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FARM INPUT, FARM OUTPUT, AND FOOD RETAIL PRICES: A COINTEGRATION ANALYSIS

by

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ABSTRACT

This paper investigates the adjustment mechanism between farm input prices, farm output prices, and food retail prices in Canada. Johansen's maximum likelihood approach is used in addition to the Engle-Granger approach to test for cointegration. Contrary to the common assumption that farm output prices are more flexible than farm input prices, it is found that farm output prices though cointegrated are weakly exogenous in the sense that they do not respond in a systematic manner to disequilibrium in farm input prices and food retail prices. Evidence was found to support "Cost Push" and "Demand Pull" theories but since food retail prices carry a heavier weight in the cointegration relations, it can be concluded that shocks manifesting themselves (first) at the retail level do not persist as long.

RESUME

Cet ouvrage se concentre sur le méchanisme d'ajustement entre les prix des intrants à la ferme, les prix payés aux producteurs agricoles, et les prix de l'aligmentation au detail au Canada. Les techniques developpées par Johansen et Engle et Granger sont utilisées pour vérifier des hypothèses reliées à la cointegration. Contrairement à la présomption que les prix payés aux agriculteurs sont plus flexibles que les prix de leurs intrants, on a découvert que les prix payés aux agriculteurs bien que cointégrés sont exogènes au moindre degré dans le sens qu'ils ne s'ajustment pas en fonction du déséquilibre dans les prix des intrants et les prix de l'alimentation. La théorie voulant que les prix montent à cause des bausses dans les prix des intrants est vérifiée de même que la théorie alternative voulant que les ajustements des prix se fassent du détail aux intrants. Les prix de l'alimentation ont des effets plus prononcés sur les méchanismes d'ajustement que les prix des intrants.

Farm Input, Farm Output, and Food Retail Prices:
A Cointegration Analysis

Introduction

It is usually assumed that an increase in the price of inputs will eventually induce a rise in the price of finished goods. This phenomenon is often referred to as "Cost-Push" and has been investigated by several prominent economists. Engle (1978) modeled the relationship between the wholesale price of raw food products, the wage rate in the manufacturing of food, and the food component of the consumer price index. Guthrie (1981) analysed the relationship between the wholesale price index and the consumer price index. In contrast, "Demand-Pull" theory suggests that macroeconomic factors affect the demand for final goods which in turn affect the demand for inputs thus inducing changes in input prices. The empirical relevance of this theory was explored by Gordon (1975), Granger, Robbins, and Engle (1986), and in more narrowly defined studies like Sephton (1989) and Devadoss and Meyers (1987) which concentrated on the impact of a particular type of shock (i.e., monetary) on agricultural prices.

The speed at which prices adjust to changes in other prices is just as controversial a topic as the causal direction between prices. Tweeten (1980) argues that the terms of trade for farmers deteriorates as the rate of inflation rises.[1] At the other end of the spectrum, Starleaf, Meyers, and Womack (1985) have found empirical evidence to support the hypothesis that farm output prices adjust faster than farm input prices and consequently, that farmers must have been net beneficiaries of increased inflation rates. Others like Chambers (1983), Gardner (1979), and Prentice and Schertz (1981) obtained results supporting the neutrality of inflation on farmers' terms of trade.

From a theoretical point of view, it can be argued that farmers will benefit (lose) from an unexpected[2] rise in inflation if farm output prices are more (less) flexible than farm input prices. Chambers (1985) explored this argument more deeply and concluded that the real effects generated by monetary policies (when the prices of certain goods are more flexible than others) should ultimately be self-correcting. This reasoning is echoed by Sephton (1989) who states that money is non-neutral in the short-run but "acts only as a veil" in the long run. Therefore, the emphasis of this paper is on the the self-correcting mechanism(s) between farm input prices, farm output prices, and food retail prices and the implied long run equilibrium relationship(s).

The objective of this paper is twofold. First, the hypothesis concerning the existence of stable (long run) relationships between farm input prices, farm output prices, and retail food prices in Canada is tested. Secondly, the long run equilibrium relationships between prices will be analyzed and tests will be performed to determine the relative speed at which prices move back toward equilibrium.

The paper is organized as follows. A description of the methodology and the source of the data will be provided in the next section. The results are presented in the third section and their implications are discussed in the fourth section. The paper concludes with a brief summary.

Methodology and Data

Cointegration and error correcting models, originally suggested in Granger (1981) and extended in Engle and Granger (1987) and Johansen (1988) are used in this study. Several factors motivated this choice. First, the estimated parameters of a cointegrating regression converge more rapidly to their true value than the estimated parameters of a regression of stationary series (Fuller, 1976; Stock, 1987). Moreover, regressing non-stationary variables on each other often generate spurious relationships. Granger and Newbold (1986) have proven this point by obtaining high R^2's when regressing independent random walks. They warn applied econometricians against the reporting of high R^2's and low Durbin-Watson's, a clear indication of model mispecification. They go further in showing that in many instances commonly used procedures to correct for autocorrelation such as Cochrane-Orcutt or differencing are ineffective.

