Price Dynamics in International Wheat Markets

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A vector autoregression model is used to evaluate dynamic relationships among international wheat prices. The effects of freight rates and exchange rates are also considered. Forecast error variance decompositions and impulse response functions are used to investigate price dynamics in six important international wheat markets. The results suggest significant dynamic relationships among prices in different international wheat markets and between the prices and exchange rates and transportation costs.

On s'est servi d'un modèle vectoriel à autorégression pour évaluer les relations dynamiques entre les cours du blé sur le marché international. On a également tenu compte de l'incidence des coûts de transport et du taux de change. La décomposition prévue de la variance de l'erreur et les fonctions de réaction ponctuelle ont également permis d'approfondir la dynamique des prix sur six grands marchés internationaux du blé. Les résultats suggèrent des relations dynamiques notables entre les prix sur différents marchés internationaux du blé, de même qu'entre les prix et le taux du change et les coûts de transport.

INTRODUCTION

Considerable attention has focused upon price relationships in international grain markets. Issues of market power, government intervention, price instability and pricing efficiency have all been addressed within the context of price behavior in international grain markets. In theory, prices of a homogeneous commodity in different international markets should be closely linked through trade and arbitrage activities. In competitive markets, arbitrage should ensure that prices are perfectly linked across space such that spatial price differentials are always less than or equal to transportation costs. In this case, markets are said to be perfectly integrated. However, in reality, such linkages may be influenced by a variety of factors, including information flows, transportation costs, exchange rates, barriers to trade and the exercise of market power by large traders. In addition, because of the significant delivery lags associated with international commodity trade, markets are often slow to adjust to new information and to changes in local and global supply and demand conditions. In light of these factors, international price linkages may be imperfect and of a dynamic nature, evolving slowly over time in response to new information.

The process by which markets respond to new information is often termed "price discovery." In international markets, the commodity price discovery process may be quite complex. Trade linkages in commodity markets are influenced by financial variables, such as the exchange rate, as well as by information in related markets. In addition, transportation costs will have a direct influence on international price linkages.

The objective of this analysis is to evaluate the dynamic elements of spatial price linkages in international wheat markets by using a vector autoregression (VAR) model. The analysis examines dynamic relationships among monthly wheat prices in six important international wheat markets: the U.S., Canadian, Australian and Argentine export markets, and the Japanese and Rotterdam import markets. The analysis also considers dynamic relationships between these prices and exchange rates and transportation costs. Forecast error variance decompositions and impulse responses are used to obtain inferences regarding price dynamics in the six markets.

The exact nature of price relationships in international wheat markets has been addressed in several investigations. Roe et al (1986) and Bredahl et al (1979) evaluated price responsiveness in light of governmental interventions using price transmission elasticities. Binkley (1983) evaluated the effects of freight charges and other marketing costs on wheat price stability. Ardeni (1989), Jabara and Schwartz (1987) and Goodwin et al (1990) addressed adherence to the law of one price in international wheat markets. Market power implications for international wheat price linkages have been investigated by McCalla (1966), Carter and Schmitz (1979) and Alaouze et al (1978). Spriggs et al (1982) examined price leadership roles for the U.S. and Canadian markets. Bessler and Babula (1987) evaluated the relationship between exchange rates, wheat export sales and shipments, and prices. In nearly every case, attention has been directed toward reasons why prices may be imperfectly linked across space and thus why markets may be imperfectly integrated. However, relatively little attention has been focused upon the dynamic elements of price adjustments in international markets.[1]

International price linkages and the price discovery process may be better understood through an evaluation of the dynamic elements influencing international commodity markets. The dynamic price adjustment paths in individual markets in response to shocks to exogenous variables, such as transportation rates and exchange rates, as well as to price shocks in alternative markets, may have important implications for the overall performance of international commodity markets.

The paper proceeds according to the following plan. The following section discusses an empirical methodology that uses a vector autoregressive (VAR) model to evaluate dynamic price relationships in international wheat markets. The next section applies this methodology to a consideration of temporal relationships between prices, exchange rates and transportation rates in the international wheat market. The final section contains a brief review of the analysis and offers concluding remarks.

EMPIRICAL PROCEDURES

Vector autoregressive models are often utilized to examine dynamic relationships among a set of interrelated economic variables. Such VAR models most often utilize a set of distributed lag equations to model each variable as a function of other variables in the system. Such an approach reduces spurious *a priori* restrictions on the dynamic relationships (Sim 1972).

A VAR system for n variables can be defined as:

$$Y_t = \sum_{k=1}^{K} \begin{bmatrix} b_{11}(k) & \ldots & b_{1n}(k) \\ \vdots & & \vdots \\ b_{n1}(k) & \cdots & b_{nn}(k) \end{bmatrix} Y_{t-k} + E_t \qquad (1)$$

where

t = time ($t = 1, \ldots, T$),
Y_t = an $n \times 1$ vector of economic variables,
K = the lag order of the system,
$b_{ij}(k)$ = the parameters to be estimated, and
E_t = a vector of random errors (innovations).

Such VAR models have realized widespread usage in evaluations of dynamic relationships in economic systems (see, for example, Bessler and Brandt 1982; Featherstone and Baker 1987; Sims 1972).

