Macroeconomic Determinants of Relative Wheat Prices: Integrating the Short Run and Long Run

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Abstract. *Prior empirical studies ignore that markets, subject to overshooting, determine farm prices and macroeconomic variables jointly So, these elasticities are statistically unreliable Using cointegration, with all variables determined simultaneously, we find that instantaneous wheat price elasticities with respect to the real exchange rate and interest rate are -1 27 and -1 97, respectively Here, we measure the amount that the wheat price overshoots its equilibrium The extent of overshooting differs for different monetary policy regimes However, 57 percent of the deviation from longrun equilibrium is corrected within two quarters*

Keywords. *Relative prices, real exchange rate, real rate of interest, cointegration, commodity prices, overshooting, and error-correction models*

Over the past decade, analysts have determined that macroeconomic developments have important effects on the agricultural economy through relative farm-to-nonfarm prices We will refer to such farm prices as relative farm prices The magnitudes of the elasticities of relative farm prices, however, with respect to such key macroeconomic variables as the exchange rate and the interest rate, are still contested—for two substantive economic reasons

First, theoretical work assessing the magnitude of the exchange-rate elasticity of the farm price has demonstrated that it is necessary to include all macroeconomic variables and treat them as endogenous in empirical models This result occurs because the range of theoretically admissible values of the exchange-rate elasticity of commodity prices expands as more macroeconomic variables are treated as endogenous In static single-market models with an exogenous exchange rate as the only macroeconomic variable, the theoretically derived elasticity of the commodity price with respect to the exchange rate is, inclusively, between -1 and 0 Orden (1986) shows, theoretically, that if the exchange rate and national income are included and treated as endogenous, this elasticity will not be restricted to values between -1 and 0 [1] With money demand depending on real income and a rapidly clearing money market, he shows that a change in the money stock induces a percentage change in the relative-farm-to-nontradeable-goods price, which may exceed the percentage change in the real exchange rate Chambers and Just (1986) stress more general models and show that in theory the admissible exchange-rate elasticities of agricultural prices may be even less restricted if interest rates were also to be endogenized They argue that making interest rates endogenous will allow a model to account for the dumping of grain stocks on world markets in response to tightening Federal Reserve policy After the above discussions, it was clear that to estimate correctly the elasticity of a farm price with respect to a macroeconomic variable, all macroeconomic variables need to be included and treated as endogenous However, the practical difficulty of estimating such a large econometric system has been overwhelming

Second, to estimate properly the relative farm price elasticities with respect to macroeconomic variables, the dynamics of relative farm prices must be accounted for This is because the magnitudes of these elasticities are affected by the atypical shortrun reaction of relative agricultural prices to changes in monetary policy That is, in the short run, relative farm prices react to monetary policy by more than they do in the long run

The atypical relative price dynamics is caused by what Dornbusch (1976) defined as overshooting Overshooting is a more-than-proportionate shortrun response of a nominal asset price, such as a farm commodity price, to a change in money growth The shortrun rigidity of manufacturing and service prices ensures this disproportionate response Because of this general price rigidity, a change in nominal money supply affects the real money supply, which, in turn, influences the real interest and exchange rates in the short run by more than required in the long run The endogenous shortrun reactions of real interest rates and exchange rates induce the more-than-proportionate reactions of flexible asset prices, such as farm commodity prices [2] The specific

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[1]Sources are listed in the References section at the end of this article

[2]Unlike prices for manufactured goods and services, the prices for farm commodities as well as financial assets are determined in auction markets and are thus highly flexible in the short run (Okun, 1981)

mechanisms by which monetary policy influences shortrun nominal farm price dynamics differs for open and closed economies For example, in an open-economy model, a dynamic farm price adjustment is caused by the farm-export-demand effects of the real exchange rate response to a change in monetary policy (Stamoulis and Rausser, 1988) While, in a closed-economy model, a dynamic farm price adjustment is caused by crop-inventory-demand effects of the real interest rate response to a change in monetary policy (Frankel, 1986) [3]