The following is a description of the Engle-Granger approach to cointegration. As in Granger and Newbold (1986), suppose that we are interested in the relationship between the components of $X_t=(x_{1t},...,x_{pt})$, a matrix made up of p series integrated of the first order (I(1)). Let R_t be the "equilibrium error" or the residuals from regressing the series in X_t on each other. The series in X_t will be cointegrated if there exists weights ß such that the following expression is integrated of order zero.

$$R_t = ß'X_t \text{ is } I(0) \tag{1}$$

Thus if R_t is stationary, the non-stationary series in X_t do not have to be differenced to obtain efficient parameters estimates. Because X_t is I(1), there will be a multivariate World representation:

$$(1\text{-}L)X_t = C(L)\epsilon_t, \tag{2}$$

where ϵ_t is white noise and L is the lag operator. Thus, the first differenced X_t is a finite variance ARMA process. By inserting equation (1) into (2) we can obtain:

$$(1\text{-}L)R_t = ß'C(L)\epsilon_t. \tag{3}$$

From equation (3), it is evident that for R_t to be integrated of degree 0, ß'C(1) must be a 1xN vector of zeros. This implies that C(1) must be singular and that equation (2) cannot take an autoregressive representation. However, Granger (1984) and Granger and Engle (1985) have shown that equation (2) can be represented by:

$$(1\text{-}L)X_t = A^*(L)(1\text{-}L)X_t\text{-}\Theta R_{t-1}+D(L)\epsilon_t, \tag{4}$$

where $A^*(0)=0$, Θ is a vector of constants, and $\det(C(L))=(1\text{-}L)D(L)$. Equation (4) is called the error correction equation. The term ΘR_{t-1} indicates the extent to which the economy (or the sector under investigation) is out of equilibrium. If $\Theta=0$, R_t will be I(1) and the components of X_t will not be cointegrated. In the case where the lag differenced X_t are excluded from (4), the restriction $\Theta=0$ is equivalent to testing for a random walk. Thus the acceptance of the null hypothesis of $\Theta=0$ corresponds to the rejection of cointegration. In terms of modeling, the rejection of cointegration implies that the usual vector ARMA approach can be used to investigate the behavior of X_t.

In practice, the methodology described above entails four steps. First, it must be established that the p series of interest have a non-seasonal unit root. This is accomplished via a series of tests (Dickey-Fuller, Augmented Dickey-Fuller, Engle-Granger-Hylleberg-Yoo, and Osborn-Chui-Smith-Birchenhall). The second step consists of estimating the parameters capturing the relationships between the variables. This is done by regressing the x_{it}'s on each other. The Dickey-Fuller and Augmented Dickey-Fuller tests on the residuals obtained from the regressions are then used to test the null of no-cointegration. The augmented Dickey-Fuller is carried out by regressing the first difference of the residuals on the lag residual and k lags of first differences. The hypothesis of no-cointegration is tested by comparing the t ratio on the lag residual to a critical value. Rejection of the null hypothesis is a signal to proceed with the fourth step which is to fit error correcting models like equation (4). Cointegration can also be tested with Phillips Z_α test which has the advantage of being more powerful than the Augmented Dickey-Fuller test in samples of moderate size (Phillips and Ouliaris, 1990). The Z_α test can be implemented by using the predicted values and coefficient estimates of the regression $R_t = aR_{t-1} + k_t$ where k_t is the residual and by computing $Z_\alpha = T(a-1) - (1/2)(S_{tl}^2-S_k^2)(T^{-2}\Sigma R_{t-1}^2)^{-1}$, where t ranges from 2 to T, $S_k^2 = T^{-1}\Sigma k_t^2$ for $t=1,...,T$, and $S_{tl}^2 = S_k^2 + 2T^{-1}\Sigma w_{sl}\Sigma k_t k_{t-s}$ for $s=1,...,l$, $t=s+1,...,T$, and some choice of lag window such that $w_{sl} = 1 - s/(l+1)$. If the Z_α statistic is smaller than the proper critical value (available in Phillips and Ouliaris (1990)), then the null of no-cointegration is rejected.

The Engle-Granger approach has the advantage of being simple but does not allow the existence of multiple equilibria and does not have a well-defined limiting distribution. The maximum likelihood cointegration approach developed by Johansen (1988) addresses these problems. Following Johansen and Juselius (1990), let $X_t = \pi_1 X_{t-1} + ... + \pi_k X_{t-k} + \mu + \Phi D_t + \epsilon_t$, where X_t is defined as before, D_t are three seasonal dummies, μ is an intercept, and ϵ_t are IINp$(0,\sigma)$. Defining L as the lag operator and Δ=1-L, the above model can be expressed as follows:

$$Z_{0t} = \Gamma Z_{1t} + \pi Z_{kt} + \epsilon_t, \tag{5}$$

where $Z_{0t} = \Delta X_t$, $Z_{1t} = (\Delta X_{t-1}, ... , \Delta X_{t-k+1}, D_t, 1)$, $Z_{kt} = X_{t-k}$, $\pi = -(I-\pi_1- ...-\pi_k)$, $\Gamma = (\Gamma_1, ...\Gamma_{k-1}, \Phi, \mu)$, and $\Gamma i = -(I-\pi_1-...-\pi_i)$. The coefficient matrix π contains information about the long run relationship(s) between the series in X_t and lends itself to hypothesis testing. A full rank π would suggest that X_t is a stationary system while a rank of zero would imply that equation (5) is simply a differenced vector time series model. In the intermediate case when $0 < r < p$, cointegration holds and π can be represented by the product $\alpha ß'$ where α and ß are pxr matrices. ß is made up of r cointegration vectors which satisfy equation (1). When r=1, the long run equilibrium is unique.