To implement the VAR system, a technique for choosing the appropriate lag order, K, of the system is required. In the empirical applications that follow, the appropriate order of the VAR system is determined using the likelihood ratio test statistic (Sims 1980) for alternative lag orders. The likelihood ratio test statistic for testing between autoregressive lag orders of j and k ($j < k$) is given by:

$$L = (T - p)\,(\ln|\Sigma_j| - \ln|\Sigma_k|) \qquad (2)$$

where

T = the number of observations,
p = the number of parameters in each of the unrestricted regressions, and
Σ_j = the residual covariance matrix for the model with j lags.

The test statistic L is distributed as a chi-square variable with degrees of freedom equal to the number of parametric restrictions being tested, given by $n(k - j)$ where n is the number of endogenous variables in the VAR system.[2] The final lag order chosen is the largest for which the null hypothesis is rejected for successive increases at the 5% level of significance (Nickelsburg 1985). To verify the final choice of lag length, the Ljung-Box Q-statistic is used to test for significant residual autocorrelation.[3] In each equation, the Q-statistic adds support for the final specification in that no significant autocorrelation is detected at the 5% level of significance. In the applications that follow, data are utilized in their levels.

Inferences regarding the dynamic adjustments in each of the variables in response to unexpected shocks to the series are undertaken by converting the system to a moving-average representation. In undertaking this conversion, care must be exercised to recognize contemporaneous correlation among the individual innovations, which could lead to misleading results. To this end, an orthogonalizing transformation of the innovation vector that transforms the covariance matrix of the innovations to an identity matrix may be undertaken using a Choleski decomposition (Bessler 1984; VanTassell and Bessler 1988). This decomposition establishes a "Wold causal chain" that avoids misleading inferences caused by contemporaneous correlation of innovations.[4] This decomposition allows use of the VAR system to forecast the time path response to exogenous shocks to any one of the variables (Hakkio and Morris 1984). These time path responses, referred to as impulse responses, are used to examine the adjustments across different markets to unanticipated innovations to variables in the system. The standard errors of the impulse responses are calculated using the Monte Carlo integration techniques outlined by Doan and Litterman (1987).

The Choleski decomposition of the VAR to a moving-average representation also allows evaluation of the forecast error variance decomposition. The forecast error variance decomposition is used to determine the degree of exogeneity of a set of variables relative to another set of variables by computing the percentage of the k-steps ahead squared prediction error of a particular variable produced by an innovation in another variable (Hakkio and Morris 1984).

EMPIRICAL RESULTS

The vector autoregressive model is applied to monthly price data covering the period from July 1975 through December 1986 for six important international wheat markets. The data are chosen to include major importing and exporting markets. In particular, export markets include the U.S. Gulf market for No. 2 Hard Winter Ordinary wheat, the Canadian Pacific market for No. 1 Western Red Spring wheat (13.5% protein), the Argentine export market for Trigo Pan wheat and the Australian export market for Australian Prime Hard (APH) wheat.

Import markets include the Rotterdam market for U.S. origin No. 2 Dark Northern Spring wheat and the Japanese market for Canadian origin No. 1 Western Red Spring wheat (13.5% protein). It is important to note that quality differences may exist in these individual wheats across markets. Such quality differences could influence international price linkages if these wheat types are imperfect substitutes for one another. However, even in light of possible quality differences, the markets should be interrelated to the extent that the individual wheat types are substitutes in consumption and thus respond to global supply and demand conditions.[5]

The Japanese, Argentine and Australian prices had a small number of missing observations. In light of the importance of the time-series structure of the data series, deletion of missing observations is considered to be too strong a step. Instead, missing observations are replaced with predicted values from a regression of the individual prices on prices in closely related markets.[6] Wheat prices are collected from the International Wheat Council's *World Wheat Statistics* series. It should be noted that such quotes are asking prices and may thus imperfectly reflect actual transactions prices.

Hathaway (1987, 14) notes that most agricultural trade in the major bulk commodities is priced in U.S. dollars because of the dominant position of U.S. agricultural exports and the U.S. dollar. Therefore, all of the prices utilized in this analysis are quoted in U.S. dollars. To represent the influences of exchange rates on dollar prices in the markets, an aggregate measure of U.S. dollar exchange rates is desirable. In the following analyses, the exchange rates between the U.S. dollar and SDRs (special drawing rights) is utilized to provide a measure of exchange rate effects in each of the dollar-denominated commodity markets.[7]

Freight rates are also an important component of international price linkages for traded commodities. In a situation of perfect arbitrage, commodity prices in corresponding import and export markets should differ by no more than the transportation costs of trade between the two markets. Changes in freight rates should therefore be reflected by equilibrating changes to prices in trading markets. An aggregate measure of transportation rates is constructed by taking a simple arithmetic average freight rate across alternative ship sizes for wheat trade between the U.S. Gulf ports and Japan; between the U.S. Pacific ports and Japan; between the U.S. Gulf ports and Rotterdam; and between the U.S. Atlantic ports and Rotterdam.[8]

