Price Dynamics in the North American Wheat Market

Jungho Baek
Research Assistant Professor
Department of Agribusiness and Applied Economics
North Dakota Sate University
Email: jungho.baek@ndsu.nodak.edu
Phone: (701) 231-7451
Fax: (701) 231-7400

Won W. Koo
Professor
Department of Agribusiness and Applied Economics
North Dakota Sate University
Email: wkoo@ndsu.edu
Phone: (701) 231-7448
Fax: (701) 231-7400

Selected Paper prepared for presentation at the American Agricultural Economics Association Annual Meeting, Providence, Rhode Island, July 24-27, 2005

Price Dynamics in the North American Wheat Market

Abstract: This study examines price dynamics in the U.S. and Canadian hard red spring (HRS) and durum wheat markets. Using monthly prices for 1979-2002, we adopt Johansen cointegration tests and a vector error-correction (VEC) model. The results show that U.S. hard red winter (HRW) and Canadian HRS are exogenous in the model consisting of U.S. HRW and HRS and Canadian HRS prices. Canadian durum is exogenous in the model of U.S. and Canadian durum prices. Therefore, the results suggest that the HRW exporting industry and Canada have been the price leader in North American wheat markets.

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

Introduction

Wheat dominates international trade in cereals with exports of 118 million metric tons (MMT) in 2001 (International Grain Council 2002). The major wheat exporting countries are the U.S., Canada, Australia, the EU, and Argentina. These five countries represent approximately 70% of the wheat traded in the world market. The U.S. and Canada are the largest exporters, followed by Australia and the EU. The U.S. leads in exports of hard red winter (HRW) and soft red winter (SRW) wheat: an annual average of 16.7 MMT of HRW and 8.2 MMT of SRW during 1997-2001. Canada is the leader in exports of hard red spring (HRS) and durum wheat; an annual average of 14.4 MMT of HRS and 3.5 MMT of durum during 1997-2001. Further, the U.S. exports HRS and durum wheat and strongly competes with Canada in the world market. As such, HRS and durum wheat have been at the core commodities of the U.S.-Canada wheat trade dispute. The objective of this study is to assess the dynamics of price relationships in the U.S. and Canadian HRS and durum wheat markets.

An understanding of price relationships in the North American wheat market is important in addressing market structure, price leadership, as well as in constructing correct models for price analysis (Goodwin and Schroeder 1991, Mohanty et al. 1996). 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 acts as the price leader and imperfect competition exists in the market. Or, if U.S. and Canadian wheat prices are cointegrated, it suggests that these two prices drift in a similar fashion in the long-run. Modeling the North American wheat market thus needs to incorporate the cointegration relationships; otherwise, the econometric models could give a biased estimation. More importantly, it is crucial to assess the price behavior to understand the on-going wheat dispute between the U.S. and Canada (Mohanty et al. 1996). For example, the discovery of the Canadian price leadership implies that the U.S. market is influenced by the Canadian market, but that the reverse does not hold. In other words, Canadian export subsidies have impacts on price changes in the U.S. market. Hence, the finding can be interpreted to support the U.S. claim that subsidized Canadian wheat has depressed the U.S. prices.

Several studies have been taken to analyze price relationships in the world wheat market (Spriggs et al. 1982, Goodwin and Schroeder 1991, Goodwin and Smith 1995, Mohanty et al. 1996 and 1999). However, relatively limited efforts have been made to estimate the dynamics of price relationships in the North American wheat market. To our knowledge, Mohanty et al. (1996) is the only study that has been done so far to estimate the relationships between U.S. and Canadian wheat prices. The researchers employ the cointegration and error correction approach to estimate long-run and short-run wheat price relationships simultaneously. This study finds that there is no significant short-run causality between U.S. and Canadian wheat prices, while Canada is the price leader in the long-run.

