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A Reexamination of Price Dynamics in the International Wheat Market

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ABSTRACT

This paper examines price relationships in the international wheat market using a cointegration and error correction approach. Price series are found to be first difference stationary and cointegrated. Our results provide evidence that the United States, Australia, the European Union, and Argentina react to Canada's pricing decisions whereas Canada does not respond to any other price changes except Australia's. Similarly, the United States also plays a strong role in the pricing of other exporters (except Canada) and its U.S. pricing decision is affected by changes in Canadian and Australian prices. The remaining exporters such as the European Union and Argentina respond to the U.S. and Canadian price changes but do not have any influence on other prices. Overall, our results suggest that there is no distinct price leader in the international wheat market.

A REEXAMINATION OF PRICE DYNAMICS IN THE INTERNATIONAL WHEAT MARKET

Introduction

The international wheat market has received much attention from researchers because of the perception that noncompetitive elements exist in the market. This perception has been reinforced by the small number of major exporting regions (the United States, Canada, Argentina, Australia, and the European Union) that account for more than 80 percent of total exports. The implications of this market structure for trade, prices, and policies have been widely studied (McCalla 1966; Alaouze et al. 1978; Paarlberg and Abbott 1986; Wilson 1989; Goodwin and Schroeder 1991; Mohanty et al. 1995; Mohanty et al. 1996).

Most of the research focuses on two broad areas. One group of studies uses structural models, which incorporate noncompetitive and institutional behavioral assumptions, to analyze trade flows, policy effects, and prices (McCalla 1966; Alaouze et al. 1978; Paarlberg and Abbott 1986). The other group of studies analyzes the pricing mechanism, particularly export price determination, to address issues regarding market structure, pricing efficiency, and government interventions (Ardeni 1989; Goodwin and Schroeder 1991; Goodwin 1992; Goodwin and Smith 1995; Mohanty et al. 1995; Mohanty et al. 1996).

McCalla (1966), using a noncolluding oligopoly equilibrium model, concluded that the international wheat market was a duopoly dominated by Canada and the United States, with Canada acting as the price leader. Alaouze et al. (1978) extended McCalla's original duopoly to a triopoly by including Australia. They argued that Canada remained as price leader in the triopoly while the United States and Australia assumed price follower roles. Subsequently, the duopoly and triopoly structures have been extended to an oligopoly because of the prominence of Argentina and the European Union (EU) in the wheat export market. Numerous later studies have characterized the international wheat market as an oligopoly (Paarlberg and Abbott 1984; Kolstad and Burris 1986).

Most of the studies cited above have used little or no empirical evidence to support their theoretical models. By contrast, a second group of studies based most of their results on empirical evidence by analyzing price relationships in the international wheat market (Spriggs et al. 1982; Perkins et al. 1984; Gilmour and Fawcett 1987; Goodwin and Schroeder 1991; Goodwin and Smith 1995; Mohanty et al. 1995; Mohanty et al. 1996). Price determination in the international wheat market has received special attention because of the small number of exporting regions as well as the presence of large multinational

grain trading companies and the fact that wheat is not an entirely homogeneous product (Mohanty et al. 1995).

Spriggs et al. (1982), Gilmour and Fawcett (1987), and Mohanty et al. (1996) have limited their analysis to U.S. and Canadian prices. Spriggs et al. (1982) utilized the Granger causality test to examine the price leadership role between daily U.S. and Canadian wheat prices. They concluded that there was no significant price leadership role between the U.S. and Canada. Similarly, Gilmour and Fawcett (1987) analyzed U.S. and Canadian wheat prices using the price reaction function and achieved results comparable to Spriggs et al. (1982). However, these studies utilized internal prices rather than export prices. Recently, Mohanty et al. (1996) analyzed export price relationships between the U.S. and Canada for Durum and Hard Red Spring wheat. Using an error correction representation, they found a strong price leadership role for Canada, both in the Durum and Hard Red Spring markets.