The VAR system is estimated using ordinary least squares. The likelihood ratio statistics initially indicate that a lag order of three months is most appropriate.[9] However, the Ljung-Box Q-statistics indicate the presence of correlation in the residuals for several of the VAR equations. The addition of a fourth lag results in Ljung-Box Q-statistics indicating that no significant residual autocorrelation is present at the 5% level in any equation.[10]

Table 1. Contemporaneous correlation coefficients of residuals for VAR system, July 1975 through December 1986[a]

	Variable							
Variable	SDRs	Shipping	U.S.	Canada	Australia	Japan	Argentina	Rotterdam
SDRs	1.00	0.06	0.23	0.16	0.20	0.22	0.15	0.07
		(0.46)	(0.00)	(0.06)	(0.02)	(0.01)	(0.08)	(0.41)
Shipping		1.00	0.27	0.30	0.32	0.24	0.11	0.31
			(0.00)	(0.02)	(0.00)	(0.00)	(0.22)	(0.00)
U.S.			1.00	0.77	0.69	0.67	0.52	0.59
				(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Canada				1.00	0.71	0.79	0.41	0.71
					(0.00)	(0.00)	(0.00)	(0.00)
Australia					1.00	0.65	0.28	0.48
						(0.00)	(0.00)	(0.00)
Japan						1.00	0.32	0.56
							(0.00)	(0.00)
Argentina							1.00	0.33
								(0.00)
Rotterdam								1.00

[a]Significance levels are reported in parentheses.

Contemporaneous correlations of the residuals of the VAR system are presented in Table 1. Of the 28 correlation coefficients, 23 are significantly different from zero at the 5% level. This indicates that a significant portion of information is reflected in price adjustments between the markets within the current month. The correlation coefficients for prices range in magnitude from 0.28 to 0.79, with most being around 0.50.[11] Residual correlation coefficients appear to be highest between individual markets and Canada and the U.S., indicating important roles for these two markets in global wheat price discovery.

Price linkages in these markets can be further investigated by considering the forecast error variance decompositions. The forecast error variance decompositions allow a consideration of which of the variables are exogenous or endogenous relative to one another at alternative forecast horizons. It should be noted that the forecast error variance decompositions are not invariant with respect to the ordering of variables in the system.

The ordering of the variables for the forecast error variance decompositions and impulse responses, reported below, are based first upon theory and second upon empirical causalities. Theoretically, the SDR variable should be the most exogenous of the factors considered (Orden 1986), followed by shipping rates.

This is consistent with the causality results. The world wheat price ordering lacks a theoretical justification and is thus based upon the causality results.[12] The overall ordering is thus selected as SDRs, shipping rates, the U.S. price, the Canadian price, the Australian price, the Japanese price, the Argentine price and the Rotterdam price.

Table 2 reports forecast error variance decompositions and standard errors for in-sample forecasts one, three, six, 12, 47 and 48 months ahead. The standard errors increase out into time but level off, thus implying a stationary system (Bessler 1984). The forecast error variance decompositions reveal a relatively large degree of price independence in most of the series.

The U.S. and Canada seem to be dominant markets. Innovations in the U.S. price have a noticeable effect on U.S., Canadian, Australian, Rotterdam, Argentine and Japanese prices. This is verified by the relatively large percentages of the forecast error variances for each of these series, which are attributable to innovations in the U.S. price. Shocks to the Canadian price have a noticeable effect on prices in Australia, Japan and Rotterdam. This effect is largest for the Japanese price. An interesting result is that forecast error variances for Argentine prices are not strongly influenced by any series except for itself, SDRs and the U.S. price.

Adjustments to innovations in freight rates appear to take place over a significant length of time for each of the series, as is evidenced by larger percentages of forecast error variances, which are explained by freight rates in the forecasts six, 12, 47 and 48 months ahead relative to the forecasts one and three months ahead. This implies that adjustments to shocks in freight rates are quite slow to occur.

Calculation of impulse responses permits an evaluation of the dynamic paths of adjustment of prices to shocks in the data series. Impulse responses for prices in the six markets generated by separate one standard deviation shocks to the exchange rate and freight rates are presented in Figure 1. Panel A illustrates price adjustments for the six wheat markets in response to a shock to the SDR exchange rate. As would be expected, a positive shock to the exchange rate between the U.S. dollar and SDRs has a positive influence on dollar prices in all of the markets. This response is significant at the 5% level in three of the six markets: the U.S., Australia and Japan. However, this effect is significant only for the contemporaneous period (and for one month after the shock for the U.S.), indicating rapid price adjustments in response to exchange rate shocks. This result underscores the forecast error variance decompositions that suggested important exchange rate influences on prices in the wheat markets.

