2-1996

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Abstract

U.S. wheat producers contend that Canadian production subsidies and implicit export subsidies have undermined the U.S. price support program. The authors examine this contention and assess the relationships between U.S. and Canadian wheat prices using a cointegration and error approach. The results suggest that both U.S. durum and hard spring wheat prices respond to restore the equilibrium relationship with the corresponding Canadian price while the Canadian price does not respond to restore equilibria.

Disciplines

Agricultural and Resource Economics | Agricultural Economics | International Economics

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Working Paper 96-WP 154
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ABSTRACT

In this paper, relationships between U.S. and Canadian wheat prices are examined using cointegration and error correction approach. The use of an error correction model is appropriate because U.S. and Canadian wheat prices are first differenced stationary and cointegrated. The results suggest that both U.S. durum and hard spring wheat prices respond to restore the equilibrium relationship with the corresponding Canadian price while the Canadian price does not respond to restore equilibria. That is, the structure of the respective policies is such that Canadian markets are largely insulated from influences flowing from the United States while U.S. markets are not insulated from Canadian influences. These results could be interpreted to support the contention of U.S. producers that Canadian production subsidies and implicit export subsidies have undermined the U.S. price support program. These results also support the price leadership role for Canada in both the durum and hard spring wheat market.

TIME SERIES EVIDENCE OF RELATIONSHIP BETWEEN U.S. AND CANADIAN WHEAT PRICES

International wheat price relationships have received considerable attention in recent years. Market structure, price leadership, the efficiency of government interventions, and many other issues have been addressed within the context of price behavior in the international wheat market (Goodwin and Schroeder 1991). Most of these studies have used time series methods to analyze the price relationships of interest (Spriggs et al. 1982; Ardeni 1989; Goodwin et al.1990; Goodwin and Schroeder 1991; Goodwin 1992; and Mohanty et al. 1993;). Spriggs et al. (1982) examined the price leadership roles of the United States and Canada using standard Granger causality tests. Mohanty et al. (1993) used Granger causality tests to determine the exogeneity of U.S. prices in the export price formation of competing exporters. The results showed that causality runs from the United States to competing exporters, including Canada, with no causality in the reverse direction.

Goodwin and Schroeder (1991) examined the dynamics of price relationships in the international wheat market using a vector autoregressive (VAR) model. In addition to prices, they included two other economic variables, exchange rates and shipping rates. The forecast error variance decomposition indicated that the U.S. price has a significant effect on the Canadian price but the Canadian price does not influence the U.S. price.

Other researchers have approached the problem from the perspective of the law of one price using the concept of cointegration (Ardeni 1989; Goodwin 1992). Ardeni (1989) used bivariate two-step cointegration testing techniques developed by Engle and Granger (1987) and found that the law of one price did not appear to hold, even in the long run, for most commodity price series including wheat. He did find that some of the wheat price series were cointegrated but argued that the overall results do not provide support for the notion that the law of one price holds for primary commodities. Goodwin (1992), on the other hand, used multivariate cointegration tests, recently introduced by Johansen (1988) and Johansen and Juselius (1990), and found that wheat price series are cointegrated.

An important point frequently overlooked is that if, in fact, prices are cointegrated, standard Granger causality tests provide misleading results (Miller and Russek 1990). In addition, the structural VAR approach used by Goodwin (1992) is likely to be misspecified, raising questions about the validity of the results. Both Granger causality and the VAR model focus on short-run dynamics rather than long-run. Although some analysts (Ardeni 1987, Goodwin 1992) have examined the existence of long-run relationships between or among wheat prices, they did not examine the mechanisms through which long-run relationships are restored.

Another shortcoming of most studies evaluating price relationships in the international wheat market is that not enough attention has been given to differentiation of wheat. Ardeni (1987) and Mohanty et al. (1993) used the U.S. and Canadian prices for hard red winter and western red spring wheat. Goodwin (1992), in his multivariate cointegration analysis, has pointed this out and chose to use the U.S. export price for dark northern spring, which may be more comparable to Canadian red spring wheat.