Consequently, overshooting could substantially distort the relative farm prices in the short run, influencing the magnitudes of their elasticities with respect to macroeconomic variables Thus, any attempt to measure the relative farm price elasticities must take into account the atypical relative farm price dynamics, which has not been done before

The objective of this analysis is to estimate the macroeconomic elasticities of the relative wheat price, measuring the magnitudes of shortrun deviations from longrun equilibrium and the speed with which the relative price approaches its longrun equilibrium level To this end, the relative wheat price is modeled using cointegration methodology that joins, in an econometrically acceptable manner, the longrun trend relationship between the relative price and its determinants, including macroeconomic variables, into a shortrun dynamic equation The dynamic equation, referred to as an error-correction model, identifies how the rate of growth of the relative price responds to its shortrun deviations and to changes in the rates of growth of its longrun determinants Thus, by accounting for the dynamics of price adjustments and treating all variables as endogenous, the new methodology resolves the two difficulties in estimating the elasticities of relative prices with respect to macroeconomic variables

Cointegration and Error-Correction Models

Advances in cointegration by Engle and Granger (1987) and Johansen and Juselius (1990) provide the tools to apply dynamic error-correction models (ECM's), first suggested by Sargan (1964), that account explicitly for the dynamics of shortrun price adjustment toward longrun equilibrium When variables in an equation are nonstationary, spurious regression results are likely, in which case correlated stochastic trends result in a high R^2, and nonstationary residuals produce a low Durbin-Watson statistic The usual solution to achieving stationarity is to estimate the model in first differences However, this first-differencing typically results in a loss of information concerning the long-term relationship between the variables

Cointegration analysis resolves this problem by identifying conditions under which a relationship is robust (Engle and Granger, 1987) If variables are cointegrated, longrun trends (secular components) of time series variables adjust in accordance with an equilibrium constraint, and the shortrun dynamics (cyclical components) conform to the class of ECM's That is, while stochastic trends cause the variables to wander apparently randomly, the time series variables eventually follow one another if they are cointegrated In this way, cointegration and error-correction modeling reintroduce, in a statistically acceptable way, the longrun information omitted from the differenced models

Consider the simple case of two endogenous time-series variables, x_t and z_t, with single-unit roots whose first differences are stationary [4] The linear combination, referred to as the cointegrating equation

$$w_t = x_t - a - bz_t, \qquad (1)$$

is generally I(1), where a and b are constants However, if there exists an a and b such that w_t is level stationary, I(0), then x_t and z_t are said to be cointegrated, and the relationship

$$x_t - a - bz_t = 0, \qquad (2)$$

is the cointegrating or equilibrium relationship with w_t representing the equilibrium error When cointegrated, the shortrun dynamic processes through which the series adjust toward their longrun equilibria are represented by constrained ECM's The ECM's specify that the first differences of x_t and z_t are functions of distributed lags of first differences of both variables as well as the once-lagged equilibrium error, w_{t-1}, referred to as the error-correction term (ECT) Because the series are

[3]Although the overshooting literature emphasizes the implications of nominal price dynamics, ' the importance is not the overshooting result *per se* but the possibility that relative prices of farm products can be affected by monetary policy' (Stamoulis and Rausser, 1988, p 185) Agricultural production and real farm income are strongly influenced by relative farm prices Monetary policy, via commodity overshooting, affects relative farm prices Thus, in the short run, monetary policy influences real farm income and agricultural production Tight monetary policy is an implicit tax on farm production and farm income

[4]A variable is integrated of order d, I(d) if its dth difference is a stationary, invertible, nondeterministic ARMA process A variable integrated of degree zero is therefore stationary in its level

cointegrated, the ECT is stationary, matching the I(0) first differences Hence, the least squares standard errors of the ECM, using the ordinary least squares residuals of equation 1 in place of the ECT, will be consistent estimates of the true standard errors (Engle and Granger, 1987, p 262)