Panel B of Figure 1 illustrates price adjustments in each of the six markets to a positive one standard deviation shock in freight rates for wheat. Each market price responds in a positive manner to increased freight rates. Such a response is expected for the CIF import market prices, but is somewhat perplexing for

Table 2. In-sample forecast error variance decompositions attributed to innovations in respective series, July 1975 through December 1986

Variable	Months ahead	Standard error	Percentage of forecast error explained by: SDRs	Shipping	U.S.	Canada	Australia	Japan	Argentina	Rotterdam
SDRs	1	0.013	100.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
	3	0.024	95.5	2.3	0.1	0.0	0.0	0.7	1.2	0.2
	6	0.034	93.8	2.4	1.1	0.0	0.8	0.6	0.6	0.7
	12	0.049	81.1	1.8	7.4	0.1	4.2	1.1	3.5	0.8
	47	0.103	32.1	1.1	24.7	14.2	5.2	0.9	18.9	2.8
	48	0.103	31.8	1.2	24.6	14.6	5.1	0.9	18.9	2.9
Shipping	1	0.899	0.4	99.6	0.0	0.0	0.0	0.0	0.0	0.0
	3	1.579	0.4	90.4	7.7	0.1	0.1	1.0	0.1	0.1
	6	2.136	1.5	79.2	14.1	1.5	0.4	1.3	1.1	1.0
	12	3.032	5.5	64.6	14.6	3.6	0.2	1.9	4.1	5.5
	47	5.000	14.2	35.4	14.5	5.6	3.2	1.2	20.8	5.2
	48	5.027	14.0	34.9	14.7	5.8	3.2	1.2	21.0	5.2
U.S.	1	4.854	5.2	6.6	88.2	0.0	0.0	0.0	0.0	0.0
	3	9.016	3.9	7.8	85.9	0.3	1.0	0.0	0.1	1.0
	6	12.288	3.3	10.5	76.2	0.2	3.9	0.1	0.5	5.2
	12	15.354	3.2	10.6	64.0	2.2	5.7	0.8	4.4	9.2
	47	22.624	15.3	10.4	38.8	10.8	6.7	1.0	8.9	8.1
	48	22.830	15.3	10.2	38.7	10.8	6.8	1.0	9.3	7.9
Canada	1	4.539	2.3	10.4	49.8	32.6	0.0	0.0	0.0	0.0
	3	10.038	2.3	10.4	49.8	32.6	3.7	0.0	0.6	0.5
	6	13.855	1.9	21.5	47.4	22.2	5.7	0.4	0.5	0.4
	12	17.618	2.1	24.4	46.8	14.7	6.7	2.4	1.0	2.0
	47	25.313	11.5	22.1	31.2	14.2	6.2	2.1	6.5	6.2
	48	25.499	11.6	21.8	31.2	14.1	6.3	2.1	6.9	6.1

Table 2. Continued

Variable	Months ahead	Standard error	Percentage of forecast error explained by: SDRs	Shipping	U.S.	Canada	Australia	Japan	Argentina	Rotterdam
Australia	1	6.779	4.1	9.5	36.3	7.1	43.1	0.0	0.0	0.0
	3	10.164	2.8	13.8	47.1	12.0	23.0	0.7	0.0	0.5
	6	12.963	2.2	18.6	49.4	9.3	17.9	0.8	0.4	1.4
	12	15.898	1.9	19.8	47.8	8.5	14.3	2.4	2.0	3.4
	47	22.738	11.9	17.8	31.9	12.8	10.3	1.8	6.9	6.6
	48	22.915	12.0	17.5	31.9	12.7	10.2	1.8	7.3	6.5
Japan	1	6.078	5.0	5.2	35.2	18.1	0.9	35.6	0.0	0.0
	3	11.755	3.1	8.6	41.2	25.7	4.4	15.9	0.1	0.9
	6	16.133	2.9	18.4	42.5	16.2	7.6	11.1	0.1	1.3
	12	21.537	2.7	22.7	41.7	10.0	8.4	10.8	0.13	3.6
	47	31.489	10.4	23.1	27.7	11.7	6.7	6.6	6.3	7.6
	48	31.704	10.5	22.8	27.7	11.6	6.7	6.5	6.7	7.5
Argentina	1	6.429	2.3	1.0	24.0	0.0	1.4	0.1	71.1	0.0
	3	11.471	4.7	1.9	34.5	0.0	0.7	0.2	57.3	0.6
	6	14.245	5.0	5.7	42.9	0.8	1.7	0.4	42.8	0.6
	12	19.088	8.8	8.8	41.9	5.2	1.9	0.2	28.7	4.5
	47	33.796	24.5	6.7	26.9	9.7	6.6	1.1	19.9	4.5
	48	34.200	24.2	6.5	27.1	9.9	6.6	1.1	20.1	4.5
Rotterdam	1	5.865	0.5	9.6	27.5	14.5	0.5	0.0	0.1	47.3
	3	10.859	2.2	14.1	39.2	16.0	2.8	0.0	0.8	24.9
	6	14.550	2.5	21.5	45.5	9.9	3.9	0.5	0.8	15.3
	12	17.731	3.4	20.2	45.9	7.4	6.1	1.9	1.6	13.5
	47	24.741	13.3	18.2	31.7	10.7	6.3	1.6	7.1	11.1
	48	24.928	13.4	17.9	31.7	10.6	6.3	1.6	7.4	11.0