This paper explores the relationships between Canadian and U.S. wheat prices, in particular the prices of hard red spring and durum wheat using cointegration and error correction techniques. These two classes of wheat have been at the heart of the U.S.-Canadian trade dispute and are the commodities over which there is the greatest competition between Canada and the United States. The use of the error correction model allows the rigorous study of long-run and short-run price relationships simultaneously. The short-run elements describe the dynamics of moving towards a new equilibrium. The long-run relationships are of special interest because they offer insight into how equilibrium relationships are restored and what new equilibrium levels would be obtained with a policy shift.

The results of the analysis of these price relationships provide information on the existence of price leadership in the world market for durum and hard spring wheat. An understanding of price leadership is important in explaining the structure of the market and also helps researchers in correctly specifying price linkage equations. For example, if Canada is the price leader, then results of the studies that model the United States as the price leader may not be correct. More important, the long-run price relationship is an important piece of information for policymakers in formulating domestic policies and understanding the recent wheat trade dispute.[1]

Error Correction Model

Recently, Granger (1983, 1986) and Engle and Granger (1987) provided a comprehensive test of causality in the presence of cointegrated variables. This test specifically allows for a causal linkage between two variables stemming from a common trend or equilibrium relationship (Miller and Russek 1990). This alternative test of causality incorporates the possibility that the lagged level of causal variables may also explain current changes in the dependent variable. The reason for including this lagged term in the causality test is that if two variables, X and Y, have a common trend or are cointegrated, current changes in X are partly the result of moving into alignment with the trend value of Y (Miller and Russek 1990). Standard causality tests miss these effects because they only include the effect of previous changes in causal variable Y on current changes in X.

The alternate test of causality developed by Granger (1983, 1986) and Engle and Granger (1987) is based on an error correction representation. The error correction equation used for testing causality between cointegrated variables is

$$\Delta X_t = \alpha_0 + \sum_{s=1}^{k} b_s \Delta X_{t-s} + \sum_{s=1}^{k} c_s \Delta Y_{t-s} + \gamma_1 \eta_{t-1} + \varepsilon_t \, , \qquad (1)$$

where η_{t-1} is the lagged error term from the following cointegrating equation,

$$X_t = \alpha_0 + \delta y_t + \eta_t \, . \qquad (2)$$

The cointegrating equation defines long-run equilibrium relations in X and Y. Residuals from the cointegrating equations (error correction term), representing departure from the long-run equilibrium, are included in the error correction model to capture the response of X to any

[1] Recent growth in U.S. imports of Canadian wheat have created growing tension between the two countries and caused the United States to attempt to block wheat imports from Canada. U.S. efforts to block wheat imports from Canada are not only inconsistent with the GATT but also with the spirit of NAFTA. Representatives from both countries worked diligently to resolve the dispute but were unable to reach a solution because of disagreement on the causes of the increased wheat trade. Each country holds the farm programs of the other country responsible for the increased trade.

disequilibrium created by the movement in Y. The size and statistical significance of the coefficient on the error correction term in each ECM measures the tendencies of each price to return to equilibrium (Baghestani and McNown 1992). For example, if γ in equation 1 is statistically significant, it could be concluded that X responds to disequilibria in the X-Y relationship. Joint significance of the coefficients of lagged changes in Y measures the short-run or flitting response of X, as with standard causality tests. Additionally, the response of Y in adjusting towards a new equilibrium and also the short-run response of X through lagged changes in Y may be measured by repeating this experiment with Y and X interchanged.