In the bivariate case, the cointegrating vector, [1, −b], must be unique since any other parameter, say (b+c), introduces the additional nonstationary term, cz_t When more than two variables are involved, the cointegrating vector may not be unique Engle and Granger's two-step procedure assumes a unique cointegrating vector So, their cointegration test does not distinguish between the existence of one or more cointegrating vectors Johansen and Juselius (1990) provide a maximum likelihood procedure to estimate the parameters of and to test for the number of cointegrating vectors

The modeling of macroeconomic and relative farm price variables is a natural application of cointegration, since the overshooting literature demonstrates that relative agricultural prices exhibit shortrun departures from longrun equilibrium Cointegration analysis determines the longrun relationships between the observed values of the relative wheat price and the other time series involved, where the residuals measure the extent of disequilibria And, the ECM describes the shortrun dynamic adjustment of the relative wheat price

Empirical Results

The solution to a typical static general equilibrium model specifies that the relative wheat price, P, depends on the real exchange rate, Q, real domestic income, Y, real foreign income, Y*, real interest rate, R, and the wheat stocks carried over from the last period, S (see app I) Assuming a log-linear function, the relative wheat price model is

$$lnP = k + alnQ + blnY + clnY^* + dlnR + elnS, \quad (3)$$

where k, a, b, c, d, and e are constant parameters

Cointegration and error-correction modeling involves three steps First, the order of integration for each variable is determined If a series is nonstationary, it will be successively differenced until stationarity is obtained Second, if nonstationary variables are integrated of the same order, a linear combination of them can be stationary The Johansen-Juselius procedure tests for cointegration, identifying the number of cointegrating vectors Finally, if the cointegrating vector is unique, the OLS residuals from equation 3 can be used to measure the equilibrium error, ECT, to proceed with the estimation of the dynamic ECM

Data

The data are quarterly and cover the 1977 4-1989 4 period The relative wheat price is measured by the ratio of the seasonally adjusted (fourth difference) wheat (Chicago no 2 soft red winter) price to the Nonfood Consumer Price Index The real exchange rate is a wheat-trade-weighted index of the real value of the U S dollar U S disposable personal income (constant 1982 dollars) represents the real domestic income The index of OECD's quarterly industrial production is a proxy for income of major U S wheat importers—a series which is not available The real interest rate is calculated by subtracting the rate of inflation (measured using the Consumer Price Index) over a quarter from the prime rate at the beginning of the quarter Beginning stocks are the deseasonalized total wheat inventory measured over noncalendar quarters, for example, December-February Because deseasonalizing prices and inventories removed the overall mean, all other series were also expressed as deviations from their means

Integrating Properties of the Variables

Unit-root test procedures developed by Fuller (1976) and Dickey and Fuller (1981) are applied to examine the orders of integration The procedure starts with the following regression

$$\Delta z_t = \alpha + \beta t + (\rho-1) z_{t-1} + \sum_{i=1}^{m} \rho_i \Delta z_{t-i} + e_t, \quad (4)$$

where z is the variable under consideration, Δz_{t-i} is the first difference at time t−i, and m is the number of lags that ensures adequate representation of the time series z, that is, when the error term, e_t, is white noise The null hypothesis for a unit root requires that $\rho=1$, in which case the variable z is said to be nonstationary The statistic used for the test, named τ_τ, is the usual t-statistic calculated under the hypothesized null However, the τ_τ statistic is not distributed as the standard t Fuller provides the critical values for the τ_τ distribution

If a unit root is detected, it is possible that a second unit root exists as equation 4 has m characteristic roots In this case, application of the same procedure to the first difference of a variable tests for possible existence of a second unit root Because the variable of interest is the first difference of z, model (4) without the deterministic time trend and drift is estimated Once again, the statistic used for the test, named τ, is the usual t-statistic calculated under the hypothesized null whose critical values are reported in Fuller If a second unit root is found, the procedure will be continued until the order of integration, that is, the appropriate number of differencing to achieve stationarity, is identified