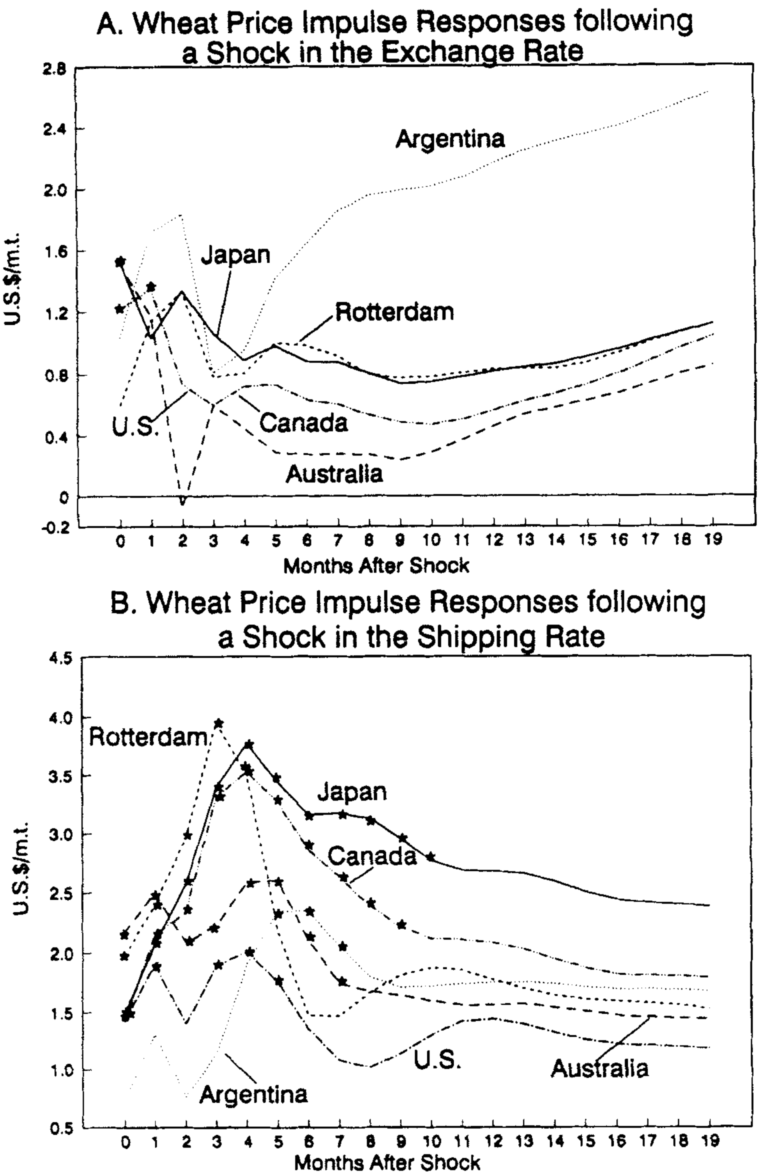
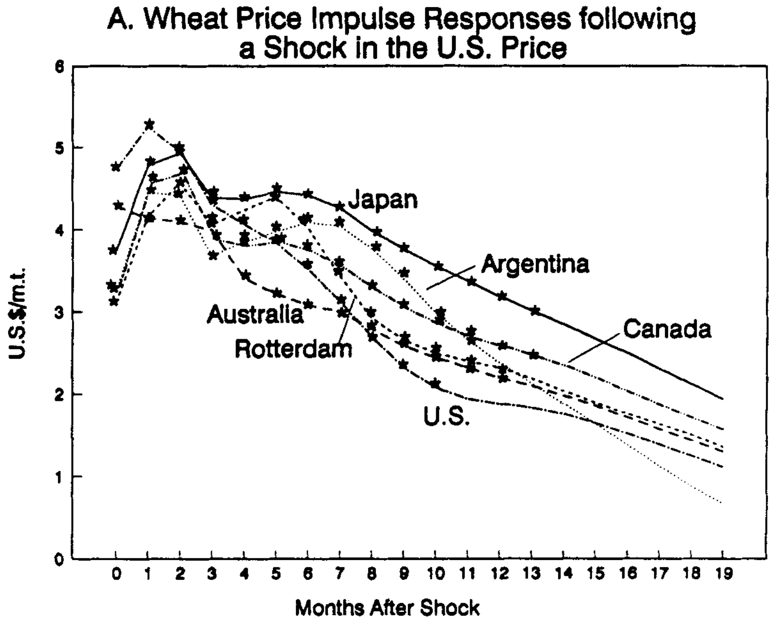


Figure 1. International wheat price impulse responses to one standard deviation shocks in exchange rates and shipping rates
*indicates response significantly different from zero at the 0.05 level

for FOB export prices. A possible explanation might be that higher global freight rates are coincidental with higher domestic transportation charges (i.e., from