Of importance for this study are the estimated dynamic responses to long-run disequilibria. If we find evidence that the U.S. price responds to disequilibria induced by a shock shifting either U.S. or Canadian price levels, but that the Canadian price does not respond, it may be concluded that independent U.S. farm programs aimed at influencing market prices may be undermined by the Canadian response. For example, supply management policies in the United States such as set-a-sides and export subsidies, which have the goal of raising domestic prices above world levels, may, in the end, not be effective in the presence of Canadian subsidies for production and their implicit export subsidy policy. Similarly, if we find evidence that Canadian prices respond to the disequilibria but the U.S. price does not, efforts on the part of Canadian authorities to develop different farm programs from those of the United States may not be effective. Finally, changes in both U.S. and Canadian prices to restore the equilibrium indicate that differences in farm programs between the United States and Canada may prevent complete effectiveness of either program.

Relative to the current dispute, the discovery that U.S. price adjusts to restore equilibrium but Canadian price does not could be interpreted to mean that U.S. programs are indeed undermined by Canadian policies designed to lower the price. But if the reverse holds –Canadian price responds to disequilibria but U.S. price does not–then it might be that U.S. policies such as Export Enhancement Programs (EEP) designed to raise domestic prices will cause the Canadian price to increase in order to restore the equilibrium relationship. In such a case, U.S. producers' claims that Canadian production subsidies and implicit export subsidies have undermined the U.S. price support program would be unfounded. Finally, responses by both prices to disequilibria suggest that the claims by both countries are not completely true.

Data and Estimation

The data used for this analysis are monthly quoted f.o.b. prices for January 1978 to June 1993. The specific price series for the U.S. include #2 dark northern spring 14 percent protein, Gulf (P_{uss}) and #3 hard amber durum, lake (P_{usd}). The corresponding Canadian prices are western red spring, 13.5% protein, Pacific(P_{cas}) and #1 amber durum, Vancouver(P_{cad}). The primary sources for the F.o.b. prices were the 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. The series is divided into two subseries, the period before implementation of EEP and the period after. The pre-EEP subseries contains observations from January 1978 to December 1985, while the post-EEP subseries is from January 1986 to June 1993. The data are studied in two subperiods according to the suggestion of Spriggs et al. (1982), who argue for the importance of splitting the time series in cases where historic policy shifts have occurred.

Cointegration is a necessary and sufficient condition for representing the series in an ECM (Engle and Granger 1987). Since previous studies provide conflicting evidence on the existence of cointegration in the international wheat market, this study tests for cointegration using both Engle and Granger's bivariate cointegration technique and a maximum likelihood procedure developed by Johansen (1988) and Johansen and Julius (1990).

The empirical test of cointegration must be preceded by a test of nonstationarity for the individual variables under consideration to determine the order of integration of each variable. The determination of order of integration of each variable is required for cointegration and of more importance, for error correction equations because each variable involved in the estimation of these models must be first-difference stationary. To verify that the first-differenced price series are indeed stationary, both Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) unit root tests were used. The (ADF) test is based on the regression

$$\Delta X_t = \alpha_0 + \beta X_{t-1} + \sum_{i=1}^{m} \delta_i \Delta X_{t-i} + \epsilon_t \, , \tag{3}$$

where Δ is the first difference operator and ϵ_t is a stationary error term. The number of lags to include in the equation was determined using the Aikaike information criterion and was found to be four for all the price series. The DF test is based on this regression without the right hand side summation. The importance of including a constant term without a time trend has been addressed by Dickey, Bell, and Miller (1986) and Miller and Russek (1990). Based on their suggestions, both DF and ADF equations were estimated with an intercept and no time trend. All the price series are expressed in logarithmic form.

Table 1 presents the results of both Dickey-Fuller and augmented Dickey-Fuller unit root tests for each price series. The null hypothesis of nonstationarity was tested using a t-test on the β_1 coefficient in both cases. The null hypothesis is rejected if β_1 is significantly negative. Based on the critical values reported by Fuller (1976), nonstationarity cannot be rejected for the level of all price series at the 5 percent significance level but nonstationarity was rejected for all the price series expressed in first differences at the same significance level for both the pre- and post-EEP periods. Because determination of lag order using statistical tests alone has been criticized, both tests were conducted using different lag orders. These alternative representations did not alter the results of the tests.