With the recent development of the wheat dispute, this study is motivated by the need for more thorough and up-to-date analysis of price relationships in the North American wheat market. As Mohanty et al. (1996) did, we use the Johansen multivariate cointegration tests and a vector error-correction (VEC) model. Unlike the previous study, however, we take into account alternative model specifications in order to find a correctly specified a vector autoregressive (VAR) model. Specifically, foremost among the limitation of the Johansen test is that the procedure is sensitive to specifications, particularly the selection of the number of lags (Maddala and Kim 1998). We thus make efforts to adopt as reasonable model specifications as possible based on a number of VAR lag selection criteria and diagnostic tests. In addition, we use the recent concept of a general-to-specific procedure to construct a VEC model (Harris and Sollis 2003). Our dynamic modeling thus starts with a general statistical model, which captures the essential characteristics of the underlying dataset. Standard testing procedures are then used to reduce model complexity by eliminating statistically insignificant variables, as well as to check the validity of the reductions. Finally, we incorporate market structural break in our testing (Maddala and Kim 1998), which can have a substantive impact on estimated results, but has been largely ignored by previous studies of agricultural products markets. This comprehensive and up-to-date analysis is expected to enhance our understanding of price dynamics in the North American wheat market and contribute to the literature of the trade dispute.

The remainder of the paper is organized as follows. The data and empirical procedures are described in the next section, following by the empirical results. Summary and conclusions are presented in the final section.

Data and Empirical Procedures

Data

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

The variables that merit a brief mention are hard red winter (HRW) wheat and the dummy for the Wheat Peace Agreement (WPA; DUM_1). First, the dominant wheat class for the U.S. exports is HRW wheat, while for Canada it is HRS wheat. In addition, HRW wheat is 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). To incorporate this relationship, therefore, U.S. HRW wheat price is included in the assessment of U.S. and Canadian HRS wheat prices. Note that since HRW wheat is not produced in Canada, Canadian HRW wheat price is not included in the model. Second, the WPA is a trade restriction to regulate wheat imports from Canada for the period of October 1994 to September 1995. Under the WPA, Canada is allowed to export 0.3 million tons of durum and 1.05 million tons of other wheat to the U.S. during the period at the existing NAFTA tariff rates. Specifically, shipments of durum between 0.3 million and 0.45 million tons are subject to a fee of $23/ton. Shipments over 0.45 million tons of durum and 1.05 million of other wheat are subject to a fee of $50/ton. As such, the WPA may cause

Canadian producers to shift export patterns toward other countries, which in turn result in changes in U.S. and Canadian export prices. To protect against bias from overlooking such an effect, therefore, the dummy variable is included in our assessments.

Finally, it should be noted that the price quotations are asking price and thus may not accurately reflect actual transactions prices. Since there is a large body of literature on the discussion on relationships between quoted and realized wheat prices (Spriggs et al. 1982, Goodwin and Schroeder 1991, Mohanty et al. 1995 and 1999, Goodwin and Smith 1995), the details of this topic are not repeated here.

Empirical Procedures

Our approach starts with the vector autoregressive (VAR) model as follows:

$$X_t = A_1 X_{t-1} + ... + A_k X_{t-k} + \mu + u_t \qquad (1)$$

where X_t is a ($1 \times n$) vector of endogenous variables; for example, $X_t = [USH_t, CAH_t, USW_t]'$ for HRS wheat prices and $X_t = [USD_t, CAD_t]'$ for durum wheat prices; A_k is an ($n \times n$) matrix of parameters; μ is a vector of constant; k is the lag length; and u_t is a vector of normally and independently distributed error terms. In many cases, price series are non-stationary. Ordinary least squares (OLS) regression between such series thus lead to a spurious regression problem (Wooldridge 2000). To avoid this problem, non-stationary variables should be differenced to make them stationary. However, Engle and Granger (1987) show that even in the case that all the variables in a model are non-stationary, it is possible for a linear combination of integrated variables to be stationary. In this case, the variables are cointegrated and the problem of spurious regression does not arise. Hence, the first requirement for cointegration analysis is that the price series must be non-stationary.

The augmented Dickey-Fuller (ADF) test is commonly used to test for unit roots of price series (Dickey and Fuller 1979). However, the usual ADF test is unable to detect the structural break in a series because of its implicit assumption that the deterministic trend is correctly specified (Maddala and Kim 1998). In other words, if there is a break in the deterministic trend, then the ADF test may have lower power and even could lead to a false conclusion that there is a unit root when in fact there is not, or vice versa. Therefore, Perron's (1989 and 1997) test is used in this study in order to overcome the shortcomings of the ADF test, as well as to examine if there is any evidence of a structural break in the price series.

