Price Dynamics in the North American Wheat Market

Jungho Baek and Won W. Koo

Perron's test, Johansen cointegration analysis, and a vector error-correction (VEC) model are used to identify structural change, as well as to examine price dynamics in the U.S. and Canadian hard red spring (HRS) and durum wheat markets. It is found that, due to the U.S. Export Enhancement Program (EEP), price instability experienced in June 1986 has resulted in structural changes for Canadian HRS and durum prices. We also find that Canadian prices have significant effects on the determination of the U.S. prices in the North American wheat market.

Key Words: Canadian wheat exports, durum wheat, hard red spring wheat, Johansen cointegration test, unit root test with a structural break, vector error-correction

The United States and Canada are the leading wheat exporters in the world market (Koo and Taylor 2005). The United States leads in exports of hard red winter (HRW) and soft red winter (SRW) wheat: an annual average of 10.6 million metric tones (MMT) of HRW and 6.8 MMT of SRW during the 2000–2004 period. Canada is the leader in exports of hard red spring (HRS) and durum wheat: an annual average of 11.6 MMT of HRS and 3.1 MMT of durum during the 2000–2004 period. Furthermore, the United States exports HRS and durum wheat and competes with Canada in the world market. As such, HRS and durum wheat have been at the core of the U.S.-Canada wheat trade dispute.[1]

An understanding of price relationships in the North American wheat market is important in addressing market structure and price leadership, as well as in constructing correct models for price analysis (Goodwin and Schroeder 1991). For example, if we find evidence that, with a shock to the North American wheat market, the U.S. price tends to recover to the long-run equilibrium relationship with the Canadian price, but that the Canadian price does not adjust, it suggests that Canada plays a key role in price-setting in the North American market. If U.S. and Canadian wheat prices are cointegrated, on the other hand, it suggests that these two prices drift in a similar fashion in the long run, and the cointegration relationships should be included in modeling the North American wheat market; otherwise, the econometric models could give a biased estimation. More important, it is crucial to assess the price behavior to understand the ongoing wheat dispute between the United States and Canada. For example, the discovery of Canada's dominant role in setting price implies that the U.S. market is influenced by the Canadian market, but that the reverse does not hold. This indicates that the Canadian wheat trading practices may have an impact on price changes in the U.S. market. Hence, the finding can be interpreted to support the U.S. claim that the export practices of the Canadian

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[1] Since implementation of the Canada-U.S. Free Trade Agreement (CUSTA) in 1989, a number of wheat trade disputes have arisen between the two countries (Mattson and Koo 2005). The very latest wheat dispute between the United States and Canada has come as a result of a petition from the North Dakota Wheat Commission, the Durum Growers Trade Action Committee, and the U.S. Durum Growers Association. In September 2002, U.S. producers filed countervailing and antidumping petitions with the Department of Commerce (DOC) and the International Trade Commission (ITC). The petitions claimed that subsidized Canadian HRS and durum wheat were being dumped on the U.S. markets, depressing the U.S. prices and net farm income. The DOC then conducted investigations and the ITC issued its finding that the HRS wheat industry in the United States is materially injured by imports of Canadian HRS wheat, but that the durum industry is not materially injured or threatened by Canadian exports of durum wheat. As a result, in October 2003, antidumping (8.87 percent) and countervailing (5.29 percent) duties were imposed on imports of Canadian HRS wheat.

Agricultural and Resource Economics Review 35/2 (October 2006) 265–275

Wheat Board (CWB) to the U.S. market are unfair and depress the U.S. prices.[2]

Several studies have analyzed price relationships in the international wheat market (i.e., U.S. and Canadian markets) (Spriggs, Kaylen, and Bessler 1982, Gilmour and Fawcett 1987, Goodwin and Schroeder 1991, Goodwin 1992, Goodwin and Smith 1995, Gardner 1999). For example, Spriggs, Kaylen, and Bessler (1982) use the Granger causality test to examine the existence of price leadership between the U.S. and Canadian wheat prices. They find that there is no significant price leadership role between the two countries. Goodwin (1992) adopts the Johansen cointegration test to evaluate the law of one price (LOP) in international wheat markets. He finds evidence of LOP in the world wheat market. These studies have typically concentrated on either the short-run price relationships (e.g., Granger causality) or the long-run price relationship (e.g., cointegration). Relatively limited efforts have been made to estimate the short- and long-run price relationship simultaneously.

To our knowledge, Mohanty, Peterson, and Smith (1996) and Mohanty and Langley (2003) are the only studies completed so far to conduct a simultaneous analysis of the short- and long-run price relationship in the North American wheat market. For example, Mohanty, Peterson, and Smith (1996) employed the cointegration and error correction approach to estimate long-run and short-run wheat price relationships simultaneously. They found that there is no significant short-run causality between U.S. and Canadian wheat prices, and that Canada is the price leader in the long run. However, the two studies did not directly examine the structural change in the U.S. and Canadian price series.