Having confirmed that the price series are stationary in first differences, we can proceed with the cointegration tests. Cointegration is tested using both Johansen's maximum likelihood procedure and the bivariate cointegration technique of Engle and Granger. Johansen's maximum likelihood procedure is based on the error correction representation

$$\Delta Z_t = \sum_{j=1}^{k} \alpha_j \Delta Z_{t-j} + \theta(r) Z_{t-1} + \epsilon \ , \tag{4}$$

where Z_t is the 2×1 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 statistics. The critical values for these statistics were obtained from Johansen and Juselius (1992). Equation 4 was estimated with U.S. and Canadian durum prices and then reestimated using spring wheat prices. Four lags for the prices of both classes of wheat were included. Estimations were performed both with and without a time trend. As reported in Table 2, both the eigenvalue and trace tests reject the null hypothesis of no cointegrating vectors at the 5 percent significance level but fail to reject the null hypothesis of one or fewer cointegrating vectors with or without the time trend in the model. Thus, Johansen's tests support the existence of cointegration between U.S. and Canadian durum prices, and U.S. and Canadian spring wheat prices for both the pre- and post-EEP period.

The bivariate cointegration test developed by Engle and Granger (1987) is based on the stationarity of the residuals of the cointegration equation. X and Y are cointegrated if the residuals from regressing X on Y and Y on X are both stationary. Cointegration for P_{uss} and P_{cas} was tested by regressing P_{uss} on P_{cas} and P_{cas} on P_{uss} and testing residuals from these regressions for unit roots. The procedure was repeated for P_{usd} and P_{cad}. Table 3 presents the results. Results of both DF and ADF test statistics on residuals are presented in the last two columns of Table 2. Using both DF and ADF tests, the nonstationarity of residuals is rejected at the 5 percent significance level in most cases. If a significance level of 10 percent is used, the null hypothesis of nonstationarity is rejected in all cases. It is reasonable to conclude that both pairs of U.S. and Canadian prices are cointegrated. Thus, a long-run ECM equilibrium relationship between these price series also exists and one or both parties to the dispute may have evidence to support its claims.

Finally, the ECM was estimated for each pair of cointegrated price series for both the pre- and post-EEP periods. Residuals of the cointegration equations were lagged and used as the error correction terms in the ECM. For example, the lagged residuals($\eta_{PussPcas}$) from regressing P_{uss} on P_{cas} were included as the error correction term in the error correction representation having ΔP_{uss} as the dependent variable. The number of lags included in the ECM was the same as in the tests for the unit roots and cointegration. Validation of the ECM estimates are obtained by examining the Box-Pierce Portmanteau Q-statistic associated with the fitted residuals. Q-statistics for residuals associated with each ECM explaining price relationships between U.S. and Canadian spring wheat prices for both pre and post-EEP are reported in Table 4. Similarly, Q-statistics for residuals associated with each ECM explaining price relationships between U.S. and Canadian durum prices are reported in Table 5. In all cases, Q statistics (8 and 12 lags) are smaller than their respective critical values. This suggests that the null hypothesis of no autocorrelation or the white noise status of residuals of the ECMs cannot be rejected.

OLS estimates of the ECMs for hard spring and durum price relationships are reported in Table 4 and 5. The most important finding is the statistical significance of the error correction or disequilibrium term ($\eta_{PussPcas}$, $\eta_{PusdPcad}$) in both the U.S. durum and spring wheat equations for both the pre- and post-EEP period. This suggests that U.S. wheat prices adjust to correct long-run disequilibria in U.S. and Canadian prices. On the other hand, the statistical insignificance of the disequilibrium term in both Canadian durum and spring wheat equations suggests that Canadian prices do not respond to long-run disequilibria between U.S and Canadian prices.