In order to build test statistics about the size of r, the residuals R_{0t} and R_{kt} from the regressions of Z_{0t} and Z_{kt} on Z_{1t} need to be saved. Alternatively, these residuals can be easily computed by using the moment's approach where $M_{ij}=T^{-1}\Sigma Z_{it}Z_{jt}'$, $(i,j)=0,1,k$ and T is the number of observations.

$$R_{0t} = Z_{0t} - M_{01}M_{11}^{-1}Z_{1t} \tag{6}$$

$$R_{kt} = Z_{kt} - M_{k1}M_{11}^{-1}Z_{1t} \tag{7}$$

From (6) and (7), the concentrated maximum likelihood can be expressed as $|\sigma|^{-T/2}\exp\{-\Sigma(R_{0t}-\pi R_{kt})'(R_{0t}-\pi R_{kt})/2\}$. Keeping in mind that $\pi=\alpha ß'$, the maximum likelihood estimator of ß can be obtained by maximizing the previous expression with respect to ß. It can be shown that this is equivalent to solving the following standard eigen value problem:

$$Lmax\text{-}2/T(H1) = |S_{00} - S_{0k\beta}(ß'S_{kk}ß)^{-1}ß'S_{k0}| = 0$$

$$= |\phi S_{kk} - S_{k0}S_{00}^{-1}S_{0k}| = 0 \quad (8)$$

$$= |\phi I - C^{-1}S_{k0}S_{00}^{-1}S_{0k}C'^{-1}| = 0, \quad (9)$$

where $S_{ij}=T^{-1}\Sigma R_{it}R_{jt}'$ for (i,j)=0,k, C is a pxp matrix such that $S_{kk}=CC'$, and ϕ is a vector of eigen values. Given the normalization of (8) into (9) and the fact that $S_{k0}S_{00}^{-1}S_{0k}$ is symmetric and positive definite, then the eigen values will be real and bounded by 0 and 1. Defining $e_1,...,e_p$ as the eigen vectors corresponding to (9) and $v_i=C^{-1}'e_i$ as the eigen vectors corresponding to (8), it follows that the estimated ß is given by $(v_1,...,v_r)$. Again, for r>1, there will be more than one long run equilibrium relationship.

Hypothesis testing about the number of cointegrating relations can be done using the likelihood ratio test. Defining H_1 as the hypothesis of r=p (i.e., no restriction on π) and $H_2(r)$ as $\pi=\alpha ß'$ for r < p, the trace test can be computed as:

$$-2\ln(Q;H_2|H_1)=-T\ \Sigma\ln(1-\phi_i), \quad (10)$$

for i=r+1,...,p and $\phi_1>...>\phi_p$. Testing $H_2(r)$ in $H_2(r+1)$ is also possible. In this case, the maximum eigen value test must be used.

$$-2\ln(Q;r|r+1)=-T\ln(1-\phi_{r+1}) \quad (11)$$

The distributions for (10) and (11) are multivariate versions of the Dickey-Fuller distribution and depend on the number p-r of non-stationary components under the hypothesis. The distributions are obtained by simulation and are given in Johansen and Juselius (1990).

The non-stationary process X_t may not have linear trends (i.e., $\mu=\alpha ß_0'$). If that was the case, the model specification used to test hypotheses regarding the cointegrating relations ß and speed of adjustment coefficients α would have to exclude linear trends. Linear trends can be excluded from the model by replacing X_{t-k} by $X^*_{t-k}=(X_{t-k},1)$ in the definition of Z_{kt} (see equation 5) and ß by $ß^*=(ß,ß_0)$ so that $\alpha ß'X_{t-k}+\mu = \alpha ß'X_{t-k}+\alpha_{\beta 0}' = \alpha^* ß^*{}'X^*_{t-k}$. The analysis can then be performed by following the steps given by equations (5) to (11). The test for the absence of linear trends is defined as:

$$-2\ln(Q;H_2^* \mid H_2) = -T\Sigma\ln\{(1-\phi_i^*)/(1-\phi_i)\}, \qquad (12)$$

where i goes from r+1 to p.

Johansen's approach would be of limited applicability if hypothesis testing about the cointegration vector(s) ß was not possible. Fortunately, this can be done quite easily. Let H_3 be the null hypothesis of $ß=H\delta$ where H is a pxs matrix of restrictions and δ is a set of cointegration vectors. Johansen and Juselius (1990) have shown that the following test can be used to verify hypotheses about the size of the individual components of ß.

$$-2\ln(Q;H_3 \mid H_2)=T\Sigma\ln\{(1-\phi_{3i})/(1-\phi_i)\} \qquad (13)$$

for i=1,...,r and where ϕ_{3i} are the eigen values associated with the solution to $|\phi H'S_{kk}H - H'S_{k0}S_{00}^{-1}S_{0k}H|=0$. Johansen (1989) has shown that the test statistic in (13) has a chi square distribution with r(p-s) degrees of freedom. Since r(p-s) corresponds to the number of restrictions on the system, it follows that the p-s restrictions being tested must hold for all r cointegrating relations. This feature greatly facilitates the interpretation of the testing of a theoretical hypothesis when the equilibrium is not unique.