To determine the order of autoregression, m, the Akaike (1977) information criterion was applied, which indicated that variables in equation 3 are generated by AR(1) processes [5] Other tests for additional lag terms indicated that AR(1) was sufficient to represent these processes [6]

The outcome of the tests are similar for all series (table 1) The null hypothesis of a unit root could not be rejected at the 10-percent significance level The results using first-differenced data unanimously rejected the hypothesis of second-unit roots So, each series is characterized as a nonstationary I(1) process

Cointegration Test

Because all variables are I(1), one or more linear combinations of these series could be stable in the long run if they are cointegrated To test for cointegration, the Johansen-Juselius maximum likelihood procedure is applied The procedure involves the identification of rank of the matrix Π in

$$X_t = \varphi + \sum_{i=1}^{k} \Pi_i X_{t-i} + e_t, \tag{5}$$

where X_t is a column vector made up of p (here six) series involved in the analysis The procedure is based on the error-correction version of equation 5 [7]

$$\Delta X_t = \varphi + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + e_t, \tag{6}$$

where $\Gamma_i = -[I - \Pi_1 - \quad - \Pi_i]$ for $i = 1, \quad ,k-1$, and, $\Pi = -[I - \Pi_1 - \quad - \Pi_k]$ Johansen and Juselius show

[5]These findings are also consistent with behaviors of autocorrelation and partial autocorrelation functions of the variables

[6]The test is based on Fuller's proof (1976, chap 8) that while the limit distributions of OLS estimators of α, β, and ρ are not normal, the distribution of such estimators for ρ_i's converge in the limit to a multivariate normal Consequently, an ordinary t-test can be used to test for the possible existence of an additional lag

[7]Equation 6 is derived from equation 5 Any autoregressive time series of order k can be written in terms of its first difference, its level lagged k times, and k-1 first differences (Dickey and others, 1991)

Table 1—Unit-root tests, I(1) and I(2)

Variable[1]	Levels, H_o I(1)[2]	Differences, H_o I(2)[3]
Relative wheat price	-1 92	-5 13
Real exchange rate	-1 94	-5 18
U S disposable income	-1 42	-4 68
OECD industrial production	-1 76	-4 55
Real interest rate	-1 67	-6 90
Wheat inventory	-2 76	-7 44

[1]All variables are in logarithm

[2]Critical values τ_τ for n = 50 are -3 18 and -4 15 at 10- and 1-percent significance levels, respectively

[3]Critical values of τ for n = 50 are -1 61 and -2 66 at 10- and 1-percent significance levels respectively

that if the rank is zero, the variables are not cointegrated However, if the rank is r, there exist r possible independent stationary linear combinations In the latter case, equation 6 represents an ECM described by Engle and Granger

The tests to determine the rank of Π involve the estimates of the ordered eigenvalues, $\lambda_1 > \quad > \lambda_p$, from the characteristic equation

$$|\lambda S_{kk} - S_{ko}(S_{oo})^{-1} S_{ok}| = 0,$$

where $S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it} R_{jt}'$ for $i, j = 0,k$, and T is the sample size The R_{ot} and R_{kt} are the OLS residuals obtained by regressing ΔX_t and X_{t-k} on an intercept and $\Delta X_{t-1}, \quad ,\Delta X_{t-k+1}$, respectively First, test that the rank of Π is less than or equal to one, that is, H_o $r \leq 1$ The likelihood ratio statistic, called the trace, is given by

$$-2ln(Q) = - T \sum_{i=2}^{p} ln(1-\lambda_i)$$

If the null hypothesis is not rejected, the hypothesis that the rank of Π is zero should be tested, or H_o $r = 0$ The trace statistic for this test is

$$-2ln(Q) = - T \sum_{i=1}^{p} ln(1-\lambda_i)$$

If the null is not rejected, then the rank is zero and the series are not cointegrated Otherwise, one would conclude that a unique cointegrating vector exists