Structural change is an important issue in time-series analysis and affects all the inferential procedures associated with unit roots and cointegration tests (Maddala and Kim 1998). More specifically, testing for the presence of unit roots in time-series data is a prerequisite for estimating an appropriate vector auto-regression (VAR) model. Given the assumption that the deterministic trend is correctly specified, the standard augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1981) is unable to detect structural breaks in the series. In other words, if there is a break in the deterministic trend, then the ADF test could falsely lead to the conclusion that there is a unit root when in fact there is not (Perron 1989). Moreover, cointegration analysis that fails to account for structural change may raise the issue of spurious long-run relationships (Harris and Sollis 2003). It is thus both desirable and necessary to perform tests for structural changes of the series to overcome the weaknesses of the ADF procedure, as well as to obtain a reliable estimation of structural relationships in the price series.

The objective of this study is to assess the dynamics of price relationships in the U.S. and Canadian HRS and durum wheat markets. For this purpose, we use the Johansen multivariate cointegration test and a vector error-correction (VEC) model to examine both short- and long-run price relationships in the North American wheat market. Unlike previous studies (Mohanty, Peterson, and Smith 1996, Mohanty and Langley 2003), we incorporate market structural break(s) in our testing, which could have a substantial impact on estimated results but has been largely ignored by previous studies of agricultural products markets. This analysis will enhance the understanding of price dynamics in the North American wheat market and contribute to the literature on the trade dispute.

The remainder of the paper is organized as follows. First, we briefly discuss the empirical methods of unit root testing in the presence of a structural break and of cointegration analysis. Next, we present our data and the results of unit root tests under structural change. Then, we report the main empirical results of our short- and long-run price analysis. Finally, we summarize principal findings and draw some concluding remarks.

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[2] The CWB, established by the Canadian Wheat Board Act of 1935, is a state trading enterprise (STE) responsible for the marketing of all western Canadian wheat, durum, and barley. The main purpose of the CWB is to provide the highest possible profits for western Canadian grain farmers. The practices of the CWB have become a crucial issue in U.S.-Canada trade disputes. A number of studies thus have examined the behavior and the effectiveness of the CWB. Koo et al. (2004) provide a comprehensive review of the economic literature on these issues. Among the literature, Lavoie (2002) and Wilson, Johnson, and Dahl (1999), for example, examine the ability of the CWB to price discriminate in wheat export markets. The results indicate that there is evidence of the CWB's price discrimination and market power. In addition, after the recent investigation of countervailing and antidumping petitions on Canadian wheat, the U.S. Trade Representative (USTR) and the ITC concluded that the CWB has exerted special monopoly rights and privileges, resulting in unfair trade advantages to Canadian producers.

Empirical Methods

Perron (1989) develops a modified ADF test for the presence of a unit root with three different models. If a structural break in a time-series is known, then it is possible to adjust the ADF test by including dummy variables that allow a one-time change in the structure occurring at a time T_B (time of break). The three alternative models are as follows:

Model A:

$$(1) \qquad y_t = \mu_0 + \beta t + \mu_1 DU_t + \varepsilon_t ,$$

Model B:

$$(2) \qquad y_t = \mu + \beta_0 t + \delta_1 DT_t^* + \varepsilon_t ,$$

Model C:

$$(3) \qquad y_t = \mu_0 + \mu_1 DU_t + \beta_0 t + \delta_1 DT_t + \varepsilon_t ,$$

where $DU_t = 1$ if $t > T_B$, and 0 otherwise; $DT_t^* = t - T_B$, if $t > T_B$, and 0 otherwise; and $DT_t = t$ if $t > T_B$, and 0 otherwise. Model A allows for a one-time change in the intercept of the trend function (crash model). Model B allows a change in the slope of the trend function without any sudden change in the intercept (changing growth model). Finally, model C allows both effects (slope and intercept) to take place simultaneously.