The vector(s) α in $\pi=\alpha ß$ can be interpreted as a set of weights containing information about the speed at which variables adjust. Let H_4 be a hypothesis about the speed of adjustment of the variables of interest. H_4 can be expressed as $\alpha_{(pxr)} = A_{(pxm)}F_{(mxr)}$ or as $B'_{((p-m)xp)}\alpha_{(pxr)}=0$ where B is orthogonal to A (i.e.,$B=A^{\perp}$). Using equations (6) and (7) as well as the definition of S_{ij} which follows equation (9), we can define the following: $S_{ab}=A'S_{00}B$, $S_{bb}=B'S_{00}B$, $S_{bk}=B'S_{0k}$, $S_{ak}=A'S_{0k}$, $R_{at}=A'R_{0t}-S_{ab}S_{bb}^{-1}B'R_{0t}$, $R_{gt}=R_{kt}-S_{kb}S_{bb}^{-1}B'R_{0t}$ and $S_{ijb}=T^{-1}\Sigma R_i'R_j$ where i,j=a,g. It can be shown that the maximum likelihood estimator of ß under H_4 (say $ß_4$) can be obtained by solving:

$$|\phi S_{ggb} - S_{gab}S_{aab}^{-1}S_{agb}|=0, \qquad (14)$$

or $|\phi I - C^{-1} S_{gab} S_{aab}^{-1} S_{agb} C'^{-1}| = 0,$ (15)

where $CC' = S_{ggb}$. The solution to (14) yields a set of eigen values $\phi_{41} > ... > \phi_{4p}$ and eigen vectors $e_{41}, ...e_{4p}$. ß4 is computed by multiplying C'^{-1} by the eigen vectors e_{4i}'s. Given the above definition of H_4 and that $F = (A'A)^{-1} S_{agb} \beta_4$, the constrained weights α_4 are obtained by multiplying A and the estimated F. The restriction(s) implied by A are tested with the following likelihood ratio test:

$-2\ln(Q;H_4 | H_2) = T\, \Sigma \ln\{(1-\phi_{4i})/(1-\phi_i)\},$ (16)

where i ranges from 1 to r. The asymptotic distribution of this test is given by a chi square distribution with r(p-m) degrees of freedom. The p-m restrictions about the weights of adjustment are tested in all r cointegrating relations which means that a hypothesis is accepted if and only if it holds in every cointegration relation.

The three quarterly data series used for this analysis were taken from Agriculture Canada's MOAD data base. They are:

FIP - Farm input prices, total index for Canada, 1981=100,

FOP - Farm product prices, total index for Canada, 1981=100,

FRP - CPI for food sold in stores, for Canada, 1981=100.

The sample contains 112 observations, from the first quarter of 1961 to the last quarter of 1988. The reason for having all three series in an error correcting mechanism as opposed to having a system made up of FIP and FOP and another with FRP and FOP is that the influence of some input prices in FIP on FRP need not be all filtered through FOP. For example, rises in the price of electricity or gasoline are likely to have a direct impact on food retail prices. Hence the analysis was carried out with the three variables in order to avoid an obvious mispecification error.

Results

As stated earlier, the first step is to verify that the series can be made stationary by performing a single non-seasonal difference. Even though seasonality is often present in times series and has been fully incorporated to traditional time series analysis, it has received little attention in cointegration analysis. This situation has been remedied by recent contributions by Engle, Granger, Hylleberg, and Yoo (1987), EGHY thereafter, and Osborn, Chui, Smith, and Birchenhall (1988), OCSB thereafter. Testing for seasonal unit roots is important because there cannot be long run relationships between variables integrated of different orders. In other words, the left and right hand sides of error correcting models must have compatible long run behaviors.

Four procedures are implemented to verify the assumption that all three series are I(1,0). The first one is the Augmented Dickey-Fuller test (ADF). The ADF test is a t statistic for B_0 in the model $\Delta x_t=B_0x_{t-1}+B_1\ x_{t-1}+...+B_k\ x_{t-k}+e_t$. In this instance the null and alternative hypotheses are Ho:x_t is I(1,0) and Ha:x_t is I(0,0). The Dickey-Fuller (DF) test is similar to the ADF test with the exception of not containing lagged differences on the right hand side.

In the EGHY test, the null hypothesis is Ho:x_t is I(0,1) and it is tested by comparing three calculated t ratios (for π_1, π_2, and π_3) to generated critical values. The test is constructed as follows:

$\Delta_4x_t=\pi_1Z_{1,t-1}+\pi_2Z_{2,t-1}+\pi_3Z_{3,t-2}+a_1\Delta_4x_{t-1}+...+a_p\Delta_4x_{t-k}+e_t$

where $Z_{1t}=\phi(L)(1+L+L^2+L^3)x_t$, $Z_{2t}=-\phi(L)(1-L+L^2-L^3)x_t$, $Z_{3t}=-\phi(L)(1-L^2)x_t$, $\phi(L)x_t=(1-\phi_1L-...-\phi_kL^k)x_t$ and the ϕ_i's are obtained by regressing Δ_4x_t on $\Delta_4x_{t-1},...,\Delta_4x_{t-k}$. As stated in OCSB (1988), the alternative hypotheses are: i)H_a: x_t is I(1,0) when $\pi_1=0$, and π_2 or π_3 is non-zero, ii) H_a: x_t is I(0,0) when π_1 is non-zero and π_2 or π_3 are non-zero. In the implementation of this test k was set to four.