The remaining studies cited in this paper have extended their analysis by including other major exporters such as Australia, Argentina, and the European Union. Goodwin and Schroeder (1991) estimated a vector autoregression model and impulse response functions, and found that both the United States and Canada have a significant effect on the prices of competing exporters. They also found that while the U.S. price has a significant effect on the Canadian price, the Canadian price does not affect the U.S. price; thus, they conclude that the United States is the price leader in the international wheat market. Mohanty et al. (1995) using a Granger causality test found that the U.S. price affects the prices of competing exporters (Canada, Australia, Argentina, and the European Union), but the reverse is not true, i.e., the U.S. price is not affected by changes in prices of competing exporters. Recently, Goodwin and Smith (1995) concluded from Granger causality tests and impulse responses that the Canadian wheat export prices were not responsive to changes in the U.S., Australian, and Argentinean wheat export prices, but that these prices were, in turn, responsive to changes in the Canadian wheat export price.

Inconsistency in past studies regarding price relationships in the international wheat market can be attributed to various factors including frequency of data, time period under consideration, and specification of model. However, most past studies failed to recognize that the price series used in the analysis are, a priori, potentially nonstationary integrated variables. As suggested by Granger (1983, 1986) and Engle and Granger (1987), standard Granger causality tests provide misleading results if prices are cointegrated. In addition, the structural vector autoregression and impulse response function approach used by Goodwin and Schroeder (1991) and Goodwin and Smith (1995) is likely to be misspecified, which would raise questions about the validity of the results. Mixed results from past studies may be due to the failure to consider the possibility that the variables are nonstationary and cointegrated.

In this study, we estimate the price relationships using a general error correction approach developed by Engle and Granger (1987). The analysis examines relationships among wheat prices for five major international wheat exporters: the United States, Canada, Australia, the European Union, and Argentina. Before we estimate the error correction model (ECM), we first establish the properties of the individual time series prior to testing for cointegration. In the second step, we estimate the cointegrating vector using Johansen's maximum likelihood procedure. In the third step, we estimate the error correction model by including the error correction term from the Johansen procedure.

Model Specification

To specify the price formation model, we employ a general model developed by Engle and Granger (1987) in order to establish the long-run equilibrium relationship and identify short-run dynamic specification:

$$\Delta P_{i,t} = \sum_{s=0}^m \alpha_s \Delta P_{i,t-s} + \sum_{j=1}^4 \sum_{s=1}^n \beta_{js} \Delta P_{i,t-s} + \psi Z_{i,t-1} + \eta_{i,t} \quad (1)$$

where P_i denotes the logarithm of price of i th competing exporter, $i, j = \text{United States (us), Canada (ca), Australia (au), the European Union (eu), and Argentina (ar)}$, and $i \neq j$. $Z_{i,t}$ is the estimated error correction term obtained from the i^{th} cointegrating equation. The cointegrating equation for i^{th} exporting country is as follows:

$$P_{i,t} = a + \sum_{j=1}^4 b_j P_{j,t} + \epsilon_{i,t} \quad (2)$$

where $i \neq j$. In equation (1) the parameter ψ is the coefficient of the error correction term. If the variables employed in equation (1) are cointegrated, then the error correction model exists, i.e., $\psi \neq 0$ (Engle and Granger 1987).

Data and Estimation

The data used for analysis are monthly quoted FOB prices for the period from January 1981 to June 1993. The specific price series for the major competing exporters includes U.S. #2 Hard Red Winter, ordinary protein, Gulf; Canadian #1 Western Red Spring, 13.5 percent, St. Lawrence; Australian Standard wheat; Standard wheat, Rouen to specified zones (EU); and Trigo Pan (Argentina). These representative wheat types are similar to the series selected by Goodwin and Schroeder (1991) and Mohanty et al. (1995), who also examined price relationships in the world wheat market. The primary data sources are various issues of *World Wheat Statistics*, published by the International Wheat Council. Prices for recent years were obtained from *World Grain Statistics*, also published by the International Wheat Council. All prices are quoted in U.S. dollars.