The Johansen cointegration procedure is used to identify the number of cointegration relationships among time-series data (Johansen and Juselius 1990, Johansen 1995). Following Johansen, the cointegrating VAR model can be defined as follows:

$$(4) \qquad \Delta X_t = \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + u_t ,$$

where X_t is an $(n \times 1)$ vector of endogenous variables (i.e., $X_t = [USH_t, CAH_t, USW_t]$ for HRS wheat price model and $X_t = [USD_t, CAD_t]$ for durum wheat price model); Δ is the difference operator; $\Gamma_1,...,\Gamma_{k-1}$ are the coefficient matrices of short-term dynamics; $\Pi = -(I - \Pi_1 + ... + \Pi_k)$ is the matrix of long-run coefficients; μ is a vector of constant; k is the lag length; and u_t is a vector of normally and independently distributed error terms. Granger's representation theorem states that if the coefficient matrix Π has reduced rank [i.e., there are $r \le (n - 1)$ cointegration vectors present], Π can be decomposed into a matrix of weights, α, and a matrix of cointegration vector, β, that is $\Pi = \alpha\beta'$. The matrix α represents the speed of adjustment to equilibrium and β' is a matrix of long-run coefficients such that the term $\beta' X_{t-k}$ represents up to $(n - 1)$ cointegration relationships in the system. The number of cointegration vectors, the rank of Π, in the model is estimated by the likelihood ratio test (Johansen 1995).

If all variables in X_t are cointegrated, a vector error-correction (VEC) model can be estimated to capture the short-run dynamics while restricting the long-run behavior of variables to converge to their cointegration relationships (Engle and Granger 1987). For this purpose, equation (4) can be reformulated as a short-run dynamic model as follows:

$$(5) \qquad \Delta X_t = \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \alpha(\beta' X_{t-1}) + \mu + u_t ,$$

where $\beta' X_{t-k}$ is a measure of the error or deviation from the equilibrium, which is obtained from residuals from the cointegrating vectors. It should be noted that equation (5) incorporates both short- and long-run effects, since variables are cointegrated. As such, if the long-run equilibrium holds, then $\beta' X_{t-1}$ is equal to zero. In contrast, during periods of disequilibrium, $\beta' X_{t-1}$ is non-zero and measures the distance the system has deviated from equilibrium during time t. An estimate of α thus provides information on the speed of adjustment and implies how the variable X_t changes in response to disequilibrium.

Data and Tests for Unit Roots under Structural Change

Data

Monthly FOB prices for U.S. and Canadian durum and HRS wheat are collected for the period of July 1979 to June 2002. The U.S. price series are No. 2 Dark Northern Spring (14 percent protein, USH_t) and No. 2 Hard Winter (13 percent protein, USW_t) in the Pacific market, and No. 2 Hard Amber Durum wheat (USD_t) in the Lake market. The corresponding Canadian price series are No. 1 Canadian Western Red Spring wheat (13.5 percent protein, CAH_t) in the Pacific market, and No. 1 Amber Durum (CAD_t) in the St. Lawrence market. Wheat prices from July 1979 to June 1989 are collected from various issues of *World Wheat Statistics*, published by the International Wheat Council. Prices for the period of July 1989 to June 2002 are obtained from various issues of *World Grain Statistics*, published by the International Grains Council. To allow for exchange rate fluctuations, Canadian prices are expressed in U.S. dollars. Hence, all price series are quoted in nominal U.S. dollars per ton. Finally, all series are in logarithmic form.

One variable that merits elaboration is hard red winter (HRW) wheat. The dominant wheat class for the U.S. exports is HRW wheat, while for Canada it is HRS wheat. In addition, HRW wheat is known as a close substitute for HRS wheat and thus could have a significant effect on HRS wheat price (Gilmour and Fawcett 1987, Koo and Mattson 2002). However, studies have paid little attention to this relationship in their dealings with the North American wheat market. As such, we believe that it is worthwhile to include the U.S. HRW wheat price in the assessment (i.e., cointegration analysis) of U.S. and Canadian HRS wheat markets. Since HRW wheat is not produced in Canada, on the other hand, the Canadian HRW wheat price is not included in the model.

Tests for Unit Roots under Structural Change

Perron's procedure (1989) is used to determine the stationarity of price series with a potential structural break. Our preliminary investigations show that a breakpoint in each price series is identified to be around June 1986.[3] It is thus possible to apply Perron's method of testing for unit roots allowing for trend break at known break points. The results show that of the five series, the null hypothesis of non-stationarity cannot be rejected even at the 10 percent significance level for three of them: the U.S. HRS price, U.S. HRW price, and U.S. durum price (Table 1). However, the null hypothesis can be rejected at the 10 percent significance level for Canadian HRS and durum prices.[4] These results lead to the conclusion that the Canadian price series are stationary around a trend. Notice that, for comparison, we also estimate the standard ADF statistics for the series. The results of the ADF test indicate that the null hypothesis of non-stationarity cannot be rejected for all five series.