The OCSB test consists of comparing the t ratios for B_1 and B_2 to computed critical values in the following model: $\Delta\Delta_4x_t=B_1Z_{4,t-1}+B_2Z_{5,t-4}+a_1\Delta\Delta 4x_{t-1},...,a_k\Delta\Delta_4x_{t-k}+e_t$, where $Z_{4,t}=\phi(L)\Delta_4x_t$,

$Z_{5,t}=\phi(L)\Delta x_t$. In this test, the ϕi's are obtained from regressing $\Delta\Delta_4 x_t$ on $\Delta\Delta_4 x_{t-1},\ldots,\Delta\Delta_4 x_{t-k}$. According to OCSB (1988), under the I(1,1) null, B_1 provides a test of the non-seasonal unit root, while B_2 examines the unit root at the seasonal lag. It can be shown that with $B_1=0$, the t ratio on B_2 is a Dickey-Hasza-Fuller test of the seasonal unit root after differencing. As for the EGHY test, k=4 was used in the construction of the test.

Table 1 contains the results of the unit root tests. The DF and ADF tests indicate that FIP, FOP, and FRP are non-stationary in the levels but have stationary first differences. The EGHY test rejects the null that x_t is I(0,1) for all three series and clearly suggests that FIP and FRP are I(1,0). The test also indicates that FOP is I(0,0) or perhaps I(1,0) given that $\pi_1 = 0$ is rejected by a small margin. The results of the OCSB test unambiguously indicate that the series are indeed I(1,0). Hence it is concluded that FIP, FOP, and FRP are all integrated of degree one.

Having met the initial requirements, it is then proper to proceed with the second step of the Engle-Granger approach. Three regressions were run to determine the relationship between farm input prices, farm output prices, and retail prices. In all three cases, the R^2 was in the neighborhood of 0.99. The t ratios in Table 2 reveal that the explanatory variables were highly significant. The hypothesis that the sum of the estimated coefficients of the explanatory variables equal one is tested and the results are presented in Table 2. This test can be interpreted in two ways. From the first regression it could be established that $dFIP = B_1\,dFOP + B_2\,dFRP$. In the event of a general price increase defined as $dP=dFOP=dFRP$, it follows that $dFIP/dP = B_1 + B_2$. If $B_1 + B_2 = 1$, the FIP will increase by the amount dP. Since inflation is a controversial subject, it might be interesting to link the above hypothesis to this concept. Because the differences between the means of the three series are not statistically significant[3], P could be defined such that $FIP\approx FOP\approx FRP\approx P$. Therefore, in the event of a general price increase $dP=dFOP=dFRP$, $dFIP/FIP = (B_1 + B_2)(dP/P)$, where dP/P is the rate of inflation. In this case, $B_1 + B_2 = 1$ implies that FIP increases as fast as the rate of inflation.

In the FIP regression, the sum of B_1 and B_2 is significantly less than one and suggests that FIP would lag behind (but not by much) if there were a general price increase dP. The negative coefficients in the FOP and FRP equations are at first glance peculiar, especially when considering that the coefficient of correlation between FRP and FOP is 0.94.[4] The latter statistic clearly indicates that both variables trend in the same direction, increasing from one year to the next. The negative signs are simply the result of differences in the behavior of FRP and FOP around their respective trends[5]. The estimates obtained for the FOP regression support the neutrality of general price increases on farmers' terms of trade. In contrast the estimated coefficients and the F test for the food retail price equation upholds the "cost push" theory since FRP increase by an amount 1.06 times larger than the general price increase.

To establish that the relationship between FIP, FOP, and FRP is not spurious, it is imperative to verify that the price series are co-integrated.[6] The Durbin-Watson statistic can be used to for this purpose even though its critical value is sensitive to the particular parameters within the null (Engle and Granger, 1987). Since cointegration implies that the residuals from regressing non-stationary series on each other be stationary, it follows that the Durbin-Watson statistic must be significantly above zero (Bhargava, 1984). As argued by Granger and Newbold (1986) the standard tables for this statistic cannot be used when testing for cointegration since the test is not whether D.W.=2 but rather that D.W. is significantly positive. The Durbin Watson statistic for the first two regressions is higher than the critical value at the 5 percent level (0.367) while for the FRP regression it is well above the critical value at the 10 percent level (0.308). This preliminary evidence suggests that the null hypothesis of no cointegration can be rejected and that our three series "move together".

The ADF test for cointegration consists of regressing the first difference of the residuals of the regressions in Table 2 on lagged dependent variables and a lag residual. If the coefficient on the lag residual is statistically different from zero, then the residuals are stationary and the null hypothesis of no cointegration is rejected. According to Table 3, the null hypothesis of no cointegration is rejected at the 5 percent level (ADF above 3.13) for all three regressions. The Z_α tests computed for a lag window equal to 4 confirm the presence of cointegration at a 5 percent level of significance for the FOP equation and at a 10 percent level of significance for the two other equations.