A common criticism of efforts to analyze wheat price relationships is that quoted prices do not accurately reflect true market prices because of subsidies and other special sales arrangements (Mohanty et al. 1995). In particular, the lack of transparency in the pricing behavior of the Canadian Wheat Board (CWB), which has a monopoly on the marketing of Canadian wheat both domestically and internationally, is often thought to give rise to large and unsystematic divergences between the prices actually realized in market transactions and the published price series most commonly used in econometric analysis.

Despite such criticisms, past studies such as Goodwin and Schroeder (1991), Mohanty et al. (1995), and Goodwin and Smith (1995) have used quoted prices in analyzing price relationships in the international wheat market on the basis that there is likely to be some relationship between quoted and realized wheat prices. However, because it is impossible to resolve this question in the absence of complete data on actual prices at which transactions have taken place over the period being analyzed, some caution must be exercised in interpreting the results of our analysis.

Before proceeding with estimation, one other aspect needs to be discussed: the implication of quality differences in wheat across markets. Wheat from one exporter differs from another in terms of its protein content, variety, and foreign material content. These quality differences, particularly protein content, could influence international price linkages by making varieties of wheat imperfect substitutes for one another. Despite such product differences, it has been argued by Goodwin and Smith (1995) and Mohanty et al. (1995) that the market should be interrelated to the extent that the individual wheat types are close substitutes in consumption and thus respond to global supply and demand conditions.

Cointegration is a necessary and sufficient condition for the representation of a series in an ECM. Prior to testing for cointegration, the price series are first tested for their order of integration, since a necessary condition for cointegration is that the series have to be integrated of the same order. To test for their order of integration, the Augmented Dickey Fuller (ADF) test was used. The ADF test is based on the following regression equation:

$$\Delta P_{i,t} = \alpha_{i,0} + \sum_{s=1}^m \beta_{is} \Delta P_{i,t-s} + \varepsilon_{i,t} \quad (3)$$

where Δ is the first difference operator and ε is a stationary error term. The number of lags to include in the equation was determined using the Akaike information criterion. The importance of including a constant term without a time trend has been addressed by Dickey et al. (1986). Based on their suggestions, ADF equations were estimated with an intercept and no time trend. All the price series are expressed in logarithmic form.

Results

Table 1 presents the results of Augmented Dickey Fuller unit root tests for each price series. The hypothesis tests are based upon the comparison of calculated statistics with the critical McKinnon (1991) statistics. The tabulated statistics indicate the order of integration for each test variable. Based on the critical values calculated from McKinnon (1991), unit roots can not be rejected for the level of all price series at the 5 percent significance level, but the unit root hypothesis was rejected for all the price series expressed in first difference at the same significance level. Because determination of the lag order using statistical tests alone has been criticized, in this study the ADF test was conducted using different lag orders. These alternative representations did not alter the results discussed above. Overall, it may be reasonable to conclude that price series are integrated of order one or I(1).

The determination of the presence of unit roots in price series indicates that there may exist a long-run relationship among such price series. In particular, if the time series are found to be integrated of order one I(1), then there may exist a long-run relationship among them. In order to investigate this possibility, it is necessary to carry out a cointegration analysis of the price series. Establishing the existence of cointegration implies that a stable long-run relationship holds among the price series and the researcher interested in investigating short-run dynamics among the prices may employ the error correction estimation technique. If, on the other hand, there is no evidence of cointegration, then the first

difference VAR estimation is appropriate. The same conclusion applies for the time series data which are not integrated of order one I(1).

Having confirmed that the price series are first differenced stationary, we can proceed with the cointegration tests. Cointegration is tested using Johansen's maximum likelihood procedure based on the following error correction representation:

$$\Delta P_t = \sum_{j=1}^k \alpha_j \Delta P_{t-j} + \theta(r) P_{t-1} + v_t \quad (4)$$