Given the results of the Perron procedure, it is no longer appropriate to use the full sample that includes stationary Canadian price series in cointegration analysis. As an alternative, therefore, we divide the full sample into two sub-samples, dataset I (July 1979–June 1986) and dataset II (July 1986–June 2002), according to the significant break point (June 1986), and check to see if this feature is stable in both cases. For this purpose, we apply the Dickey-Fuller generalized least squares (DF-GLS) test (of the null of non-stationarity) (Elliott, Rothenberg, and Stock 1996, Ng and Perron 2001) and the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) test (of the null of stationarity) (Kwiatkowski 1992) for the two subsamples.[5] The results show that, for all five price series in dataset I, the DF-GLS test rejects the null of non-stationarity, while the KPSS test cannot reject the null of stationarity, at least at the 10 percent level (Table 1). For all five series in dataset II, on the other hand, the DF-GLS test cannot reject the null of non-stationarity, while the KPSS test rejects the null of stationarity. From these findings, we conclude that the price series in dataset II are non-stationary, while the price series in

[3] The break points found here coincide with the period for the implementation of the U.S. Export Enhancement Program (May 1985–July 1995). The objectives of the EEP were to help U.S. farm products become more competitive with those from countries that receive subsidies (particularly the EU), and to expand U.S. agricultural exports. As the major commodity, wheat comprised more than 80 percent of the value of all EEP-assisted sales during the period of implementation.

[4] Our investigations show that the Canadian durum price is potentially characterized by a trend function with both sudden changes in the level and in the slope (model C). For the other price series, on the other hand, it appears that there is a change in the slope of the trend function with a constant level (model B). To test for unit roots in the presence of structural breaks, therefore, we apply equation (3) (model C) for the Canadian durum price and equation (2) (model B) for the other four prices.

[5] It should be emphasized that, when dealing with finite samples (i.e., small numbers of observations), the power of the standard ADF test is notoriously low (Harris and Sollis 2003). In addition, Maddala and Kim (1998) argue for the usefulness of performing tests of the null hypothesis of stationarity in addition to tests of the null hypothesis of a unit root (confirmatory data analysis). As such, it is prudent to conduct the DF-GLS and KPSS tests to provide overwhelming evidence of non-stationarity in the two sub-samples.

Table 1. ADF, Perron, and DF-GLS Tests of the Null Hypothesis of Non-Stationarity, and KPSS Test of the Null Hypothesis of Trend Stationarity [a]

	Full sample (July 1979–June 2002)		Dataset I (July 1979–June 1986)			Dataset II (July 1986–June 2002)		
	ADF	Perron	ADF	DF-GLS	KPSS	ADF	DF-GLS	KPSS
USH_t	-2.79	-3.32	-2.61	-2.52*	0.062	-2.65	-1.99	0.21**
CAH_t	-3.03	-3.67*	-2.60	-2.53*	0.061	-3.07	-2.29	0.17**
USW_t	-2.86	-2.96	-3.11	-3.31**	0.078	-2.38	-2.10	0.15**
USD_t	-2.87	-3.47	-2.68	-2.85**	0.077	-2.64	-1.72	0.14**
CAD_t	-3.02	-3.99*	-2.85	-2.56*	0.079	-2.63	-2.09	0.15**

[a] "ADF" is Augmented Dickey-Fuller test. "DF-GLS" is Dickey-Fuller Generalized Least Squares test. "KPSS" is Kwiatkowski, Phillips, Schmidt, and Shin test of the null of stationarity (Kwiatkowski et al. 1992).

Notes: ** and * denote rejection of the null hypothesis of non-stationarity at the 5 percent and 10 percent significance levels, respectively. The 5 percent and 10 percent critical values for the ADF including a constant and a trend are -3.44 and -3.14, respectively. The 5 percent and 10 percent critical values for the Perron test are -3.87 and -3.58 for Model B, and -4.17 and -3.87 for Model C. The 5 percent and 10 percent critical values for the DF-GLS test are -2.75 and -2.47 for dataset I, and -2.80 and -2.58 for dataset II. The 5 percent and 10 percent critical values for the KPSS test of trend stationarity are 0.15 and 0.12, respectively.

dataset I are stationary, which means that they cannot be used for the standard cointegration analysis.

It should be noted that, since one of the key motivations of this study is to examine the long-run relationship of the price series, it is also interesting to know if a break occurs in this long-run relationship rather than in the individual price series. For this purpose, we employ the most recent Johansen cointegration technique that allows for structural breaks at known points in time (Johansen, Mosconi, and Nielson 2000). Our implementation of the cointegration tests thus proceeds as follows: (i) the standard Johansen method is conducted with dataset II (case I), and (ii) the new Johansen technique is implemented with the full sample (case II).[6] This multi-step approach will help us understand the economic and policy implications contained in the long-run relationship, as well as lead to more robust empirical findings.