Since it was established that FIP, FOP, and FRP are I(1) and cointegrated, there has to be a Granger causal relationship between the variables. The next step aims at analyzing the response of the three price series to deviations from the long run equilibrium and hence will provide information about the direction and nature of the causal relationships. For each series, variations of $(1\text{-}L)x_t = A^*(L)(1\text{-}L)X_t + \Theta R_{t-1} + D(L)\epsilon_t$ are estimated with OLS. A series of tests are then conducted to check for omitted variables such as lagged variables, time trends, and seasonal dummies and to verify that the ϵ_t's are well behaved. Time trends and seasonal dummies improved the fit of two of the error correcting models but did not significantly affect the estimated parameters and t ratios of $(1\text{-}L)X_t$ and R_{t-1}. Heteroscedasticity is tested with ARCH tests of various order, the Breuch-Pagan-Godfrey test and the Glejser test while the Durbin M and the Box-Pierce Q tests are used to verify the absence of autocorrelation. Normality is tested with the Jarque-Bera test. For all three series, the correlogram of the OLS residuals suggests the presence of third or fourth order correlation. Consequently, re-estimation is carried out with an iterative Cochrane-Orcutt procedure.

Table 4 shows the magnitude of the error correcting coefficients as well as their corresponding t ratios. As expected, the error correcting coefficients are all negative. This implies that the error correcting mechanism will eventually bring back a series to its long run equilibrium after a shock.

The low t ratio for FOP means that the parameter of the error correcting term is not statistically different from zero. The error correcting term for FRP is not significant at the 95 percent level but greatly contributes to the stability and robustness of the model. It follows that to avoid a mispecification problem, the error correcting term must be included.

The importance of the error correcting terms for FIP and FRP indicate that FOP and FRP Granger cause FIP and that FOP and FIP Granger cause FRP. As for FOP, the error correcting term is not significant but the first lag of the first difference of FIP is. Therefore, farm output prices are weakly exogenous from farm input prices and food retail prices in the sense that changes in farm output prices are not affected by the extent by which the three price series are out of equilibrium. The weak exogeneity of FOP implies that following an innovation, disequilibrium will persist until FRP and FIP adjust. The error correcting equation for FOP boils down to the standard equation used in Granger causality tests and confirms that FIP Granger causes FOP. The direction of the causal relationship is not surprising given that the prices of supply-managed commodities (which are included in the FOP index) are based on cost of production formula. Similarly, the weak exogeneity of FOP can be rationalized by the generally accepted notion that governmental agencies cannot react quickly enough to market disturbances to implement optimal state-contingent policies. Consequently, during the time farm input and food retail prices adjust to a shock, it is possible for the FPO/FPI ratio to fluctuate around its long run equilibrium.

The results from the maximum likelihood approach to cointegration are presented in Table 5. Both the trace test and maximum eigen value test, as defined in (10) and (11), suggest the presence of two cointegrating relationships. The relevant cointegrating vectors have the highest degree of correlation with the stationary part of the model and are given by the columns of ß's under the the two largest eigen values. The error correcting mechanisms implied by these vectors can be represented as follows:

$$\text{FRP} = 4.01\ \text{FIP} - 3.95\ \text{FOP}, \tag{17}$$

$$\text{and FIP} = 0.37\ \text{FOP} + 0.57\ \text{FRP}. \tag{18}$$

As explained earlier, the matrix α distributes the influence of the stationary error-correction variables $\beta' X_t$ to the components of X_t. Thus the weights (α's) can be interpreted as coefficients capturing the speed at which the variables adjust to equilibrium (Johansen and Juselius, 1990). A high coefficient indicates rapid adjustment. The first column of α's are the weight with which FRP enters all three equations. Similarly, the second column of α's can be regarded as the weight with which FIP enters all three equations. It should be noted that the weights of FIP and FRP in the FOP equation are particularly small which is consistent with the FOP error correcting model estimated in the context of the Engle-Granger approach.

The hypothesis about the absence of linear trend in the non-stationary part of the model is formally tested by using equation (12). As indicated by the likelihood ratio test in Table 6, the hypothesis is strongly rejected. As explained earlier, hypotheses about the coefficients of the explanatory variables in (17) and (18) can also be tested. The parameters of the two cointegration vectors are restricted such that their sum is zero. This restriction is consistent with the tests conducted on the cointegration regressions reported in Table 2 and therefore, should be interpreted in a similar manner. Using the following H matrix,

$$H = \begin{bmatrix} 1 & 1 \\ -1 & 0 \\ 0 & -1 \end{bmatrix} \tag{19}$$

and applying the test described by equation (13), a likelihood ratio of 87.09 is obtained and exceeds the critical value of 5.99 for a 5 percent level of significance. The null hypothesis is strongly rejected as it is for the FIP and FRP cointegration regressions of the Engle-Granger approach. It was shown earlier that because the means of FIP, FOP, and FRP do not significantly differ, testing null hypothesis associated with (19) is tantamount to testing the neutrality of inflation. Therefore, the result of our test rejects that inflation is neutral.