An additional statistic, called the maximal eigenvalue statistic, provides evidence that should confirm the inference obtained by the trace statistics For example, given that $r \leq 1$, the maximal statistic for the null hypothesis that the rank is zero is

$$-2ln(Q, r=0|r\leq1) = - T\ ln(1-\lambda_1)$$

Similarly, if the trace statistics cannot reject the hypothesis that $r \leq 2$, then the result that $r=1$ can be confirmed by the maximal statistic

$$-2ln(Q,\ r=1|r\leq 2) = -\ T\ ln(1-\lambda_2)$$

The distributions of these statistics are not the usual chi-square Johansen and Juselius provide the asymptotic critical values

The lag structure of equation 6 must be determined to conduct the test One lag proved to be sufficient using the Akaike information criterion The trace statistic for the null hypothesis that $r \leq 1$ was 62 4, indicating that the hypothesis of at most one cointegrating vector cannot be rejected at the 10-percent significance level Because the dimensions of the distribution tables in Johansen and Juselius are limited to five series, the trace and maximal tests for $r = 0$ could not be performed Instead, the trace test was used for the hypothesis that $r \leq 2$ At 34 64, the null could not be rejected at the 20-percent level Having accepted this null, we used the maximal statistic to test that $r = 1$ against the alternative that $r = 2$ At 21 84, the statistic could not, at the 50-percent significance level, reject the null that there exists a unique cointegrating vector

Error-Correction Model

Engle and Granger proved that cointegration implies an ECM Since the variables in equation 3 are cointegrated, the shortrun dynamics of the relative wheat price follows an ECM that relates its growth rate to its past deviations from longrun equilibrium, that is, the ECT, and to the growth rates of the other variables (see appendix II) Uniqueness of the cointegrating vector means that the estimated residual of the cointegrating equation represents the equilibrium error

A major decision is the choice of lag length Because of the complexity of dynamic relationships, the orders of autoregressive-distributed lag (ADL) structure of ECM's may be complicated (Engle and Granger, 1987)[8] To find the lag lengths, Hendry's general-to-specific modeling strategy is followed, which estimates an unrestricted ADL version of the model first and, then, simplifies the representation by eliminating the lags with insignificant parameters[9] Since the data are quarterly, four lags of each variable were included initially However, because of high correlation (0 95) between the logarithms of U S disposable personal income, y, and OECD industrial production, y*, the X matrix was singular Only current y* could be included for the matrix to be invertible In addition, lags 2-4 were insignificant for all other variables Subsequently, the analysis of lag structure was performed for four lags of y, current y*, and one lag of all other series Based on these results and tests on the significance of each variable and each lag, the basic model was obtained by eliminating y*, the lagged dependent variable, all but lags 3 and 4 of y, the first lag of the ECT, and the constant term

The final stage is to transform the basic equation such that all variables are I(0), and so that the standard inference procedure applies to all tests As Hendry (1989) points out, doing so results in a nearly orthogonalized specification of the ADL The earlier unit root tests established that all time series are AR(1), so that their first differences are I(0) (table 2)

All estimated coefficients are statistically significant and have the expected theoretical signs (see appendices I and II) A battery of tests are used to validate the model As far as the residuals go, the DW statistic provides no evidence of serial autocorrelation, the LM test supports a white noise process, and the Jarque-Bera test indicates an approximately normal distribution The RESET and White tests provide no evidence of heteroscedastic misspecification[10]

Table 2—The error-correction model

Variable	Coefficient	Standard error	t-value
Δq	–1 27	0 407	–3 11
Δy_{t-3}	1 70	863	1 97
Δr	–1 97	593	–3 33
Δs	– 32	075	–4 22
ECT_{t-2}	– 57	088	–6 39

R^2 = 62 σ = 066 F(5, 44) = 14 52	DW = 2 32
Jarque-Bera test of normality	Chi²(2) = 1 37
LM test of 4th order autoregressive errors	F[4, 40] = 1 59
White's test of heteroscedastic errors	F[10, 33] = 68
RESET specification test	F[1, 43] = 2 22