where P_t is a 5x1 vector of I(1) processes. This formulation, with a one-period lag on the vector of levels, is equivalent to Johansen's equation, where the levels are lagged k periods (Baghestani and McNown, 1992). The rank of $\theta(r)$ equals the number of cointegrating vectors, which is tested by the maximum eigenvalue and trace test statistics. The critical values for these statistics are obtained from Johansen (1988) and Johansen and Juselius (1990). The null hypothesis is that there are r or fewer cointegration vectors. The alternative hypotheses are $r+1$ and at least $r+1$ cointegration vectors for the maximum eigenvalue and trace statistics, respectively. Among five variables there is a possibility of zero, one, two, three, or four cointegrating vectors. As shown in Table 2, the trace test rejects the null hypothesis of zero cointegrating vectors ($H_0=0$) at the 5 percent significance level but fails to reject the null hypotheses $r \neq 1$, $r \neq 2$, $r \neq 3$, and $r \neq 4$. Similarly, the maximum eigenvalue test results also reject the null hypothesis of $r=0$ in favor of the alternative hypothesis of $r=1$. On the other hand, none of the alternative hypotheses $r=1$, $r=2$, $r=3$, and $r=4$ can be rejected in favor of the alternative hypotheses $r=2$, $r=3$, $r=4$, and $r=5$, respectively. Both the trace and eigenvalue tests indicate that there is a unique statistically significant cointegrating vector, i.e., one linear long-run equilibrium relationship among the variables, and any departure from this relationship may be due to temporary disequilibrating forces. The fact that the maximum eigenvalue test rejects the null hypothesis is noteworthy since Johansen and Juselius (1990) observe that one should expect the maximum eigenvalue test to produce more clear-cut results. The unique cointegrating vector, corresponding to the largest eigenvalue of the stochastic matrix, may be normalized with respect to i^{th} export price and interpreted as the long-run price response function for P_i .

The cointegrating vectors are normalized on each of the export prices. This is done by setting the coefficient of normalized price equal to -1 and dividing each cointegrating vector by the negative of the estimated coefficient of normalized variable. For example, the cointegrating vector normalized on U.S. price is

$$P_{us} = 1.51 + 0.19(P_{ca}) - 0.28(P_{au}) - 0.44(P_{eu}) - 0.18(P_{ar}) \quad (5)$$

After having confirmed that the price series are indeed first differenced stationary and cointegrated, we estimated the error correction model (equation 1). The residuals of the cointegrating equation for each price, which had been generated from the Johansen maximum likelihood procedure, are lagged one period and included as regressors in the error correction model (equation 1). The results from the ECM are summarized in Table 3. Four lagged differences of each variable are initially included in each regression equation. Starting with the longest lagged difference, lagged terms that are not statistically significant are eliminated gradually¹. A battery of diagnostic tests was performed to check the adequacy of the error correction model. These tests include the Lagrange Multiplier (LM) test (η_1) for autocorrelation in the residuals, LM test (η_2) for autoregressive conditional heteroskedasticity (ARCH) effect, i.e., autocorrelation in squared residuals, and the RESET test (η_3) for functional form misspecification. The test results are presented in the lower part of Table 3. The LM tests fail to reject the null hypotheses of autocorrelation in residuals and squared residuals. The Ramsey RESET test reveals no serious violation in the linearity assumption in the structure of the error correction model.

Considering that the dependent variable in each equation is at first differenced, the higher adjusted R square (0.41 to 0.75) in the five equations suggests that the statistical fit of the model is satisfactory. It is important to note that the error correction term in each of the export price equations is statistically significant and also has appropriate (negative) signs. A significant and negative error correction term validates the existence of an equilibrium relationship among the variables and suggests that ignoring the nonstationarity and cointegration of the variables would introduce misspecification in the underlying dynamic structure (Arize 1995). The size of this coefficient provides a measure of speed of adjustment towards long-run equilibrium. For example, in the U.S. export price equation, 11 percent of adjustment towards long-run equilibrium occurs within the first month. Similarly, lower values of the coefficient of error correction terms in Canadian, Australian, and Argentinean export price equations suggest that these prices are very slow in adjusting towards long-run equilibrium.