If price series are non-stationary, a test for cointegration is identical to a test of long-run equilibrium. The cointegration approach used in this study is the maximum likelihood estimation method developed by Johansen (Johansen and Juselius 1990, Johansen 1995). In order to impose the cointegration constraint, equation (1) is reformulated as follows:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + u_t \qquad (2)$$

where Δ is the difference operator; $\Gamma_1,...,\Gamma_{k-1}$ are the coefficient matrices of short-term dynamics; and $\Pi = -(I - \Pi_1 + ... + \Pi_k)$ are the matrix of long-run coefficients. Equation (2) is a standard first-difference VAR process except for the term ΠX_{t-k}. In this model, all terms are stationary and the standard asymptotic results apply (Engle and Granger 1987). The coefficient matrix 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 $\beta^{'}$ is a matrix of long-run coefficients such that the term $\beta' X_{t-k}$ represents up to $(n-1)$ co-integration relationships in the system. The number of cointegration vectors, the rank of Π, in the model is determined by the likelihood ratio test (Johansen 1995).

If all variables in X_t are co-integrated, equation (2) can be reformulated into a vector error-correction (VEC) model as follows:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \alpha(\beta^{'} X_{t-1}) + \mu + u_t \tag{3}$$

where $\beta' X_{t-1}$ is a measure of the error or deviation from the equilibrium, which is obtained from residuals from the cointegrating vectors. Equation (3) incorporates both short-run and long-run effects. In other words, if the long-run equilibrium holds at any time, $\beta' X_{t-1}$ is equal to zero. During periods of disequilibrium, on the other hand, this term is non-zero and measures the distance the system is away 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.

Empirical Results

Unit Root and Diagnostic Tests

Perron's procedure (1989 and 1997) is introduced to determine the stationarity of price series with a structural break. Our preliminary investigations show that all the five price series are contains only one break, with June 1986 as the break point. Like other studies (Mohanty et al. 1996, Gardner 1999), it coincides with the implementation of the U.S. Export Enhancement Programs (EEP). The EEP was announced in May 1985 and continued until July 1995 in the case of wheat. The purpose of the EEP is to expand U.S. agricultural exports, as well as to challenge unfair trade practices (i.e., the EU) by U.S. farmers meet competition from subsidizing countries. Since a structural break in the series is known, it is possible to apply Perron's method of testing for unit roots.

The results show that of the five series, the unit root hypothesis cannot be rejected even at the 10% significance level for three of them: U.S. HRS price, U.S. HRW price, and U.S. durum price (Table 1). However, the null hypothesis can be rejected at 10% significance level for Canadian HRS and durum wheat prices. Notice that, for comparison, we also estimate the usual ADF statistics for the series. The results show that the null hypothesis for all the series cannot be rejected even at the 10% significance level. The findings thus indicate that the usual ADF test without considering the structural change in the series tends to under-reject the null hypothesis. Therefore, we conclude that the underlying process for Canadian HRS and durum wheat prices can be characterized by stationary fluctuations around a deterministic trend function.

According to the results of the Perron's procedure, it is no longer appropriate to use the full sample that includes stationary price series in our cointegration analysis. As an alternative, we divide the full sample into two sub-samples according to the break point. Then, the standard ADF procedure is applied for the two sub-samples. The results show that the unit root hypothesis cannot be rejected for the level series, but can be rejected for the first differences of them at the 5% significance level (Table 1). The ADF test statistics are estimated from a model that includes a constant and a trend variable. The test is performed with 1 to 10 lags, and the results are found to be invariant to the use of different number of lags. From these findings, we conclude that the price series in the sub-samples are non-stationary and integrated of order one, or $I(1)$; therefore, cointegration analysis is pursued on two sub-samples such as dataset I (1979:07-1986:06) and dataset II (1986:07-2002:-6).

Someone may wonder that since the unit root hypothesis is rejected only at 10% significance level for Canadian prices, the results do not provide strong evidence to support structural change and thus the division of the full sample into two sub-samples. However, as one of the usual statistical significance levels (Wooldridge 2000), the 10% significance level, if not strong, seems sufficient to provide statistical evidence that the EEP has caused structural change for Canadian prices. Failure to take into account of this structural change thus raises concerns the problems of parameter stability and spurious regression (Maddala and Kim 1998). Further, Mohanty et al. (1996) identify December 1985 as the break point induced by the EEP and estimate two models with sub-samples. However, they did not test the effect of EEP within the structural change framework. Accordingly, their division of the full sample into two sub-samples could result in efficiency loss and thus undermine the credibility of their findings.