These results indicate that the U.S. market is not insulated from the Canadian market but that the reverse does not hold. Thus, any U.S. policy designed to alter wheat prices directly or indirectly without taking account of the Canadian market situation is likely to be ineffective. For example, both the United States and Canada have policies designed to affect their domestic prices but the expected effects of these policies run in opposite directions. Canadian input subsidies should lower the domestic wheat prices while EEP would be expected to raise U.S. prices. The Canadian policies could be effective tools because their prices do not respond to disequilibria, policy-induced not. Supply management policies in the United States, however, may not be effective because higher U.S. prices relative to equilibrium values in time t trigger responses that lead to a fall in U.S. prices in the following periods. The similarity of the long-run relationships for both the pre- and post-EEP periods suggest that EEP has not been an important factor in altering the long-run equilibrium relationship between U.S. and Canadian prices. From the viewpoint of the current wheat trade disputes between the United States and Canada, these results support the allegations of U.S. producers.

Of secondary interest are the causality results implied by the lagged difference terms. The short-run dynamics are characterized by unidirectional causation, with Canadian durum price significantly affected by lagged changes in U.S. durum price, whereas the U.S. price is unaffected by lagged changes in Canadian durum price during the pre-EEP period. Apart from this, there seems to be no significant short-run causality between U.S. and Canadian durum prices in either direction for the post-EEP period. In the case of spring wheat prices, there is no significant short-run causation between U.S. and Canada both for the pre- and post-EEP periods.

The fact that U.S. prices respond to restore the equilibrium relationship with the Canadian price while the Canadian price does not respond to restore equilibria would suggest that Canada is the price leader both in the durum and spring wheat markets. This is in contrast to the belief that the United States is the price leader (Perkins et al. 1984; Bailey 1987; Thursby and Thursby 1989).

Conclusions

The long-run relationship between Canadian and U.S. durum and spring wheat markets is uni-directional. That is, the structure of the respective policies is such that Canadian markets are largely insulated from influences from the United States while U.S. markets are not insulated from Canadian influences. These results could be interpreted to support the contention of U.S. wheat producers that Canadian production subsidies and the implicit export subsidy have undermined U.S. price support programs. However, without further investigation, we cannot say that wheat imports from Canada are directly responsible for price changes in the United States because no account has been taken of the high quality and standardization of Canadian wheat.

On the other hand, this interpretation does not necessarily mean that the Canadian government has implemented predatory policies targeted at the U.S. market. The likely effect of an export subsidy program like the EEP is to drive a wedge between internal prices, which are raised, and world prices, which are depressed if the subsidizing country is a major exporter of the subsidized good. The Canada-U.S. trade agreement of 1989 and the recently-enacted NAFTA are designed to eliminate trade barriers between the United States. and Canada. With no trade barriers, we would expect Canadian wheat to be diverted from the depressed world market to U.S. markets where favorable prices, induced by the EEP, prevail. This study shows that, with the current policy structure in Canada and the United States, efforts by the United States to pursue independent durum and spring wheat price policy objectives may well be undermined. In addition to policy implications, these results also support the price leadership role of Canada in both the durum and spring wheat markets.

Table 1. Nonstationarity results using Dickey-Fuller and Augmented Dickey-Fuller tests

	DF Test Statistics		ADF Test Statistics	
Variables	Levels	1st Difference	Levels	1st Difference
		1978:01 - 1985:12		
P_{uss}	-2.42	-7.38*	-2.26	-5.37*
P_{cas}	-2.81	-7.69*	-2.14	-4.82*
P_{usd}	-1.66	-5.76*	-2.09	-3.97*
P_{cad}	-1.74	-5.90*	-2.23	-4.43*
		1986:01 - 1993:06		
P_{uss}	-1.32	-5.28*	-1.94	-3.89*
P_{cas}	-1.07	-4.86*	-1.59	-4.17*
P_{usd}	-2.08	-6.44*	2.40	-4.42*
P_{cad}	-1.56	-5.03*	1.44	-5.21*

*Indicates rejection of null hypothesis of nonstationarity or unit root at 5 percent significance level; critical values for n=100 is -2.89 and n=50 is -3.58 percent at 5 percent significance level (Fuller).