Empirical Results

Johansen Cointegration Test

Before implementing cointegration analysis, we first determine lag length (k) and conduct diagnostic tests for the residuals of price series. Using the likelihood-ratio (LR) test, the VAR models with k = 2 are determined for both HRS and durum wheat markets in cases I and II. Diagnostic tests on the residuals of each equation and corresponding vector test statistics support the VAR models with two lags (k = 2) as sufficient descriptions of the data (Table 2). More specifically, in a serial correlation test using the F-form of the Lagrange Multiplier (LM) procedure, the null hypothesis of no serial correlation cannot be rejected at the 5 percent significance level. In the heteroskedasticity test, the null hypothesis of no heteroskedasticity cannot be rejected at the 5 percent level. Normality of the residuals is tested with the Doornik-Hansen method (Doornik and Hendry 1994), and the null hypothesis of normality can be rejected at the 5 percent level for the five price series in both cases. However, since non-normality of residuals does not bias the results for the Johansen's cointegration test, the test results can be considered valid (Gonzalo 1994).

[6] Since Canadian HRS and durum prices in the full sample are found to be stationary at the 10 percent significance level, we change their level of significance from 10 percent to 5 percent to treat them as non-stationary in case II. In addition, for case II, equation (4) is reformulated as follows:

$$\Delta X_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} X_{t-1} \\ tE_t \end{pmatrix} + \mu E_t + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \sum_{i=1}^{k} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + u_t,$$

where q is the number of divided sample periods, $D_{j,t-1} = 1$ if $t = T_{j-1} + i$ (for $j = 2, .., q$ and $i = 1, ..., k$) and 0 otherwise, and $E_t = 1$ if $T_{j-1} + k + 1 \le t \le T_j$ and 0 otherwise [see Dawson and Sanjuán (2006) for a detailed application].

Table 2. Diagnostic Tests for Residuals of Wheat Price Series

	Serial Correlation		Heteroskedasticity		Normality	
	Case I	Case II	Case I	Case II	Case I	Case II
ΔUSH_t	0.81 [0.58]	0.71 [0.66]	1.45 [0.19]	0.81 [0.58]	29.39 [0.00]**	44.25 [0.00]**
ΔCAH_t	0.94 [0.48]	0.81 [0.58]	0.58 [0.78]	0.41 [0.89]	88.88 [0.00]**	112.22 [0.00]**
ΔUSW_t	0.85 [0.55]	0.94 [0.48]	1.77 [0.11]	0.86 [0.52]	17.28 [0.00]**	36.13 [0.00]**
System	0.94 [0.61]	0.82 [0.82]	--	--	225.26 [0.00]**	311.36 [0.00]**
ΔUSD_t	1.06 [0.40]	1.29 [0.15]	0.46 [0.50]	1.21 [0.25]	27.74 [0.00]**	29.89 [0.00]**
ΔCAD_t	1.32 [0.25]	1.59 [0.12]	0.47 [0.49]	1.58 [0.14]	30.76 [0.00]**	33.93 [0.00]**
System	1.38 [0.11]	1.22 [0.13]	--	--	76.44 [0.00]**	81.62 [0.00]**

Notes: Δ denotes the first differences of the variables. *p* values are given in parentheses. ** indicates that the null hypothesis is rejected at the 5 percent significance level. Serial correlation of the residuals of individual equations and a whole system is examined using the *F*-form of the Lagrange-Multiplier (LM) test, which is valid for systems with lagged independent variables. Heteroskedasticity is tested using the *F*-form of the LM test. Normality of the residuals is tested with the Doornik-Hansen test (Doornik and Hendry 1994).

As mentioned above, two Johansen cointegration methods are used to determine the number of cointegrating vectors. The results show that, with both case I and case II, the trace tests reject the hypothesis of no cointegrating vector ($r = 0$) at the 5 percent level, but fail to reject the null of one cointegrating vector ($r = 1$) (Table 3). This result suggests that there is one stable long-run equilibrium relationship between U.S. and Canadian HRS and between U.S. and Canadian durum prices.

Identifying one cointegrating vector in both cases, the test for the long-run exclusion is conducted to examine whether any of the variables can be excluded from a cointegrating vector. The null hypothesis is formulated by restricting the matrix of long-run coefficients to zero ($\beta_i = 0$) (Johansen and Juselius 1992). With case I, for example, the null hypothesis can be rejected for U.S. and Canadian HRS and durum prices in both models. However, the null hypothesis cannot be rejected even at the 10 percent level for U.S. HRW in the HRS wheat model (Table 4). Similarly, with case II, the null hypothesis cannot be rejected only for U.S. HRW price. These findings indicate that U.S. HRW is not statistically relevant to the cointegrating space in both cases and can be excluded from the long-run relationship.

