.052070	0.052050 *
Shipments	1	0.322790	0.28867 *	0.17658	0.120740 *
	2	0.295200	0.28572 *	0.14937	0.130630 *
	3	0.316640	0.29368 *	0.13389 *	0.140720
	4	0.344680	0.33507 *	0.14330 *	0.162420
	5	0.335700	0.32136 *	0.15371 *	0.171380
	6	0.353640	0.34529 *	0.17246 *	0.182910
	7	0.287580	0.28432 *	0.17988 *	0.194810
	8	0.313540	0.30666 *	0.11091 *	0.124030
	9	0.228310 *	0.228890	0.10722 *	0.116860
	10	0.207420 *	0.211230	0.10764 *	0.115390
	11	0.017162	0.016872 *	0.05655 *	0.060410
	12	0.014933 *	0.018211	0.04583 *	0.051144

[a] k denotes the kth forecast horizon or step. [b] Univar. denotes the univariate specification's RMSEs. * Denotes which model's specification (VAR or univariate), generated the smaller RMSE, and hence the relatively more accurate forecast, at a particular horizon.

1979–1984 period), but still are less important than prices. Both of these effects are captured in the historical decompositions in Figs. 5 and 6.

7. Forecasts beyond the sample

As further evidence of the dynamic relationships among the four series in each model, we investigate the accuracy with which the VAR equations predict beyond their information sets. More specifically, we compare a multivariate VAR equation's out-of-sample forecast performance relative to that of the univariate specification of the same equation.

The early and recent VAR models were estimated over the periods of 1978:2–1984:1 and 1986:2–1991:12, respectively. The models were then forecasted over the out-of-sample periods — 1984:2–1985:1 for the early model and 1992:1–1992:12 for the recent model. Forecasts at each date were evaluated and the models updated using a Kalman filter. The root mean square errors (hereafter RMSEs) are used to evaluate performance.

As a competitor to these VAR forecasts we consider forecasts from univariate models identified and fit to both the early and late subperiods. Specification search methods similar to those applied to the VAR models were used to formulate the following univariate specifications: one or two lags on the dependent variable, a constant, and the seasonal indicator variables. The early subperiod univariate specifications included DUM (see Footnote 3).

Table 5 provides the RMSEs. Early period results show the univariate price equation predicting more accurately at most (nine) of the twelve horizons, suggesting that exchange rates have little to say in formulating price. Yet this situation changed by period 2, when the VAR price equation predicted more accurately than the exchange-rate-excluding univariate price equation at seven of the twelve equations. Hence, perhaps exchange rate has become an increasingly important determinant of price behavior. For the recent period, we conclude that there is evidence suggesting that exchange rate influences prices.

The results shown in Table 5 suggest that exchange rates have no clear influence on sales. In both periods, the univariate sales equations predict as well as or better than the multivariate sales specification. Univariate sales RMSEs are smaller at exactly half (six) of the early period horizons and are as accurate or more accurate for most (nine) of the recent period horizons. These results and the modest nature of the exchange rate contributions to sales FEV suggest that exchange rate movements have had little influence on corn export sales in either period.

The out-of-sample results fail to contradict the in-sample evidence that shipments have become more dependent on sales and less dependent on the other modeled (economic) variables. In period 1, most (nine) of the univariate RMSEs are less than the multivariate RMSEs. Yet by the second period, the multivariate shipment forecasts, inclusive of sales influences, out-predict the univariate shipments forecasts at ten of the twelve horizons. Combined with the above-noted FEV results suggesting that price and exchange rates have come to matter less, and that sales have come to matter noticeably more, in formulating/explaining shipments over time; the shipments' out-of-sample results reinforce the notion of a stronger link with shipments, and of shipments' dependence on the modeled economic variables over time.

8. Conclusions

A priori theory is not particularly helpful in answering the question of whether exchange rates matter in the determination of agricultural sales and shipments. Under a pure purchasing power parity story, exchange rates serve to adjust the real purchasing power of currencies and thus have no real effect. Accordingly, sales and shipments of real agricultural products should not be expected to respond to changes in exchange rates.

Advocates of the "exchange rates matter" position argue that the purchasing power parity argument is a long-run argument that will hold after differential inflation rates run their courses, after trade barriers are lifted, and after all other market imperfections are accounted for (Machlup, 1980). The agricultural economics literature generally follows this argument. Supporting empirical studies, however, often rely on ad hoc lag-selection procedures and within sample tests of fit. Bessler and Babula (1987) is an exception. They applied alternative VAR lag structures, of real exchange rates, real wheat price, U.S. wheat sales, and U.S. wheat shipments. Their results from in-sample fit and out-of-sample forecast performance suggest that exchange rates have little effect on wheat sales or shipments.

The fact that we find no cointegration between exchange rates, price, sales, and shipments may explain much of the debate that has gone on in the agricultural economics literature. What influences we do find are all short-run or between stationary variables. So policy analysts who look to "getting the exchange rate right for agriculture" are likely to be continually frustrated, as we can find only short-run connections between exchange rates and prices, two nonstationary variables, and sales and shipments, two stationary variables. Relying on the exchange rate to "bail-out" agriculture is likely to be met with confused results. Movement to a new level of exchange rates will have effects on price and sales in the short run, but once at the new level, there is no underlying equilibrium to keep sales at higher levels. It is only through continuous perturbations in the exchange rate (operating through the Γ matrices of Footnote 1) that policy effects on price, sales, and shipments can be maintained. Such changes in market exchange rates are not likely to be forthcoming.

It is plausible that discrete changes in exchange rates in a fixed-rate world may have set into motion a different set of events than changes in subperiods of market-determined rates. That is, one may be able to go back in time and across space and find environments when changes in policy regimes set about rather drastic changes in agricultural prices, sales, and shipments. But to extend the argument to periods of market determined rates is unwise.

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