Notes: P_{uss}= U.S. #2 dark northern spring 14% protein, Gulf price; P_{cas}= #1 Canadian western red spring 13.5% protein, Pacific port price; P_{usd}= U.S. #3 hard amber durum, f.o.b. Lakes; and P_{cad}= Canadian #1 CW amber durum, f.o.b. Vancouver.

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

Test	Variables	Null hypothesis	Cointegration test statistics	Critical value at 5 percent
		1978:01 - 1985:12		
Trace test	P_{uss} & P_{cas}	$r = 0$	32.89*	17.95
		$r \leq 1$	5.84	8.17
Maximum eigenvalue test	P_{usd} & P_{cad}	$r = 0$	27.06*	14.90
		$r \leq 1$	5.84	8.18
		1986:01 - 1993:06		
Trace test	P_{uss} & P_{cas}	$r = 0$	19.79*	17.95
		$r \leq 1$	4.02	8.18
Maximum eigenvalue test	P_{usd} & P_{cad}	$r = 0$	15.77*	14.90
		$r \leq 1$	4.03	8.18

*Indicates rejection of the null hypothesis at the 5 percent significance level.

Note: Both tests have alternative hypotheses of H_a:r>n.

Table 3. Cointegration results

Dependent Variable	Const.	Coefficient of P_{cas}	P_{uss}	P_{cad}	P_{usd}	DF	ADF
				1978:01 - 1985:12			
P_{uss}	.57	.872	-	-	-	-3.93^{*}	-4.34^{*}
P_{cas}	.92	-	.514	-	-	-4.01^{*}	-4.10^{*}
P_{usd}	.065	-	-	.984	-	-3.83^{*}	-3.82^{*}
P_{cad}	.421	-	-	-	.946	-3.40^{**}	-3.56^{*}
				1986:01 - 1993:06			
P_{uss}	.007	.958	-	-	-	-3.10^{*}	-3.48^{*}
P_{cas}	.414	-	.959	-	-	-2.83^{**}	-2.95^{**}
P_{usd}	1.010	-	-	.762	-	-4.14^{*}	-3.29^{*}
P_{cad}	-.312	-	-	-	1.106	-3.71^{*}	-2.96^{**}

* and ** indicates rejection of null hypothesis of nonstationarity of residuals at 5 and 10 percent significance levels.

Note: For ADF test critical values for n=100 are -3.17 and -2.91 for 5 percent and 10 percent significance level and for DF test the corresponding significance values are -3.05 and -2.71, respectively (Engle and Granger, and Fuller).

Table 4. OLS estimates for U.S. and Canadian spring wheat prices using ECM for both pre- and post-EEP period

	Pre-EEP Period		Post-EEP Period	
Dependent Variable	ΔP_{uss}	P_{cas}	ΔP_{uss}	ΔP_{cas}
	(1978:1-1985:12)		(1986:1-1993:06)	
$\Delta P_{uss}(-1)$	0.49(3.63)	-0.25(0.26)	-0.88(0.38)	0.25(1.2)
$\Delta P_{uss}(-2)$	0.19(1.5)	-0.84(0.89)	0.31(1.37)	0.15(0.75)
$\Delta P_{uss}(-3)$	-0.04(0.33)	-0.58(0.63)	0.23(0.1)	0.01(0.06)
$\Delta P_{uss}(-4)$	0.21(1.65)	-0.2(0.21)	-1.0(0.43)	-0.58(0.28)
$\Delta P_{cas}(-1)$	-0.13(0.73)	0.29(2.14)	0.65(2.36)	0.77(3.19)
$\Delta P_{cas}(-2)$	-0.18(1.08)	0.03(0.21)	-0.26(0.92)	-0.99(0.39)
$\Delta P_{cas}(-3)$	0.19(1.13)	0.16(1.27)	-0.29(1.1)	-0.27(1.15)
$\Delta P_{cas}(-4)$	-0.29(1.72)	-0.15(1.23)	0.28(1.04)	0.24(0.96)
$\eta_{PussPcas}(-1)$	-0.51(3.95)		-0.32(2.39)	
$\eta_{PcasPuss}(-1)$		-0.14(1.53)		0.07(0.54)
Q-Statistics				
Q(8)	3.43	2.54	1.12	2.72
Q(12)	6.93	8.07	6.74	6.5
F-Statistics	1.29	0.3	2.18	0.57