A parameter in speed-of-adjustment is then restricted to zero ($\alpha_i = 0$) to test long-run weak exogeneity (Johansen and Juselius 1992). The results show that the null hypothesis of weak exogeneity cannot be rejected for Canadian prices in the HRS and durum wheat markets in cases I and II (Table 4). These findings indicate that the Canadian prices are the driving variables in the system and significantly affect the long-run movements of U.S. HRS and durum wheat prices, but are not influenced by U.S. HRS and durum wheat prices. In other words, the Canadian prices are the determining factors and the U.S. prices are the adjusting variables of the long-run relationships in the models. Notice that U.S. HRW price is also found to be weakly exogenous in the system. Combined with the finding of the exclusion test, this implies that U.S. HRW price does not adjust to deviation from any equilibrium state defined by the cointegration relation, but is rather determined outside the model system. Hence, it seems safe for us to exclude U.S. HRW price from the HRS wheat model. In addition, with case II, we test whether the structural break occurs in the long-run relationships in the HRS and durum wheat markets. The results show that the null hypothesis of no existence of structural break can be

Table 3. Johansen Cointegration Tests of Wheat Price Series

		Null hypothesis	Eigenvalue	Trace statistics
Case I	USH_t and CAH_t and USW_t	H_0: r = 0	0.168	55.78 [0.00]**
		H_0: r ≤ 1	0.058	20.73 [0.19]
		H_0: r ≤ 2	0.048	9.43 [0.16]
	USD_t and CAD_t	H_0: r = 0	0.098	27.18 [0.03]**
		H_0: r ≤ 1	0.041	7.87 [0.20]
Case II	USH_t and CAH_t and USW_t	H_0: r = 0	0.146	67.85 [0.00]**
		H_0: r ≤ 1	0.053	22.68 [0.12]
		H_0: r ≤ 2	0.035	9.73 [0.14]
	USD_t and CAD_t	H_0: r = 0	0.113	34.62 [0.00]**
		H_0: r ≤ 1	0.006	1.92 [0.96]

Notes: ** denotes rejection of the null hypothesis at the 5 percent significance level. *p* values are given in parentheses.

Table 4. Exclusion and Weak Exogeneity Tests of Wheat Price Series

	Case I		Case II	
	Exclusion	Weak Exogeneity	Exclusion	Weak Exogeneity
	$H_0 : \beta_i = 0$	$H_0 : \alpha_i = 0$	$H_0 : \beta_i = 0$	$H_0 : \alpha_i = 0$
Variable	(LR test statistic)	(LR test statistic)	(LR test statistic)	(LR test statistic)
USH_t	15.82 [0.00]**	4.96 [0.03]**	21.02 [0.00]**	6.24 [0.01]**
CAH_t	20.53 [0.00]**	2.52 [0.11]	26.23 [0.00]**	2.05 [0.15]
USW_t	1.51 [0.22]	0.26 [0.62]	1.44 [0.23]	0.11 [0.74]
USD_t	6.16 [0.01]**	6.14 [0.04]**	25.08 [0.00]**	6.20 [0.01]**
CAD_t	8.29 [0.00]**	3.09 [0.21]	29.95 [0.00]**	1.52 [0.22]

Notes: β_i and α_i represent a matrix of long-run coefficients and the speed of adjustment to equilibrium, respectively. LR test statistic is based on the χ^2 distribution and parentheses are *p*-values. ** denotes the rejection of the null hypothesis at the 5 percent significance level.

rejected in both markets: χ^2 = 9.68 (*p*-value = 0.00) for the HRS market and χ^2 = 12.22 (*p*-value = 0.00) for the durum market. This suggests that, due to the U.S. Export Enhancement Program, price shifts that occurred in June 1986 significantly affect the long-run relationships between the U.S. and Canadian wheat prices.

Finally, the long-run coefficients (β) explain the cointegrating relationships between the price series. For example, the long-run equilibrium relationships in the HRS and durum markets in both cases are represented as follows:

Case I:

(6) $USH_t = 0.84CAH_t$ and $USD_t = 0.79CAD_t$

Case II:

(7) $USH_t = 0.83CAH_t + 0.96E_{1,t} + 1.03E_{2,t}$

and $USD_t = 0.92CAD_t + 0.56E_{1,t} + 0.65E_{2,t}$.

Since CAH_t and CAD_t in both cases are weakly exogenous, we normalize the cointegrating vector on USH_t and USD_t. In addition, because the cointegrating relationships in equations (6) and (7) are identified, the coefficients can be interpreted as the long-run elasticities; for example, a 1 percent increase in $CAH_t(CAD_t)$ causes a 0.83–0.84 percent (0.79–0.92 percent) increase in $USH_t(USD_t)$ in both cases. Further, positive coefficients of the Canadian prices on the U.S. prices in equations

(6) and (7) suggest that U.S. HRS and durum wheat are substitutes for Canadian HRS and durum wheat, respectively. Notice that equation (7) verifies that the structural break has a significant impact on the long-run relationship between the U.S. and Canadian HRS and durum prices. In the HRS wheat market, for example, the breakpoint in June 1986 leads to an increase of approximately 0.07 (1.03–0.96) of the stationary difference (USH_t – 0.83CAH_t), or to an increase of approximately 7 percent in the price ratio between the United States and Canada. In addition, the structural break in the durum market results in an increase of approximately 9 percent in the price ratio.