To define a correctly specified VAR model for cointegration analysis, we determine lag length and conduct diagnostic tests for the residuals of price series. Maddala and Kim (1998) note that Johansen procedure is sensitive to changes in lag structure. The lag lengths (k) of the VAR model is determined by the Schwarz (SC), Hannan-Quinn (HQ) and Akaike (AIC) information criteria using likelihood ratio tests (Doornik and Hendry 1994). With HRS and HRW prices in dataset I, for example, we start from $k = 6$ and a reduction of the VAR from $k=6$ to $k=5$ is rejected. This reduction sequence is then conducted until we find that the reduction from $k=4$ to $k=3$ is accepted.

Similarly, with HRS and HRW prices in dataset II and durum prices in both datasets, the VAR models with k =2 are accepted for cointegration analysis.

Diagnostic tests on the residuals of each equation and corresponding vector test statistics support the VAR models with two lags (k =2) or three lags (k =3) as sufficient descriptions of the data (Table 2). In 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% significance level. In the heteroskedasticity test, the null hypothesis of no heteroskedasticity cannot be rejected at the 5% 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 for some of the price series at the 5% level. However, since non-normality of residuals does not bias the results for the Johansen's co-integration test, the test results can be considered valid (Gonzalo 1994).

Johansen Co-integration Test

The Johansen co-integration estimation is used to determine the number of co-integrating vectors in datasets I and II. The specification tests indicate that a linear trend is necessary but seasonal dummies are not. The results show that the trace tests reject the hypothesis of no cointegrating vector (r =0) at the 5% level, but fail to reject the null of one cointegrating vectors (r =1) in both datasets (Table 3). This result suggests that there is a stable long-run equilibrium relationship between U.S. and Canadian HRS and between U.S. and Canadian durum prices in datasets I and II.

With identifying one cointegrating vector in both datasets, a parameter in speed-of-adjustment (α) is restricted to zero to test long-run weak exogeneity (Johansen and

Juselius 1992). The results show that with both datasets, the null hypothesis of weak exogeneity cannot be rejected for U.S. HRW and Canadian prices in the HRS wheat markets (Table 4). Furthermore, the null hypothesis that both prices are weakly exogenous cannot be rejected for dataset I ($\chi^2(2)$=1.24, p-value=0.54) and dataset II ($\chi^2(2)$ = 4.83, p-value=0.11). These findings indicate that both U.S. HRW and Canadian HRS wheat prices do not adjust to deviations from any equilibrium state defined by the cointegration relation. In other words, these two prices are the driving variables in the system and significantly affect the long-run movements of U.S. HRS wheat prices, but these two prices are not influenced by U.S. HRS wheat prices. Similarly, it is found that the null hypothesis that Canadian durum prices are weakly exogenous cannot be rejected in both datasets (Table 4). This result also suggests that the Canadian durum price is the determining part and the U.S. durum price is the adjusting part of the long-run relationship. Notice that U.S. HRW and Canadian prices are consistently found to be weakly exogenous for both pre- and post-structural break periods. This implies that the structural break (EEP) does not change the long-run equilibrium relationship between U.S. and Canadian prices.

Finally, the long-run coefficients (β) explain the cointegrating relationships between the price series. For example, the long-run equilibrium relationship in the HRS market is represented as follows:

$$\text{Dataset I:} USH_t = 0.77CAH_t + 0.16USW_t \qquad (4)$$

$$\text{Dataset II:} USH_t = 0.83CAH_t + 0.24USW_t \qquad (5)$$

Since CAH_t and USW_t are both weakly exogenous, we normalize the cointegrating vector on USH_t. In addition, because the cointegrating relationships in equations (4) and (5) are identified, as Johansen (2002) notes, the coefficients can be interpreted as the long-run (price transmission) elasticity; for example, a 1% increase in CAH_t causes a 0.77-0.83% increase in USH_t. Further, positive coefficients of USW_t on USH_t in both equations support that HRW wheat is a substitute for HRS wheat.