Final specifications for all the export price formations include at most two period lags of their own and competitors' prices. For each export price equation, parameter estimates are presented along with the t-statistics (in parentheses). In the U.S. export price formation equation, the Canadian and Australian export prices have statistically significant coefficients. This implies that the U.S. export price is affected

¹Elimination of statistically insignificant variables or lags from an estimating equation is very common practice in econometric literature. This provides a model that is parsimonious and easy to interpret. For more discussion on this issue see Turnovsky and Wohar (1987).

by changes in Canadian and Australian export prices, but movements in the European Union and Argentinean export prices seem to have no significant effect on the U.S. export price. Similarly, the Canadian export price equation suggests that it is affected by the Australian export price, but the prices of other competing exporters do not seem to have a significant effect in the formation of the Canadian export price. Only Australian, European Union, and Argentinean export prices appear to respond significantly to changes in U.S. and Canadian export prices but do not respond to each other's prices.

Unlike Canadian wheat exports, U.S. wheat exports are not coordinated through a single marketing board, but rather are driven by the actions of private commodity trading companies such as Cargill, Continental, and Louis-Dreyfus (Goodwin and Smith, 1995). Thus, in effect, the Canadian Wheat Board (CWB) is the largest single operator in world wheat markets. In that case, our results are consistent with the fact that the United States responds to the largest exporter in the international market—Canada. The follower role of the U.S. price to Canadian wheat prices may be arguable because quality differences exist among wheat classes sold by these exporting countries; for example, Canadian Hard Red Spring wheat sells at a premium compared to the Hard Red Winter wheat of the United States. Thus, it is logical to expect the United States to follow the price set by the CWB, which sells higher-quality wheat.

The finding that the United States responds to Canadian price changes, but Canada does not respond to U.S. price changes, supports the concerns of U.S. wheat producers and exporters who argued that the Export Enhancement Program (EEP) process does not provide a level playing field for U.S. wheat exporters vis-a-vis the Canadian Wheat Board. They contend that the EEP process instead allows U.S. wheat exporters to react to Canadian pricing decisions when these decisions clearly threaten U.S. wheat exports. With regard to the current wheat trade dispute between the United States and Canada, our results may also support the allegation by U.S. producers that Canadian production subsidies and implicit export subsidies have undermined the U.S. price support program, including EEP.

The CWB responds only to Australian prices. On the other hand, each of the other market prices responds significantly to Canadian prices. Bidirectional causality between Australian and Canadian prices may be a reflection of the fact that Canada and Australia are important competitors in some Asian wheat markets and thus respond to each other's price changes. Responsiveness of the Canadian price to the Australian price but not to the U.S. price may be due to the higher quality Canadian wheat (#1 Western Red Spring) being a better substitute for Australian (Standard) wheat than U.S. (Hard Winter #2) wheat.

Our results also suggest that the EU responds to U.S. price changes with no significant response in the opposite direction. This may be in contrast to the conventional belief that U.S. export subsidies are based on EU export prices and might mean that since 1985, at least, U.S. prices may have responded to

EU prices. We also repeated the above price analysis using the data from June 1985 and the U.S./EU price relationships were identical to the one discussed above. This suggests that the U.S. might have started the EEP because of EU export subsidies, but after the program was initiated the EU might be responding to EEP subsidies in determining its subsidy level. The similarity in results in both periods may imply that EEP has not been instrumental in altering U.S./EU wheat price relationships.

Our results do not coincide with past studies (Goodwin and Schroeder 1991; Goodwin and Smith 1995; Mohanty et al. 1995) that found either the United States or Canada to be the price leader in the international wheat market. Goodwin and Schroeder (1991) and Mohanty et al. (1995) affirmed the price leadership role of the United States. By contrast, Goodwin and Smith (1995) supported the price leadership role of Canada based on findings that Canada significantly affects other market prices but is not affected by other prices. Our results are somewhat similar to the findings of Goodwin and Smith (1995) except that we found a significant response of Canadian prices to changes in Australian prices, which weakens the hypothesis that Canada is a price leader. Overall, our results imply that there is no significant price leader in the international wheat market.