Notes: Numbers in parentheses are the absolute values of t-ratios.

$\eta_{PussPcas}$ and $\eta_{PcasPuss}$ are the residuals series from the OLS cointegrating regressions (P_{uss} on P_{cas} and P_{cas} on P_{uss} respectively) reported in Table III.

The Q-Statistic denotes Box-Pierce-Ljung Portmanteau tests for autocorrelation, which are distributed as chi squares with degree of freedom equal to the lags provided within the parenheses. The critical values at the 5 percent significance level are 15.51 and 21.03 at 8 and 12 degree of freedom.

F-Statistics measure the joint significance of c_s (s=1,...,k). Critical value for F-statistic is 2.49 at the 5 percent significance level.

Table 5. OLS estimates for U.S. and Canadian durum wheat prices using ECM for both pre- and post EEP period

	Pre-EEP Period		Post-EEP Period	
Dependent Variable	ΔP_{usd}	ΔP_{cad}	ΔP_{usd}	ΔP_{cad}
	(1978:1-1985:12)		1986:1-1993:06)	
$\Delta P_{usd}(-1)$	0.69(3.8)	0.38(3.04)	0.37(1.78)	0.19(1.26)
$\Delta P_{usd}(-2)$	0.12(0.67)	0.01(0.04)	-0.11(0.54)	0.03(0.17)
$\Delta P_{usd}(-3)$	-0.04(0.23)	-0.15(1.23)	0.23(1.25)	0.19(1.4)
$\Delta P_{usd}(-4)$	0.3(1.73)	0.21(1.73)	-0.2(1.1)	-0.22(1.6)
$\Delta P_{cad}(-1)$	-0.15(0.007)	0.17(1.06)	0.38(1.53)	0.4(2.1)
$\Delta P_{cad}(-2)$	-0.25(1.15)	-0.22(1.43)	-0.27(1.14)	-0.15(0.8)
$\Delta P_{cad}(-3)$	-0.03(0.14)	0.22(1.49)	-0.93(0.41)	-0.31(1.8)
$\Delta P_{cad}(-4)$	-0.15(0.76)	-0.3(2.21)	0.16(0.79)	0.27(1.66)
$\eta_{PusdPcad}(-1)$	-0.42(3.59)		-0.52(2.67)	
$\eta_{PcadPusd}(-1)$		-0.15(1.31)		-0.02(0.2)
Q-Statistics				
Q(8)	3.08	5.74	4.17	5.6
Q(12)	9.77	16.75	6.08	8.87
F-Statistics	0.52	4.3	1.38	1.99

Notes:Numbers in parentheses are the absolute values of t-ratios.

$\eta_{PusdPcad}$ and $\eta_{PcadPusd}$ are the residuals series from the OLS cointegrating regressions (P_{usd} on P_{cad} and P_{cad} on P_{usd} respectively) reported in Table 3.

The Q-Statistic denote Box-Pierce-Ljung Portmanteau tests for autocorrelation, which are distributed as chi squares with degree of freedom equal to the lags provided within the parenheses. The critical values at the 5 percent significance level are 15.51 and 21.03 at 8 and 12 degree of freedom.

F-Statistics measure the joint significance of c_s (s=1,...,k). Critical value for F-statistic is 2.49 at the 5 percent significance level.

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