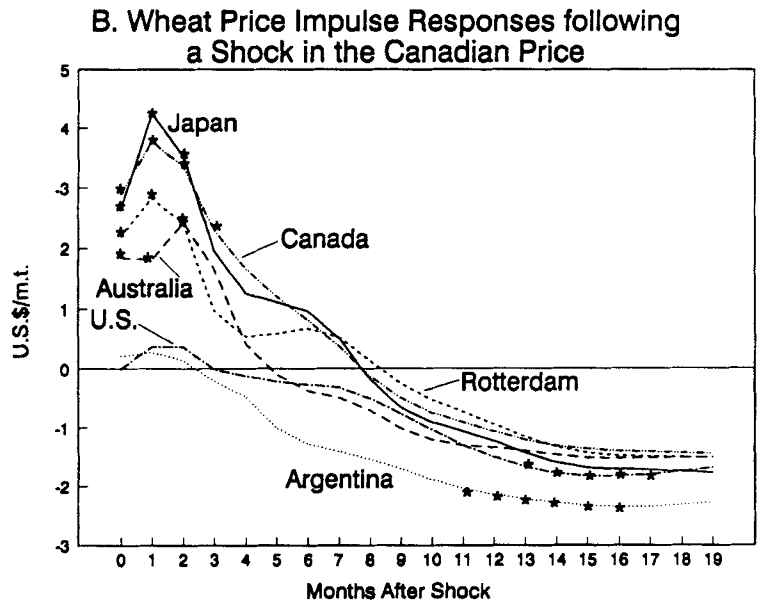


Figure 2. International wheat price impulse responses to one standard deviation shocks in U.S. and Canadian export wheat prices
*indicates response significantly different from zero at the 0.05 level

hinterland to export markets) leading to higher prices at export markets. The largest responses appear two to four months following the shocks. Interestingly, significant lags in responses to the freight rate shocks are apparent in each of the price series. In particular, large price responses to increased freight rates persist until eight or 10 months after the shock, which is consistent with the results of the error decompositions. This may result from the fact that international transactions are often contracted several months in advance, significantly slowing price responses. The response for Argentina persists longer than that for any other market, perhaps suggesting greater fixities in shipping arrangements for wheat exports from Argentina.

Panels A and B of Figure 2 illustrate price adjustments for each of the six markets in response to separate exogenous one standard deviation shocks in the U.S. and Canadian prices, respectively. Significant responses are revealed in each of the markets through 10 to 13 months following the shock to U.S. prices. The greatest response in each of the markets occurs in the first and second months following the initial shock. Significant positive responses to shocks in the Canadian price (Figure 2, panel B) are revealed for all of the countries except the U.S. and Argentina. The positive responses are not significant at lags greater than three months. The U.S. and Argentina have significant negative responses at lags of 11 to 17 months.

Panels A and B of Figure 3 illustrate price adjustments for the six markets in response to exogenous one standard deviation shocks in the Australian and Japanese prices, respectively. Significant positive responses to a shock to the Australian price are realized in every market except the U.S. and Argentina. These responses are significant for lags of two months except for the Japanese market, which had significant responses for five months following the shock. Japan is an important market for Australian wheat exports. The persistence of the Japanese response may reflect contractual arrangements between traders in Japan and Australia. A shock to the Japanese price brings about a significant response for the Japanese price only. This response is significant only for the contemporaneous period and for one lag.

Figure 4 contains price adjustments for the individual wheat markets in response to one standard deviation shocks to the Argentine and Rotterdam markets. Panel A illustrates the wheat price responses to a one standard deviation shock to the Argentine price, while panel B contains responses to a Rotterdam price shock. The responses for Argentine and Rotterdam price shocks are similar to those for Japan in that significant positive responses occur only for the markets whose prices are shocked. The responses are also rapid. The Argentine price responds significantly to a shock to itself for three lags, while the Rotterdam price responds significantly to an own-price shock for one lag. These results suggest that the Japanese, Argentine and Rotterdam markets exert limited influence on prices in other markets. Furthermore, own-price responses are short-lived in nature.