The last hypothesis being tested relates to the speed at which variables adjust following a shock. The unrestricted weights are displayed in Table 5 and it appears that the coefficients associated with the FOP equation (i.e., the second row of α's) may not be significantly different from zero. This hypothesis is tested with the procedure defined in equations (14)-(16). The matrices A and B necessary to implement the likelihood ratio test are as follows:

$$A = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}, \quad B = \begin{bmatrix} 0 \\ 0 \\ 1 \end{bmatrix}. \tag{20}$$

The computed likelihood ratio in Table 6 is 5.14 and is smaller than the critical value of 5.99 for a confidence level of 95 percent. Consequently, the null hypothesis that FIP and FRP have zero weights in the FOP cointegration equation (i.e., the second row of α's is a vector of zero in Table 5) is accepted. This result implies that farm output prices are not affected by the extent with which farm input prices and food retail prices are out of equilibrium. The acceptance of the null hypothesis confirms the weak exogeneity of FOP under the Engle-Granger approach reported in Table 4. In terms of modeling, the above result legitimizes the estimation of the α's and ß's with only two equations. The bottom half of Table 6 contains the eigen values, cointegration vectors, and weights associated with the restriction on the weights. The cointegration relations can be expressed as:

$$FRP = 0.27\ FIP + 1.58\ FOP \tag{21}$$

and $$FIP = 0.43\ FOP + 0.54\ FRP. \tag{22}$$

The absolute value of the coefficients within the columns of weights in Table 6 are almost identical. The coefficients of adjustment in first column are larger than the ones in the second column and show that food retail prices have a greater influence than farm input prices in the adjustment process for food retail price and farm input prices.

Implications of the Results

The single equation Engle-Granger cointegration approach and Johansen's maximum likelihood cointegration approach have indicated that Canadian farm input prices, farm output prices, and food retail prices are cointegrated. During the period covered by our analysis, these variables move together and should be expected to continue to move together as long as they remain integrated of degree one and the error terms in the cointegrating relations remain stationary. Barring the occurrence of events or adoption of policies that could induce the violation of these conditions, efficient forecasts of any of these variables can be obtained by using the two other variables of the system as instruments.

The results obtained in the previous section support the "Cost Push" and "Demand Pull" theories since disequilibrium at the retail level is transmitted to the input level and vice versa. However, since the weights on the cointegrating vectors are larger for FRP than FIP, it can be concluded that farm input prices and food retail prices adjust faster to disequilibrium in food retail prices than in farm input prices. This result gives a slight edge to the "Demand Pull" theory and is consistent with the conclusions reached by Sephton (1989), and Granger, Robbins, and Engle (1986).

Not surprisingly, it was found that farm output prices do not respond to changes in the magnitude by which the farm input prices and food retail prices are out of equilibrium. That is not to say that farm output prices do not adjust. It appears that the timing and magnitude of the adjustments of farm output prices are not consistent from one market disturbance to another. It is conjectured that this result is largely due to the importance of the prices of supply-managed commodities in the farm output price index.

The contention that farm output prices are more flexible than farm input prices and that farmers must gain from general price increases is clearly rejected. The weak exogeneity or inflexibility of farm output prices does not imply that farmers are worse off in the event of a general increase in prices. It is possible that the adjustment mechanism of farm output prices, though different from the adjustment mechanism of farm input prices and food retail prices, be such that it overshoots and undershoots price increases in other sectors brought about monetary or non-monetary shocks.

Conclusion

This study has used new time series procedures to analyze the behavior of farm input prices, farm output prices, and food retail prices in Canada. The Engle-Granger-Hylleberg-Yoo and Osborn-Chui-Smith-Birchenhall tests for non-seasonal and seasonal unit roots were implemented to show that all three series have a non-seasonal unit root. Cointegration was tested in four ways. The Augmented Dickey-Fuller test and the Phillips Z_α test were computed prior to the estimation of error-correcting models a la Engle-Granger. The third and fourth tests were the trace and maximum eigen value tests developed by Johansen (1988). Without exception, the tests revealed that the three variables were cointegrated. Hypothesis testing on the error correcting models estimated with the Engle-Granger approach and Johansen's maximum likelihood approach suggested that farm output prices were not affected by the extent by which the system was out of equilibrium.

The weak exogeneity of farm output prices implies that the two cointegrating vectors and their corresponding weights could be estimated with only two equations. Given the existence of multiple equilibria, Johansen's approach is more appropriate than the Engle-Granger cointegration approach and should be the choice of applied economists testing for cointegration and estimating error correcting models.

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Table 1. Unit Root Tests on Individual Series

TEST		Series FIP	FOP	FRP	C.V.[a] (5%)
AUGMENTED[b] DICKEY-FULLER		0.95	-0.22	2.36	-2.89
DICKEY-FULLER		0.91	-0.38	2.56	-3.37
DICKEY-FULLER (ON SECOND DIFFERENCE)		-19.58	-13.38	-19.96	-3.37
ENGLE-GRANGER HYLLEBERG-YOO	Π_1	-2.17	-3.54	-2.30	-2.96
	Π_2	-2.12	-3.47	-2.32	-2.95
	Π_3	-7.21	-5.29	-7.20	-3.51
OSBORN-CHUI-SMITH-BIRCHENHALL	β_1	-1.76	-0.94	-1.68	-1.83
	β_2	-3.89	-4.58	-3.61	-2.03

a The critical values were taken from OCSB (1988) except for the Dickey-Fuller test which were taken from Engle and Granger (1987).

b As in Sephton, the tests were also conducted with seasonal dummies. Because the results were very similar to the ones above, they are not reported.