<u>VEC Model</u>

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 in both datasets. For this purpose, with the identified co-integration relationship, the short-run VAR model in equation (3) is estimated. 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 (Tables 5 and 6). This thus 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, with datasets I and II, the error-correction terms for U.S.

prices are negatively significant at the 5% level in both the U.S. HRS and durum equations (Tables 6 and 7). This suggests that U.S. wheat prices adjust to correct long-run disequilibria in U.S. and Canadian prices; that is, for HRS prices, about 12-37% of the adjustment occurs in one month and about 12-24% for durum prices. On the other hand, the error-correction terms for U.S. HRW and Canadian prices are not significant at the 5% level in both the U.S. HRS and durum equations and for both datasets. This implies that U.S. HRW and 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; that is, HRW and 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, U.S. HRW price in dataset I is statistically significant and positively correlated with one period lagged Canadian price and own price, but negatively correlated with one and two period lagged U.S. prices. In addition, one period lagged Canadian price in dataset II is positively correlated with HRS and HRW wheat prices and own price. In the durum market, on the other hand, one period lagged U.S. and Canadian prices in dataset I are positively correlated with their own prices, while one period lagged Canadian price in dataset II is positively correlated with U.S. and Canadian prices. These results thus indicate that Canadian prices seem to have significant short-run dynamic effects on U.S. prices in the HRS and durum markets during 1986-2002. Our results do not coincide with Mohanty et al. (1996) who find that there is no significant short-run dynamic effect between U.S. and Canada. Notice that due to insignificant coefficients, the dummy for the WPA is dropped in the PVEC model. This implies that the WPA has little impact on U.S. and Canadian prices.

Summary and Conclusions

This study examines the dynamics of price relationships in the U.S. and Canadian HRS and durum wheat markets. Using monthly prices for 1979-2002, the Johansen cointegration analysis and VEC model are adopted. Unlike previous studies, we first pay close attention to the issue relate to how we should conduct unit root tests with a possible structural change, which could affect all the 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 over which the U.S. Export Enhancement Program for wheat implemented. To consider the structural break in our cointegration analysis and VEC model, therefore, two models are estimated with sub-samples such as dataset I (1979:07-1986:06) and dataset II (1986:07-2002:06).

The results of our cointegration tests and VEC models show that Canadian HRS prices and U.S. HRW prices are weakly exogenous in the HRS markets, implying that these two prices significantly influence U.S. HRS prices, but are not affected by U.S. HRS prices. Similarly, as a weakly exogenous variable, Canadian price have a significant impact on the determination of U.S. price in the durum markets. Therefore, we conclude that U.S. HRW exporting industry and Canada act as the price leaders and U.S. as the follower in the HRS market, while Canada is the price leader in the durum market. Our results substantiate those of Mohanty et al. (1996) and Ghoshray and Lloyd (2003). One of possible explanations for Canada's price leadership 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). In Canada, on the other hand, the Canadian Wheat Board (CWB) has responsibility for the marketing of all western Canadian wheat and durum. As such, the CWB enables to set export prices by responding to the international market situation and exercises a certain degree of market power in the North American market. Another explanation is that the Canadian wheat is superior in quality than the U.S. wheat and tends to lead the prices of other wheat in the international market (Ghoshray and Lloyd 2003). Additionally, in our cointegration analysis, U.S. HRW wheat is consistently found to be a substitute for HRS wheat. Further, U.S. HRW wheat traditionally has been the dominant product in U.S. wheat exports; for example, an annual average of 27.2 MMT of all wheat was exported in 1997-2001, of which 16.7 MMT were HRW and 8.2 MMT were HRS. Hence, U.S. HRW wheat has been an important factor in significantly affecting HRS wheat prices.

This study has important implications for econometric models of the North American wheat market. First, our unit root tests demonstrate that structural change does affect inference on stationarity. Hence, as Spriggs et al. (1982) elaborate, with estimating behavioral relationships with historical data, it is crucial to test for unit roots incorporating major policy shocks as structural break. Second, according to our cointegration analysis, it seems safe for us to treat the Canadian price as a weakly exogenous variable in the model. When estimating the North American wheat market, researchers thus need to consider this relationship in the model; otherwise, the econometric models, such as results of studies that model the U.S. as the price leader, could give a biased estimation.