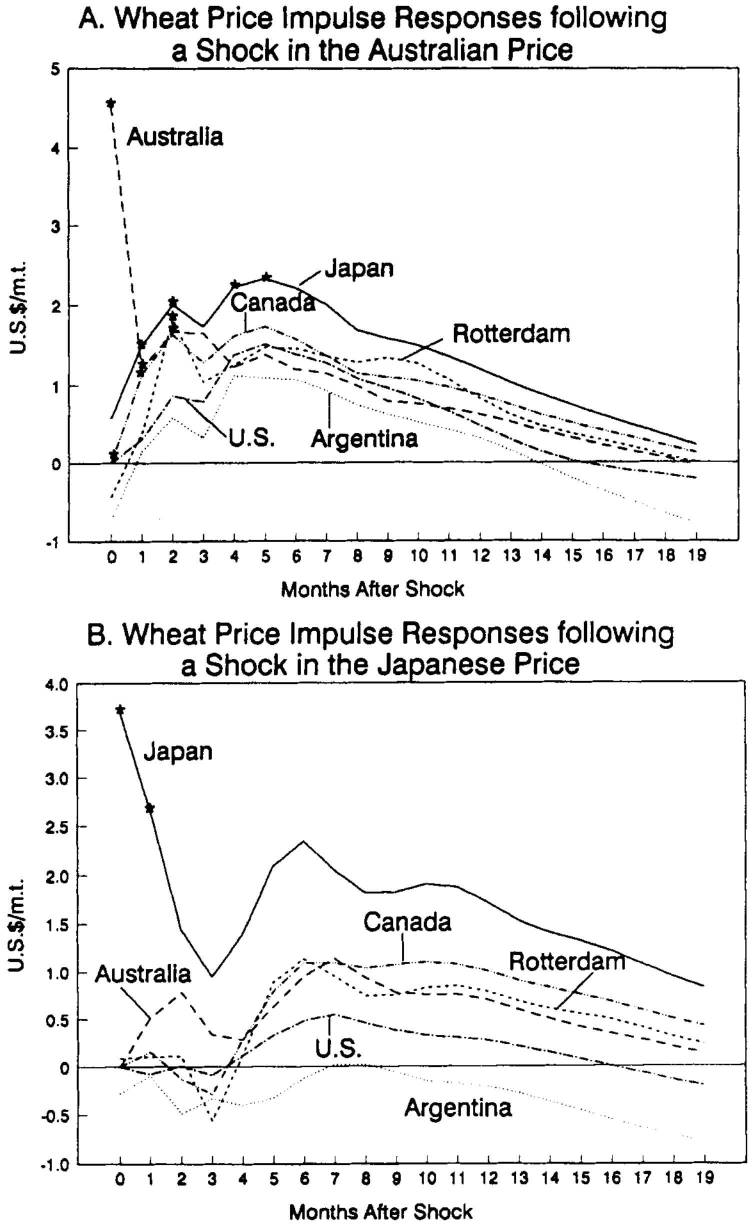


Figure 3. International wheat price impulse responses to one standard deviation shocks in Australian export and Japanese import wheat prices
*indicates response significantly different from zero at the 0.05 level

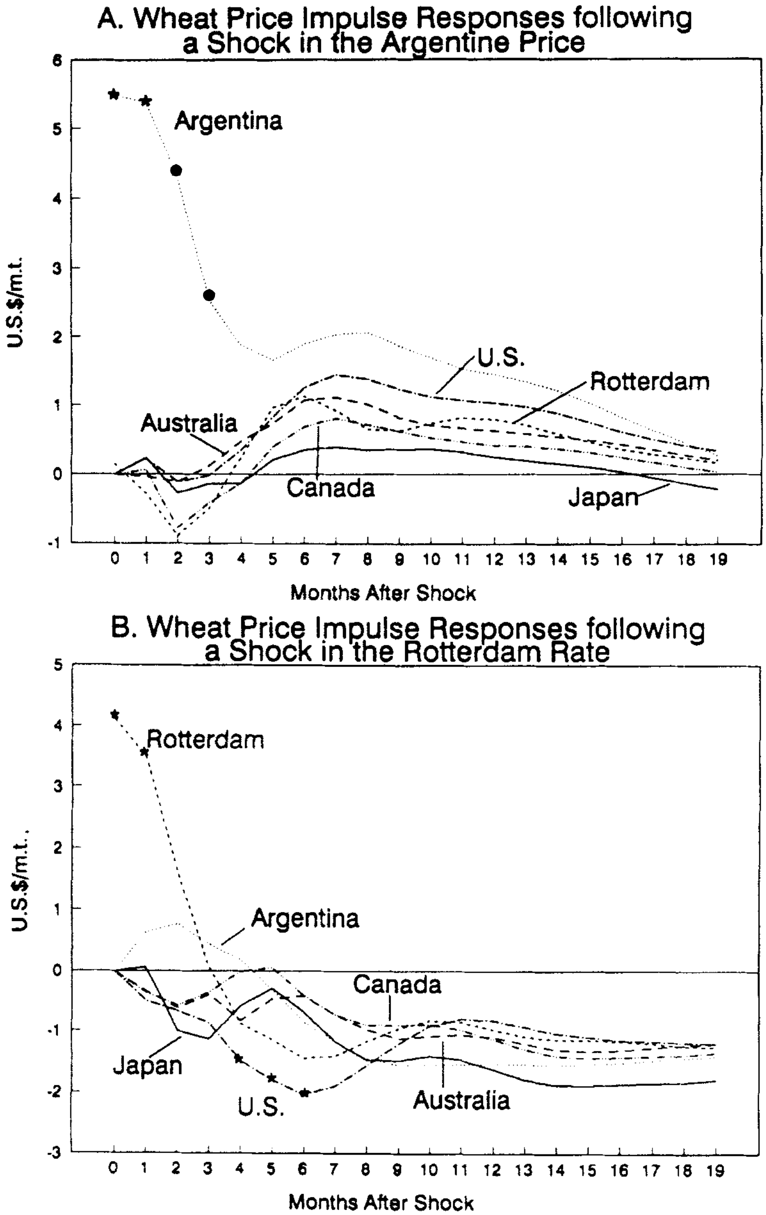


Figure 4. International wheat price impulse responses to one standard deviation shocks in Argentine export and Rotterdam import prices
*indicates response significantly different from zero at the 0.05 level

CONCLUDING REMARKS

This study examines dynamic elements of spatial price linkages in international wheat markets. The analysis considers relationships among monthly wheat prices in six important international wheat markets and evaluates interactions between these prices and exchange rates and transportation costs. A vector autoregressive model is estimated, and forecast error variance decompositions and impulse response functions are used to examine dynamic elements of the price discovery process in the international wheat market.