Conclusion

In this paper, price dynamics in the international wheat market have been examined. Time series properties of all the price series were examined and found to be first differenced stationary and cointegrated. Using an error correction model, it was found that there is no distinct price leader in the international wheat market. However, Canadian price significantly affects other prices including the U.S. price but is not affected by other prices except Australia's. Interestingly, our results suggest that the U.S. price affects other prices except Canada's but is affected by changes in Canadian and Australian prices. Similarly, while the three remaining prices (Australia, the European Union and Argentina) respond to both Canadian and U.S. price changes they do not seem to have significant interaction among themselves.

The finding that neither the United States nor Canada is the price leader in the international wheat market is not consistent with past studies (Goodwin and Schroeder 1991; Goodwin and Smith 1995; Mohanty et al. 1995) that showed Canada or the United States as the price leader. These past studies could be compared with our studies because of similarity in data and method. However, these studies failed to examine the price series for nonstationarity and cointegration and could be misspecified.

Table 1. Nonstationarity results using Augmented Dickey Fuller tests

Variable	ADF Test Statistics	
	Levels	First Difference
P _{us}	-2.61	-4.73*
P _{ca}	-2.72	-4.89*
P _{au}	-2.49	-4.49*
P _{eu}	-2.25	-6.07*
P _{ar}	-2.44	-4.25*

* Indicates rejection of null hypothesis of nonstationarity or unit roots at 5 percent significance level. For n=150, critical value calculated from McKinnon (1991) is -2.88 at 5 percent significance level.

Table 2. Maximum eigenvalue and trace test statistics on number of cointegrating vectors (r)

Null Hypothesis	Maximum Eigenvalue Statistic	95% Critical Value	Null Hypothesis	Trace Statistics	95% Critical Value
r = 0	40.96	34.40	r = 0	89.85	71.86
r = 1	25.61	28.24	r = 1	48.88	53.11
r = 2	11.22	22.01	r = 2	23.28	34.91
r = 3	6.78	15.67	r = 3	12.06	19.96
r = 4	5.28	9.24	r = 4	5.28	9.24

* Indicates rejection of null hypothesis at 5 percent significance level. Both tests have alternative hypotheses of H_a: r = n

Table 3. OLS estimates of exporter's price formation using error correction model

	Dependent Variable				
	ΔP_{us}	ΔP_{ca}	ΔP_{au}	ΔP_{eu}	ΔP_{ar}
ΔP_{us}			0.62(11.67)		0.24(1.96)
$\Delta P_{us}(-1)$				0.54(4.37)	
$\Delta P_{us}(-2)$	-0.10(-1.98)				
ΔP_{ca}	0.13(2.23)		0.24(4.74)		0.33(2.81)
$\Delta P_{ca}(-1)$		0.23(3.39)			
$\Delta P_{ca}(-2)$	0.11(2.62)			0.25(2.14)	
ΔP_{au}	0.79(11.91)	0.76(11.28)			
$\Delta P_{au}(-1)$	0.11(1.91)		-0.03(2.01)		
$\Delta P_{au}(-2)$					
ΔP_{eu}			0.03(-1.75)		
$\Delta P_{eu}(-1)$				-0.04(1.93)	
$\Delta P_{eu}(-2)$					
ΔP_{ar}					
$\Delta P_{ar}(-1)$					0.30(4.41)
$\Delta P_{ar}(-2)$					
$Z_{us}(-1)$	-0.11(-2.59)				
$Z_{ca}(-1)$		-0.09(-1.94)			
$Z_{au}(-1)$			-0.07(-2.52)		
$Z_{eu}(-1)$				-0.39(-8.15)	
$Z_{ar}(-1)$					-0.06(-3.41)

Table 3. (continued)

	Dependent Variable				
	ΔP_{us}	ΔP_{ca}	ΔP_{au}	ΔP_{eu}	ΔP_{ar}
ADJ. R^2	0.72	0.52	0.74	0.41	0.38
DW	2.06	1.92	2.18	2.06	2.03
$\eta_1(23)$	21.96	28.06	17.95	26.06	15.21
$\eta_2(1)$	0.92	0.12	1.09	2.21	1.41
$\eta_3(1, 144)$	2.31	1.61	1.80	0.19	2.06

Notes: Numbers in parentheses are the t-statistics.

In the first column, numbers in the parentheses are lag order.

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