The dependent variable is Δp Δ is the first difference operator p, q, y, r, and s are the logarithms of the relative wheat price real exchange rate, U S disposable income real interest rate, and beginning inventory respectively The sample period is 1977 4-1989 4

[8]In a money demand study, Hendry and Ericsson (1991), for example, estimated an ECM which includes nonlinear ECT's, first differences, second differences of lagged levels, and the rate of growth over the past two periods

[9]Hendry's software package, PC-GIVE, is used to estimate the ECM

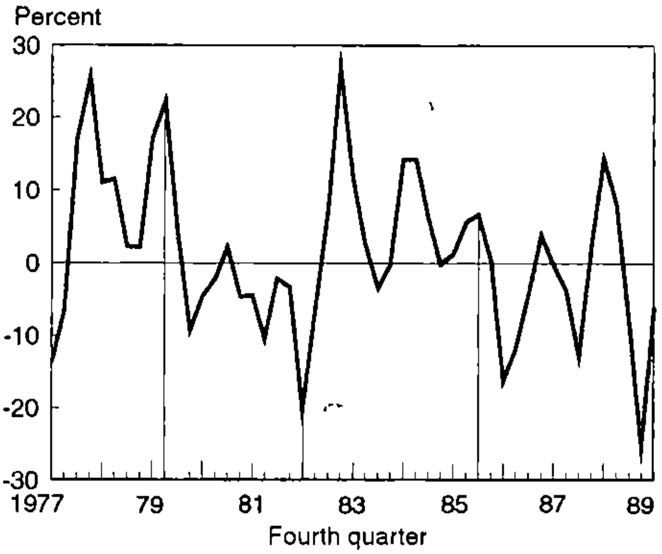
[10]An extensive battery of parameter constancy tests using recursive estimation for the out-of-sample period 1983 3 1989 4 is also conducted Chow tests, recursively estimated parameter values, and residuals along with their standard errors strongly suggest that the parameters are constant The results are available on request

The estimated ECM quantifies the effects of macroeconomic shocks characterized as significant upon relative agricultural prices (for example, Rausser and others, 1986) Consistent with Orden's theoretical result, the elasticity of the relative wheat price with respect to the real exchange rate exceeds unity The significant negative price influence of real appreciation of the dollar through its effect on the wheat export demand is even more profound if the exchange rate itself overshoots its equilibrium in response to a monetary shock

The relative wheat price is even more elastic with respect to the real interest rate The statistical significance and magnitude of the elasticity confirm the theoretical expectation that interest rate movements have important negative effects on current commodity prices via their influence on the demand for inventory (for example, Frankel (1986), Chambers and Just (1986), and Gardner (1979))

While the coefficients indicate large immediate responses to changes in the dollar's value and the interest rate, the negative coefficient of the ECT ensures, consistent with overshooting, that longrun equilibrium is achieved The adjustment toward equilibrium is not instantaneous, however Fifty-seven percent of any quarter's deviation from equilibrium is incorporated into the next two quarters' growth rate of the relative wheat price The direction of departures from equilibrium reported (fig 1) is also consistent with the conclusions of the overshooting analysis (for example, Stamoulis and Rausser, 1988) During the accommodative monetary policy of the late 1970's the relative wheat price was usually above its equilibrium During the tight monetary policy and high budget deficits of the early 1980's, the relative wheat price was usually below its equilibrium values When the Federal Reserve began to ease monetary policy in the fourth quarter of 1982, the relative wheat price rose above its equilibrium level The Federal Reserve pursued a relatively tight monetary policy between 1986 and 1989, and the relative wheat price was below its equilibrium for much of that period