Forecast error variance decompositions indicate that a large proportion of forecast error variances for prices in wheat markets are explained by transportation costs and prices in the U.S. and Canada. Forecast error variances for Argentine wheat prices are explained only by the Argentine and U.S. prices. The proportion of forecast error variances explained by shipping rates rises as the forecast horizon increases.

Innovations in the U.S. price have a significant effect on prices in the other markets. Price adjustments in response to a shock to U.S. prices are strongest in the month immediately following the shock and persist for 10 or more months. Canadian prices also exert significant influences on prices in the other markets, except for the U.S. and Argentina.

Innovations to exchange rates produce significant positive responses to prices in the U.S., Australia and Japan. Adjustments to exchange rate shocks are rapid, occurring within a single month. Freight rate shocks produce long-lasting price adjustments in each of the wheat markets. Innovations in the freight rate series take two or more months to produce a large response in the price series. This adjustment persists for eight or more months after the shock. This delayed adjustment is likely due to the fact that international transactions are commonly contracted several months ahead.

Future research may benefit from greater attention to alternative variables that may influence spatial price linkages in international markets. In this light, recent work (Orden 1986) identified important roles for financial variables such as monetary growth rates and interest rates in international trade linkages. The inclusion of such variables in a VAR system with various international prices may further clarify temporal relationships in international wheat markets.

NOTES

[1]An exception is the work of Spriggs et al (1982), who addressed lead/lag relationships between daily U.S. (Minneapolis) and Canadian (Thunder Bay) wheat prices. Our work differs from that of Spriggs et al in that we give greater attention to the dynamics of international price adjustments over time and we consider a broader range of export markets and related variables.

[2]In this analysis, we evaluate successive one-period increases in the lag order, implying $k - j = 1$. Thus, the likelihood ratio test statistic has eight degrees of freedom.

[3]The Ljung-Box *Q*-statistic is a chi-square random variable with degrees of freedom equal to the number of lags being evaluated. Details regarding the application of the Ljung-Box *Q*-statistics are presented in the text below.

[4]Results for impulse responses and forecast error decompositions are dependent upon the ordering of variables in this causal chain. In the applications, the variables are ordered as: SDRs, shipping rates, the U.S. price, the Canadian price, the Australian price, the Japanese price, the Argentine price and the Rotterdam price. Justification for this ordering is presented below.

[5]To verify that the chosen prices are representative of each market's wheat prices, correlation coefficients between the individual wheat chosen for use in the VAR and alternative wheats in each market are evaluated. For Australia, the correlation coefficient between Australian Prime Hard and Australian Standard White wheats is 0.96. In Japan, correlation coefficients between U.S. origin Hard Spring wheat, U.S. Hard Winter Ordinary wheat, Australian Prime Hard wheat and Canadian Western Red Spring wheat range from 0.96 to 0.99. In the U.S., the correlation coefficient between Dark Northern Spring and Hard Winter Ordinary wheats (both at the U.S. Gulf) is 0.94.

[6]In particular, 11 missing Japanese prices are replaced with predicted values from a regression of Japanese prices on Rotterdam prices. Price data for Australian Prime Hard wheat are unavailable for July 1975 through June 1977 and are replaced by predicted values from a regression of APH wheat prices on Australian Standard White wheat. Five missing values of the Argentine prices are proxied by a regression of Argentine prices on U.S. prices. Details regarding the proxy measures of these missing prices are available from the authors on request.

[7]Monthly average exchange rates between the U.S. dollar and SDRs are from the International Monetary Fund's *International Financial Statistics* series.

[8]Monthly freight rates for wheat trade between these markets are from the International Wheat Council's *World Wheat Statistics*.

[9]A lag order of three is the largest for which the likelihood ratio test statistic is rejected at the 5% level of significance.

[10]The residual series are evaluated at a lag of 33. The test statistics range from 19.16 to 42.00, all of which are below the chi-square critical value of 47.39 at the 0.05 level with 33 degrees of freedom. Thus, no evidence of residual autocorrelation is detected in the estimated VAR system. The out-of-sample forecasting performance of the VAR model is evaluated by evaluating Theil's *U*-statistic for Kalmanized forecasts over the period covering January 1981 through December 1986. Theil's *U*-statistic compares out-of-sample predictability of the model with naive forecasts. Forecasts are evaluated at one- through 24-month horizons. Each equation generates Theil *U*-statistics that equal or are less than one at most of the horizons. Theil's *U*-statistic also appears to fall at longer horizons, indicating satisfactory out-of-sample predictability.

[11]Strong contemporaneous correlation may indicate that markets respond rapidly (i.e., within the one-month sampling interval) to new information. Alternatively, this result may imply the absence of significant information flows across markets. In the following analysis, it is important to recognize this limitation of the empirical analysis.

[12]In-sample and out-of-sample causality results suggest a causal ordering as follows: SDRs, freight rates, the U.S. price, the Canadian price, the Australian price, the Japanese

price, the Argentine price and the Rotterdam price. Because the causality tests do not indicate a clear direction of causality between the U.S. and Canadian prices, an alternative ordering that reverses the order of these two prices is also considered. This ordering gives results that are similar to those presented. Details regarding the causality tests and results for alternative orderings are available on request.

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