Finally, another implication from out findings is that the U.S. responds to Canadian price change but Canada does not respond to U.S. price changes may support the concerns of U.S. wheat producers who contend that the CWB-led subsidized Canadian wheat unfairly markets Canadian wheat and undercut U.S. price. However, because we do not take into account issues related to the high quality and standardization of Canadian wheat, without further investigation we cannot say that Canadian wheat directly influences changes in U.S. prices; thus, our interpretation should be taken with cautious.

References

Dickey, D.A. and W.A. Fuller. 1979. Distribution of the estimators for auto-regressive time series with a unit root. Journal of the American Statistical Association 74: 427-431.

Doornik, J. and D., Hendry. 1994. Interactive econometric modeling of dynamic system (PcFiml 8.0). University of Oxford Institute of Economics and Statistics. London: International Thomson Publishing.

Engle, R. and C.W.J Granger. 1987. Co-integration and error-correction: representation, estimation and testing. Econometrica 55: 251-276.

Gardner, B.L. 1999. Canadian/US farm policies and the creation of a single North American grain market. In: J.M. Antle and V.H. Smith, (Ed.), The Economics of World Wheat Markets. CAB International, pp. 273-311.

Ghoshray, A. and T. Lloyd. 2003. Price linkages in the international wheat market. In: Proceedings of the 25th International Conference of Agricultural Economists. Durban, South Africa, pp. 1095-1108.

Gilmour, B. and P. Fawcett. 1987. The relationship between U.S. and Canadian wheat prices. Canadian Journal of Agricultural Economics 35: 571-589.

Gonzalo, J. 1994. Five alternative methods of estimating long-run equilibrium relationships. Journal of Econometrics 60: 203-233.

Goodwin, B.K. and T.C. Schroeder. 1991. Dynamics in international wheat markets. Canadian Journal of Agricultural Economics 39(2): 237-54.

Goodwin, B.K. and V.H. Smith. 1995. Price discrimination in international wheat market. A report prepared for the wheat export trade education committee. Raleigh: North Carolina State University, Department of Agricultural Economics.

Harris, R. and R. Sollis. 2003. Applied Time Series Modeling and Forecasting. Chichester, W. Sussex: John Wiley and Sons.

Hendry, D. 1995. Dynamic Econometrics. Oxford, UK: Oxford University Press.

Johansen, S. 1995. Likelihood-based inference in co-integrated vector auto-regressive models. Oxford, UK: Oxford University Press.

Johansen, S. and K. Juselius. 1990. Maximum likelihood estimation and inference on co-integration — with application to the demand for money. Oxford Bulletin of Economics and Statistics 52: 169-210.

Johansen, S. and K. Juselius. 1992. Structural tests in a multivariate co-integration analysis of the PPP and the UIP for UK. Journal of Econometrics 53: 211-244.

Koo, W.W. and J.M. Mattson. 2002. U.S.-Canada wheat trade and its effects on U.S. price and income. Agribusiness & Applied Economics Report No. 492: Center for Agricultural Policy and Trade Studies, North Dakota State University.

Maddala, G.S. and In-Moo Kim. 1998. Unit roots, co-integration, and structural change. Cambridge, UK: Cambridge University Press.

Mohanty, S.E., W.F. Peterson and D.B. Smith. 1996. Price relationships between U.S. and Canadian wheat prices: Co-integration and error-correction approach. Canadian Journal of Agricultural Economics 44(3):265-277.

Mohanty, S.E., W.H. Meyers and D.B. Smith. 1999. A re-examination of price dynamics in international wheat market. Canadian Journal of Agricultural Economics 47(1): 21-29.

Perron, P. 1989. The great crash, the oil shock and the unit root hypothesis. Econometrica 57: 1361-1402.

Perron, P. 1997. Further evidence on breaking trend functions in macroeconomic variables. Journal of Econometrics 80: 355-385.

Spriggs, J., M. Kaylen and D.Bessler. 1982. The lead-lag relationship between Canadian and US wheat prices. American Journal of Agricultural Economics 64(3):569-572.

Wooldridge, J.M. 2000. Introductory Econometrics: A Modern Approach. South-Western College Publishing.