VEC Model[7]

The VEC model is estimated to identify the short-run adjustment to long-run steady states as well as the short-run dynamics between U.S. and Canadian HRS wheat prices and between U.S. and Canadian durum prices. For this purpose, we estimate the short-run VAR model in equation (5), with the identified cointegration relationship. We adopt a general-to-specific procedure to estimate the VEC model (Hendry 1995, Harris and Sollis 2003). Specifically, the VEC models are first estimated with the same number of lags used in our cointegration analysis. The dimensions of the parameter space are then reduced to the parsimonious VEC (PVEC) models based on tests of the significance of the variables. The multivariate diagnostic tests on the estimated PVEC as a system show no serious problems with serial correlation, heteroskedasticity, and normality (Table 5). This suggests that the PVEC specifications do not violate any of the standard assumptions.

The negatively significant coefficients of error-correction terms represent the short-run adjustment speed of the dependent price series to the long-run equilibrium position. The results show that the error-correction terms for U.S. prices are negatively significant at the 5 percent level in both the U.S. HRS and durum equations (Table 5). This suggests that U.S. wheat prices adjust to the long-run equilibrium of U.S. and Canadian prices; approximately 12–13 percent of the adjustment occurs in one month for both the U.S. HRS and durum wheat prices. On the other hand, the error-correction terms for Canadian prices are not significant at the 5 percent level in both the U.S. HRS and durum equations. This implies that Canadian prices do not adjust to correct long-run disequilibria between U.S. and Canadian prices. These findings substantiate the results of our cointegration analysis; Canadian prices are weakly exogenous to the HRS and durum markets.

The coefficients of the lagged variables in the PVEC models show the short-run dynamics (causal linkage) between U.S. and Canadian wheat prices. In the HRS market, one-period lagged Canadian price is positively correlated with HRS wheat prices and its own price. Likewise, in the durum market, one-period lagged Canadian price is positively correlated with U.S. and Canadian prices. These results indicate that Canadian prices seem to have had significant short-run dynamic effects on U.S. prices in the HRS and durum markets for the period of 1986–2002.

It should be pointed out that the price quotations represent asking price and may not accurately reflect actual transaction prices, due mainly to government subsidies and other export promotion policies (Spriggs, Kaylen, and Bessler 1982, Goodwin and Schroeder 1991, Goodwin and Smith 1995, Mohanty, Peterson, and Smith 1996). For example, the asking prices do not include the U.S. Export Enhancement Program (EEP) bonus, since they are asking prices for all countries, not just for those targeted by the EEP. In addition, the average monthly EEP bonus provided by the U.S. government was substantial in monetary values over the 10-year span during which significant quantities of wheat shipments received the bonuses. As such, if we include the EEP bonus in the average prices series, we might obtain different empirical results; our findings should thus be viewed with caution.

Conclusions and Implications

This study examines the dynamics of price relationships in the U.S. and Canadian HRS and durum wheat markets. The Johansen cointegration analysis and VEC model are adopted with monthly prices for 1979–2002. Unlike previous studies,

[7] Since similar results are obtained from the two cases, we report the results of only case I here.

Table 5. Parsimonious VEC Models for U.S. and Canadian HRS and Durum Wheat Prices with Dataset II (July 1986–June 2002)

	HRS Wheat Price Series		Durum Wheat Price Series	
	ΔUSH_t	ΔCAH_t	ΔUSD_t	ΔCAD_t
ΔUSH_{t-1}	0.08 (0.73)	-0.12 (-1.32)	--	--
ΔCAH_{t-1}	0.26 (2.14)**	0.45 (4.29)**	--	--
ΔCAD_{t-1}	--	--	0.42 (5.33)**	0.53 (7.31)**
ΔCAD_{t-2}	--	--	-0.10 (-1.22)	-0.15 (-2.07)**
Error-correction	-0.13 (-2.15)**	0.09 (1.20)	-0.12 (-2.47)**	0.05 (1.18)
Constant	-0.13 (-2.14)**	0.09 (1.77)*	-0.03 (-2.28)**	0.02 (1.25)
Serial Correlation	$F_{AR}(28,318) = 0.67[0.90]$		$F_{AR}(28,340) = 1.26\ [0.18]$	
Heteroskedasticity	$F_{ARCH}(18,467) = 0.45\ [0.86]$		$F_{ARCH}(18,498) = 1.29\ [0.26]$	
Normality	$\chi^2(4) = 5.30\ [0.48]$		$\chi^2(4) = 7.43\ [0.11]$	