Table 2. Engle-Granger Cointegration Regressions

DEPENDENT VARIABLE	EXPLANATORY VARIABLES	ESTIMATED COEFFICIENTS	t RATIOS	F[a]	D.W.[b]	D.W. C.V. 5%
FIP	FOP	0.37	18.65	15.03	0.41	0.37
	FRP	0.60	39.71			
FOP	FIP	2.06	18.65	0.13	0.38	0.37
	FRP	-1.06	-11.05			
FRP	FIP	1.56	39.71	27.43	0.35	0.37
	FOP	-0.50	-11.05			

a H_o: The coefficients of the explanatory variables sum to one. $F_{0.95,1,109} = 3.94$

b A high Durbin-Watson suggests that the residuals of the cointegrating regression are stationary. The critical value are from Granger and Newbold (1986, p. 264). The 10% critical value is 0.308.

Table 3. Tests of Cointegration

DEPENDENT VARIABLE IN COINTEGRATING REGRESSION	AUGMENTED DICKEY-FULLER	C.V.[2] (5%)	$Z\alpha$	C.V.[c] (5%)
FIP	-3.43	-3.13	-24.47	-26.09
FOP	-3.90	-3.13	-26.23	-26.09
FRP	-3.28[b]	-3.13	-22.65	-26.09

a Critical values for the ADF test for p=3 were taken from Granger and Newbold (1986 p.264).

b Unlike the ADF test of cointegration for FIP and FOP, the ADF test with seasonal dummies (-3.28) was significantly different than the ADF test without seasonal dummies (-2.79).

c As indicated by Phillips and Ouliaris (1990), the null hypothesis of no-cointegration is rejected if the computed value of the statistic is smaller than the appropriate critical value. The critical value for a 10% level of significance is -22.19 which is larger than all of the calculated $Z\alpha$'s.

Table 4. t Statistics for the Error Correcting Terms $\ominus$ in $(1\text{-}L)X_t = A^*(L)(1\text{-}L)X_t + \ominus R_{t\text{-}1} + D(L)\epsilon_t$

EQUATION	COEFFICIENT	t-RATIO[a]
FIP	-0.2470	-3.4722 (-2.09)
FOP	-0.0780	-1.0851 (-0.2184)
FRP	-0.0790	-1.4964 (-2.2025)

a t ratios in parentheses were obtained from the OLS regressions while the other ratios were obtained from regressions corrected for various order of autocorrelation.

Table 5. Maximum Likelihood Cointegration Results and Tests about the Number of Cointegrating Vectors

EIGEN VALUES	0.2124	0.1563	0.0057	
EIGEN VECTORS β's				
FIP	4.01	1.00	1.00	
FOP	3.95	-0.37	-0.15	
FRP	1.00	-0.57	-0.84	
WEIGHTS α's				
FIP	-0.42	0.08	0.05	
FOP	0.10	-0.08	0.18	
FRP	-0.27	-0.30	0.03	
HYPOTHESIZED NUMBER OF COINTEGRATING VECTORS	TRACE TEST	C.V.[a] 5%	MAX.EIGEN VALUE TEST	C.V.[a] 5%
$r = 0$	45.18	31.26	26.03	21.28
$r \leq 1$	19.15	17.84	18.53	14.60
$r \leq 2$	0.62	8.08	6.21	8.08

a The critical values were taken from Johansen and Juselius (1990).

Table 6. Hypothesis Testing with the Maximum Likelihood Approach to Cointegration and Final Parameter Estimates

NULL HYPOTHESIS	LIKELIHOOD RATIO	C.V. 5%
No trend in the non-stationary part of the process	321.26	3.84
The parameters in the cointegration vectors add up to zero	87.09	5.99
The speed of adjustment coefficient for FOP is zero	5.14	5.99
FINAL PARAMETER ESTIMATES		
Eigen Values	0.9124	0.1375
Eigen Vectors β's		
FIP	-0.27	1.00
FOP	-1.58	-0.43
FRP	1.00	-0.54
Weights α's		
FIP	0.36	0.21
FOP	0	0
FRP	0.37	-0.17

Endnotes

1 The expression "terms of trade" is not used in its usual trade theory sense. As in Starleaf, Meyers, and Womack (1985), terms of trade is defined as the ratio of prices received by farmers and prices paid by farmers.

2 Modern macroeconomy assumes that economic agents have enough time to adjust when inflation is expected. However, expected inflation may cease to be neutral if (1) information is not accessible to all to the same degree, (2) there are different expectation mechanisms (ability to process the information) and, (3) persistent rigidities.

3 An ANOVA showed that the means of the three series are not statistically different. The F test was 0.2824 and clearly suggest it is not true that at least one mean differ.

4 The coefficients of correlation between FRP and FIP and between FOP and FIP are 0.99 and 0.97 respectively.

5 This could be explained by the fact that the coverage of the food retail series covers imported and processed items in addition to unprocessed domestic items.

6 If they were not co-integrated, the series would be differenced to achieve stationarity and regressed in the context of a multivariate Box-Jenkins analysis.