World Grain Statistics. Various years. International Grain Council.

World Wheat Statistics. Various years. International Wheat Council.

Table 1. Unit root tests of wheat price series with a structural break and ADF tests with sub-samples

	Perron's test	ADF test				
		Full sample (1979:07-2002:06)	Dataset I (1979:07-1986:06)		Dataset II (1986:07-2002:06)	
			Level	First Difference	Level	First Difference
USH_t	-3.25	-2.79	-2.61	-4.04**	-2.65	-8.58**
CAH_t	-3.88*	-3.03	-2.60	-4.07**	-3.07	-8.57**
USW_t	-2.98	-2.86	-3.11	-4.52**	-2.38	-6.01**
USD_t	-3.46	-2.87	-2.68	-5.59**	-2.64	-9.21**
CAD_t	-4.01*	-3.02	-2.85	-5.86**	-2.63	-8.17**

1. ** and * denote rejection of null hypothesis of a unit root at the 5% and 10% significance levels, respectively.
2. The 5% and 10% critical values for unit root tests with a structural break (Perron's method) are -4.17 and -3.87, respectively. Critical values are obtained from Table 6B in Perron (1989).
3. The 5% and 10% critical values for the ADF including a constant and a trend are -3.44 and -3.14, respectively.

Table 2. Diagnostic tests for residuals of wheat price series

	Dataset I (1979:07-1986:06)			Dataset II (1986:07-2002:06)		
	SC test	Hetero test	Normality	SC test	Hetero test	Normality
ΔUSH_t	0.42 [0.83]	0.66 [0.52]	6.00 [0.06]*	0.81 [0.58]	1.45 [0.19]	29.39 [0.00]**
ΔCAH_t	0.30 [0.91]	1.27 [0.29]	4.11 [0.13]	0.94 [0.48]	0.58 [0.78]	88.88 [0.00]**
ΔUSW_t	1.73 [0.15]	0.03 [0.97]	0.21 [0.90]	0.85 [0.55]	1.77 [0.11]	17.28 [0.00]**
System	1.38 [0.11]	-	1.58 [0.95]	0.94 [0.61]	-	225.26 [0.00]**
ΔUSD_t	1.27 [0.29]	1.58 [0.18]	13.50 [0.00]**	1.06 [0.40]	0.46 [0.50]	27.74 [0.00]**
ΔCAD_t	0.44 [0.82]	0.48 [0.79]	12.68 [0.00]**	1.32 [0.25]	0.47 [0.49]	30.76 [0.00]**
System	1.14 [0.32]	-	12.78 [0.01]**	1.38 [0.11]	-	76.44 [0.00]**

1. SC and hetero test represent serial correlation and heteroskedasticity test, respectively.
2. Δ denotes the first differences of the variables.
3. ** and * indicate that the null hypothesis is rejected at the 5% and 10% significance levels, respectively.
4. Serial correlation of the residuals of individual equations and a whole system was examined using the F-form of the Lagrange-Multiplier (LM) test, which is valid for systems with lagged independent variables.
5. Heteroskedasticity was tested using the F-form of the LM test.
6. Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hendry 1994).

Table 3. Johansen cointegration tests of wheat price series

Dataset I (1979:07-1986:06)	Null hypothesis	Eigenvalue	Trace statistics	5% critical value
$USH_t \& CAH_t \& USW_t$	H_0: r = 0	0.294	47.27**	42.44
	H_0: r ≤ 1	0.126	19.12	25.32
	H_0: r ≤ 2	0.097	8.25	12.25
$USD_t \& CAD_t$	H_0: r = 0	0.230	28.17**	25.32
	H_0: r ≤ 1	0.083	6.97	12.25
Dataset II (1986:07-2002:06)	**Null hypothesis**	**Eigenvalue**	**Trace statistics**	**5% critical value**
$USH_t \& CAH_t \& USW_t$	H_0: r = 0	0.168	55.78**	42.44
	H_0: r ≤ 1	0.058	20.73	25.32
	H_0: r ≤ 2	0.048	9.43	12.25
$USD_t \& CAD_t$	H_0: r = 0	0.098	27.18**	25.32
	H_0: r ≤ 1	0.041	7.87	12.25

1. ** and * denote rejection of the hypothesis at the 5% and 10% significance levels, respectively.