Figure 1
Relative wheat price departure from equilibrium



Over the sample period, the magnitude of the deviation from equilibrium has been large at times, reaching 28 percent in absolute value The relevant policy question is Should there be an agricultural policy response to such large shortrun relative price departures? The present analysis does not provide a clear-cut answer to this question As our analysis demonstrates, 57 percent of any shortrun departure is corrected for in the following two periods If a monetary shock, for example, is temporary, then no agricultural policy action is called for Just and Rausser (1984) have discussed, however, alternative agricultural public policy options for situations under which continued adverse macroeconomic conditions cause relative farm prices to fall below their longrun equilibrium for extended periods of time

Conclusions

There is no question that macroeconomic developments alter the economic well-being of farmers The theory tells us that changes in macroeconomic policy produce real economic consequences for the agricultural sector through generating an atypical relative farm price dynamic The theory also tells us that the relative farm price impact, for example, is carried through the real exchange rate and the real interest rate In particular, because of the general price level rigidity, the relative farm price overshoots its longrun equilibrium level in the short run

But, how significant are the macroeconomic effects? No one knew Any empirical assessment of the above theory depends on the ability to join the shortrun and the longrun dynamics to measure the size and duration of the relative-price overshooting, as well as to estimate the macroeconomic elasticities of relative farm prices If the deviations from longrun equilibrium are small, the economic effects will be insignificant no matter how long the duration If the size of the overshooting is large, then the economic effects will be significant, especially if the duration is long Here, the significance of these macroeconomic impacts for the wheat market is measured The largest overshooting happened in 1983 3 when the relative wheat price overshot its equilibrium by 28 percent during a period of accommodative monetary policy Almost 60 percent of a departure from equilibrium in any quarter is incorporated into the growth rate of the relative wheat price in the following two quarters

Our empirical study reveals the extent by which monetary policy can affect the relative wheat price in the short run, particularly through its effect on the real exchange rate and the real interest rate During the periods of expansionary monetary policy, the wheat price rises relative to its equilibrium level Specifically, the relative wheat price immediately increases by 1 27 percent to a 1-percent depreciation in the real value of the dollar and by 1 97 percent to a 1-percent decline in the real interest rate This means that expansionary monetary policy disproportionately benefits wheat producers, relative to noncommodity sectors, in the short run as relative wheat prices overshoot and real interest rates decline Conversely, tight monetary policy hurts wheat producers in the short run

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Appendix I—The Canonical Static Model

Consider an economy that consumes two distinct types of goods tradeable commodities, which are internationally arbitraged, and nontradeable goods with sticky prices. Excluding information or transportation costs and assuming no trade barriers, the law of one price (LOP) applies to commodities at any time. Under such conditions, the real exchange rate, defined as the deviation from purchasing power parity, is determined by the relative international price of nontradeables (Dornbusch, 1985). Stated in real terms, the LOP can be expressed as

$$\left(\frac{PC^*}{PN^*}\right) = \left(\frac{E \times PN}{PN^*}\right)\left(\frac{PC}{PN}\right), \tag{A1}$$

where "*" denotes foreign variables, PC and PN are the domestic currency prices of the commodities and nontradeables, respectively, and E is defined as the domestic currency price in world money. Denoting the real exchange rate with "Q," and domestic and foreign relative commodity prices with "P" and "P*," respectively, the equilibrium condition (A1) is restated as

$$P^* = Q \times P \tag{A1.1}$$

Let export demand, X, be represented with the following function:

$$X = x(PC^*, PN^*, YN^*),$$

$$\frac{\partial X}{\partial PC^*}<0, \ \frac{\partial X}{\partial PN^*} \text{ and } \frac{\partial X}{\partial YN^*}>0, \tag{A2}$$

where YN* is nominal foreign income. The function x is homogeneous of degree zero in nominal prices and income. So, (A2) can be rewritten as

$$X = x(P^*, Y^*), \ \frac{\partial X}{\partial P^*} <0 \text{ and } \frac{\partial X}{\partial Y^*} >0, \tag{A2.1}$$

where Y*=(YN*/PN*) is real foreign income and P*=(PC*/PN*). Substituting from (A1.1) into (A2.1), we have

$$X = x(Q \times P, Y^*) \tag{A2.2}$$

Similarly, domestic demand, D, is given by

$$D = d(P, Y), \ \frac{\partial D}{\partial P} <0 \text{ and } \frac{\partial D}{\partial Y} >0, \tag{A3}$$

where Y is real domestic income, and, as before, P is the domestic price of tradeable commodities relative to nontradeable goods.