Notes: ** and * indicate significance at the 5 percent and 10 percent levels, respectively. Parentheses in multivariate diagnostic tests are *p*-values.

we first pay close attention to the issue of how we should conduct unit root tests with a possible structural change, which could affect all inferential procedures associated with unit roots and cointegration tests (Maddala and Kim 1998). The results provide statistical evidence that the price instability witnessed in June 1986 has caused structural change for Canadian HRS and durum prices. The break point coincides with the period for the implementation of the U.S. Export Enhancement Program. The results of our cointegration tests also show that the structural break indeed has a significant effect on the long-run relationship between the U.S. and Canadian HRS and durum prices. Furthermore, the Canadian HRS and durum prices are consistently found to be weakly exogenous in the HRS and durum wheat markets, implying that the U.S. HRS and durum prices are affected by the Canadian prices but that the reverse does not hold. Therefore, we conclude that the Canadian prices have a significant impact on the determination of the U.S. prices in both the HRS and durum markets.[8] One possible explanation for Canada's dominant role in price-setting is that U.S. exports of HRS and durum wheat are mostly driven by a number of private companies such as Cargill, Continental, and Louis-Dreyfus (Goodwin and Smith 1995), while the CWB sets its export prices in response to the international market situation and exercises a certain degree of market power in the North American market (Wilson, Johnson, and Dahl 1999, Lavoie 2002). Another explanation is that Canadian wheat is superior in quality to U.S. wheat and tends to lead the prices of other wheat in the international market (Ghoshray and Lloyd 2003).

This study has important implications for understanding and dealing with market behavior in North America. First, our unit root and cointegration tests demonstrate that identifying correct stationarity properties in time-series data is critical to proper estimation of structural relationships. Specifically, non-stationarity implies that shocks to any time-series data have permanent effects, while stationarity means that fluctuations

[8] In other words, our conclusion can be interpreted to mean that Canada plays a dominant role in price-setting in the North American wheat market. For this interpretation, we would use the term "price dominance" rather than "price leadership" as seen in previous studies (i.e., Mohanty, Peterson, and Smith 1996), because the term "price leadership" generally conveys a high degree of market power by Canadian producers or marketing agencies (i.e., Canadian Wheat Board). In this respect, the interpretation by Goodwin and Smith (1995) that the CWB acts as a price leader in international markets is appropriate, because the CWB is a marketing agency with the potential for market power in international markets through price discrimination.

are transitory. Time-series econometric models that do not explicitly identify structural breaks could classify all innovations as permanent (i.e., unit roots), when in fact the only permanent event is a structural shift and the remaining innovations are transitory (Oehmke and Schimmelpfennig 2004). Without correctly accounting for structural breaks, therefore, standard tests for cointegration and an error-correction model could lead to spurious short- and long-run relationships.

From these inferences, we can further draw a few statements regarding the Canadian price series: (i) given the different mean of the innovations for dataset I (July 1979–June 1986) and dataset II (July 1986–June 2002), it seems reasonable to conclude that the EEP has resulted in permanent structural change for the Canadian price series, which substantiates the findings of Dawson and Sanjuán (2006); and (ii) the discovery of the Canadian price series being stationary before the break and non-stationary after the break implies that any market shocks to the Canadian price series (e.g., any export promotion program such as the EEP) after June 1986 would have significant long-run effects on the North American wheat market. From the U.S. perspective, on the other hand, this suggests that, without taking into account the Canadian response, any export promotion programs implemented by the United States to enhance its wheat exports may have little impact on the market. For example, the EEP drove the wedge between lower foreign price and higher domestic price. The Canada-U.S. Free Trade Agreement (CUSTA) of 1989 and the North American Free Trade Agreement (NAFTA) of 1994 were to eliminate trade barriers in order to increase bilateral agricultural trade between the United States and Canada. Given Canada's dominant role in price-setting in the North American market after the structural break and the existence of an integrated wheat market in North America, it is apparently possible for the CWB to sell more wheat at more advantageous prices in a U.S. market that has no trade barriers, and to undermine the effectiveness of the export promotion policy (Mohanty, Peterson, and Smith 1996, Dawson and Sanjuán 2006). This further supports the concerns of U.S. wheat producers who claim that the CWB provides an unfair advantage to Canadian producers in the world market. However, it is important to recognize other factors that affect the North American wheat market, such as the high quality and low production costs of Canadian wheat.

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