Table 4. Weak exogeneity tests of wheat price series

Weak exogeneity $H_0 : \alpha_i = 0$	Dataset I (1979:07-1986:06)	Dataset II (1986:07-2002:06)
USH_t	4.141 [0.04]**	4.955 [0.03]**
CAH_t	0.046 [0.83]	2.522 [0.11]
USW_t	0.015 [0.90]	0.255 [0.62]
USD_t	6.93 [0.03]**	6.14 [0.04]**
CAD_t	4.38 [0.11]	3.09 [0.21]

1. LR test statistic is based on the χ^2 distribution and parentheses are p-values.
2. ** and * denote significance at the 5% and 10% levels, respectively.
3. α_i represents the speed of adjustment to equilibrium.

Table 5. Results of parsimonious VEC models for U.S. and Canadian HRS wheat prices

	Dataset I (1979:07-1986:06)			Dataset II (1986:07-2002:06)		
	ΔUSH_t	ΔCAH_t	ΔUSW_t	ΔUSH_t	ΔCAH_t	ΔUSW_t
ΔUSH_{t-1}	0.43 (1.48)	0.33 (1.21)	0.56 (2.12)**	0.07 (0.69)	-0.12 (-1.34)	-0.15 (-1.19)
ΔUSH_{t-2}	0.11 (0.72)	-0.01 (-0.07)	0.54 (3.96)**			
ΔCAH_{t-1}	-0.05 (-0.15)	0.04 (0.12)	0.84 (2.87)**	0.26 (2.15)**	0.46 (4.36)**	0.35 (2.41)**
ΔUSW_{t-2}	-0.13 (-1.03)	-0.03 (-0.22)	0.23 (2.01)**			
Error-correction	-0.37 (-1.99)**	0.02 (0.08)	0.34 (1.63)	-0.12 (-2.22)**	0.09 (1.33)	-0.03 (-0.40)
Constant	-0.01 (-1.62)	-0.01 (-0.70)	0.08 (0.32)	-0.14 (-2.21)**	0.11 (1.91)*	-0.03 (-0.38)
SC test	$F_{AR}(45{,}128) = 1.19$ [0.22]			$F_{AR}(63{,}484) = 0.88$ [0.73]		
Hetero test	$F_{ARCH}(60{,}235) = 0.87$ [0.73]			$F_{ARCH}(36{,}714) = 1.23$ [0.13]		
Normality	$\chi^2(6) = 5.73$ [0.45]			$\chi^2(6) = 5.32$ [0.50]		

1. ** and * indicate significance at the 5% and 10% levels, respectively.
2. SC and hetero test represent serial correlation and heteroskedasticity test, respectively.
3. Parentheses in multivariate diagnostic tests are p-values.

Table 6. Results of parsimonious VEC models for U.S. and Canadian durum wheat prices

	Dataset I (1979:07-1986:06)		Dataset II (1986:07-2002:06)	
	ΔUSD_t	ΔCAD_t	ΔUSD_t	ΔCAD_t
ΔUSD_{t-1}	0.38 (3.62)**	0.29 (2.88)**		
ΔCAD_{t-1}			0.42 (5.33)**	0.53 (7.31)**
ΔCAD_{t-2}			-0.10 (-1.22)	-0.15 (-2.07)**
Error-correction	-0.24 (-2.19)**	0.09 (0.88)	-0.12 (-2.47)**	0.05 (1.18)
Constant	-0.03 (-2.36)**	0.01 (0.50)	-0.03 (-2.28)**	0.02 (1.25)
SC test	$F_{AR}(20,112) = 0.98$ [0.49]		$F_{AR}(28,340) = 1.26$ [0.18]	
Hetero test	$F_{ARCH}(12,188) = 1.17$ [0.31]		$F_{ARCH}(18,498) = 1.29$ [0.26]	
Normality	$\chi^2(4) = 8.24$ [0.18]		$\chi^2(4) = 7.43$ [0.11]	

1. ** and * indicate significance at the 5% and 10% levels, respectively.
2. SC and hetero test represent serial correlation and heteroskedasticity test, respectively.
3. Parentheses in multivariate diagnostic tests are p-values.