Finally, allow the inventory demand, I, to be described by

$$I = \iota(R), \ \frac{\partial I}{\partial R} <0, \tag{A4}$$

where R is real rate of interest.

Given this structure, relative price is determined by the equilibrium condition that markets clear in the short run. Because agriculture is the focus of this analysis, in the short run, that is from quarter to quarter or month to month, expected production is assumed constant. Therefore, intrayear supply for a given period is the total stocks carried over from the last period. That is,

$$X + D + I = S, \tag{A5}$$

where S is the predetermined current supply.

Now, substituting for X, D, I from (A2.2), (A3), and (A4) into (A5) and solving for domestic relative price yields

$$P = f(Q, Y, Y^*, R, S), \tag{A6}$$

where P is the equilibrium level of current relative price. Comparative statics show readily that, given the assumptions made so far about the signs of the partial derivatives, we must *a priori* expect to have

$$\frac{dP}{dQ} <0, \ \frac{dP}{dY^*}>0, \ \frac{dP}{dY} >0, \ \frac{dP}{dR} <0, \text{ and } \frac{dP}{dS} <0$$

Appendix II—Shortrun Dynamics of an Error-Correction Model

The first discussion of ECM's appeared in Sargan (1964), before Engle and Granger developed the concept of cointegration. ECM's are built around the notion that available data summarize the forces involved in a dynamic process of convergence toward longrun equilibrium values. As Engle and Granger show, if an I(1) vector of economic variables is generated by an ECM, the series must necessarily be cointegrated. In other words, as in the context of cointegration, ECM's define longrun equilibrium as a stationary linear relationship similar to equation 2. However, Sargan motivated ECM's by defining longrun equilibrium as in the steady state. In the context of our price model, equation A6, the steady-state equilibrium would be

$$P = K\ Q^a\ Y^b\ Y^{*c}\ R^d\ S^e \tag{A7}$$

Equation A7, which represents the stable longrun relationship, is linear in the logarithms of the variables, that is

$$p = \nu_1 + z\nu, \tag{A8}$$

where p is the logarithm of relative wheat price, ν_1 is the logarithm of the intercept K in equation A7, z is the logarithm of the row vector containing the determinants of relative wheat price, and ν is the column vector [a, b, c, d, e,]' To allow convergence to longrun equilibrium, some sort of shortrun dynamics is needed To illustrate the mechanics of convergence, assume the simplest case of an AR(1) type process

$$p_t = \alpha p_{t-1} + \mu + z_t\theta + \xi, \tag{A9}$$

where $|\alpha| < 1$, μ is the intercept, θ is a column vector of shortrun price elasticities, and ξ is a serially uncorrelated error term with a constant variance and zero mean

Given the shortrun dynamic model A9, the steady-state solution can be obtained when longrun equilibrium is defined as a dynamic steady state, in which all equilibrium values grow at a constant rate To see this, rearrange A9 by subtracting p_{t-1} from both sides, and adding and subtracting $z_{t-1}\theta$ from the right-hand side to obtain

$$g_p = \mu + g_z\theta + (\alpha-1)\,[p_{t-1} - z_{t-1}(1-\alpha)^{-1}\theta] + \xi, \tag{A10}$$

where g_p is the growth rate of the relative wheat price, and g_z is the row vector of growth rates of the variables in z The term inside the brackets in equation A10 provides the error-correction mechanism If the demand, p, rises above its longrun equilibrium level at time t–1, the term in the brackets becomes positive However, because $(\alpha-1)$ is negative, its effect at time t is to reduce the growth rate of the observed p toward its steady-state path For this reason, equation A10 is referred to as an error-correction model