# POLICY IMPLICATIONS OF UNIT ROOT NONSTATIONARITY IN MULTIPRODUCT ACREAGE RESPONSE SYSTEMS

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## ABSTRACT

This paper explores the theoretical and empirical issues in unit root non-stationarity for multiproduct acreage response systems. The policy implication of a unit root in a multivariate autoregressive time series like a system of acreage responses is very disturbing; it implies that even a one-shot government policy that influences acreage will have a permanent response associated with it. A wheat/barley acreage response system for the Prairie province region of Canada is estimated and the existence of a unit root cannot be rejected empirically. The response of the stationary system of acreage responses to a policy shock is shown to be radically different from the response of a non-stationary system of acreage responses even though the empirical estimates of the reduced form parameters are only marginally different between the two models.

## 1. INTRODUCTION

It is well-known that the economic implications of an economic time series that contains a unit root are radically different from that of a stationary process (Nelson and Plosser). In a time series containing a unit root, the response to any shock in the system can be decomposed into "permanent" and "transitory" components. That is, any shock in the system will cause a permanent response. In contrast, a stationary time series contains only a transitory component. In this paper, we use the concepts of permanent and transitory components to argue that a system of acreage responses for agricultural crops can contain a unit root. The policy implication associated with the existence of a unit root in a system of acreage response functions is very disturbing: it implies that even a one-shot government policy that influences acreage will have a permanent response associated with it. Therefore, the response to government policies is not simply a "policy on", "policy off" choice, as is the underlying assumption of static models. Once a government policy is put into place, there will be a permanent response, even if it is only a temporary policy. Further, the policies put in place today may just be a reaction to permanent responses created by policy "mistakes" made in the past. This argument is quite similar to Friedman's regarding the impact of changes in the money supply on aggregate output. If the time series generating a system of acreage responses contains a unit root, then the response of the system to even an unanticipated policy shock will <u>never</u> return to the pre-shock level without an equal shock in the opposite direction.

The gravity of the policy implications when unit roots are present in acreage response systems are as wide-ranging as they are disturbing. For example, if systems of acreage responses in the United States and the European Community have unit roots, it implies that even a successful conclusion to the latest round of the GATT negotiations in eliminating the price war between these two regions may not be able to undo the damage created by the agricultural policies of governments in the two regions. Or, given the predisposition of governments to provide output-enhancing subsidies to agriculture, large misallocations of resources can be permanently embedded in markets. The effects of these domestic policies can in turn be passed to the world market so that the permanent misallocation of resources may spread to all trading nations. Retaliatory measures introduced by governments to mitigate the permanent effects of policies instituted by other governments may only compound the permanent responses in the system.

The purpose of this paper is to theoretically motivate and then test for the existence of a unit root in a vector-autoregression (VAR) model of supply responses. Both our theoretical and empirical models are applied to a multiproduct system of acreage responses. Treating acreage response in a multiproduct system is preferable to treating it in the more traditional way using single-product relationships in that the former explicitly recognizes the interrelationship among the acreages of different crops.

## 2. THEORETICAL CONSIDERATIONS

In this section, we develop two alternative but observationally equivalent theoretical structures that can generate a linear dynamic system of acreage response functions: 1) a Nerlovian partial adjustment model and 2) a linear-quadratic optimal control problem with costly adjustment. In doing so, our interest is not to empirically distinguish between the two, but to motivate the multiproduct system of equations on the basis of some underlying theoretical structure.

Supply Model with Nerlovian Partial Adjustment: The simpler of the two, although somewhat ad hoc, is the multiproduct partial adjustment model developed by Clark, Siemans and Fleming (1990). In this model, a vector of desired acreages is defined in the typical Nerlovian manner

$$a_t^* = \alpha + \beta E_{t-1} p_t \tag{1}$$

where $a_t^*$ is a vector of desired acreages in year t, $E_{t-1} p_t$ is a vector of expectations made in year t-1 of the prices of all crop alternatives in year t, and $\beta$ is a matrix of response coefficients associated with $E_{t-1} p_t$. Now assume that the change in actual acreage is governed by the multiproduct partial adjustment principle. Thus:

$$a_t - a_{t-1} = D(a_t^* - a_{t-1}) + \epsilon_t \tag{2}$$

where $a_t$ is a vector of actual acreage in year t, D is a matrix of partial adjustment coefficients, and $\epsilon_t$ is a vector of error terms in year t. Note that if D is diagonal, then the model reduces to the special case of the single product partial adjustment principle. Substituting equation (1) into equation (2) and rearranging terms, we have

$$a_t = c + \Gamma a_{t-1} + RE_{t-1}\, p_t + \epsilon_t \qquad (3)$$

where $c = D\alpha$, $\Gamma = (I-D)$ and $R = D\beta$. The feedback matrix, $\Gamma$, given in equation (3) is, in general, not diagonal except in the special case where matrix D in equation (1) is diagonal.

A reduced form a vector autoregression (VAR) for the system of acreage responses can be found once an explicit price expectations process is specified. In this paper, we assume economic agents have rational expectations. Thus, suppose the VAR generating prices is given by

$$p_t = \gamma_o + \gamma_1(L)p_{t-1} + \gamma_2(L)\; x_{t-1} + \mu_t \qquad (4)$$

where $\gamma_o$ is a vector of intercept terms, $p_t$ is a vector of prices in year t, $x_{t-1}$ is a vector of all variables helpful in predicting prices other than prices themselves, $\gamma_1(L)$ and $\gamma_2(L)$ are matrices of prediction coefficients associated with $p_{t-1}$ and $x_{t-1}$ respectively, $\mu_t$ is a vector of error terms in year t and L is the lag operator. Given rational expectations then

$$E_{t-1}\, p_t = \gamma_o + \gamma_1(L)\; p_{t-1} + \gamma_2(L)\; x_{t-1} \qquad (5)$$

Substituting equation (5) into equation (3), we have

$$a_t = \Pi_o + \Gamma\; a_{t-1} + \Pi_1(L)p_{t-1} + \Pi_2(L)\; x_{t-1} + \epsilon_t \qquad (6)$$

where $\Pi_o = c + R\gamma_0$, $\Pi_1(L) = R\gamma_1(L)$ and $\Pi_2(L) = R\gamma_2(L)$. Equation (6) is the reduced form VAR for a system of acreage response functions that can be derived from a multiproduct partial adjustment model. The VAR given by equation (6) will have at least one unit root if det $(I-\Gamma)$ = det $D = 0$. Therefore, if the partial adjustment matrix D is such that it has an eigenvalue that is equal to zero, then $\Gamma$ in equation (6) will have an eigenvalue equal to one. Note that this result is independent of the solution to the prediction problem we postulated in equation (5). Therefore, the existence of the unit root is independent of the assumed expectation formation hypothesis.

Supply Model with Linear-Quadratic Costs: The derivation of a reduced form model observationally equivalent to equation (6) using optimal control techniques is more complicated but has the advantage that it is derived from explicit optimizing principles. The model we specify is similar to that derived by Eckstein (1984) and Tegene, Huffman and Miranowski (1988). Although the derivation in this paper uses the three-output case to simplify the notation, the model could be expanded to as many outputs as desired.

Consider an agricultural producer who produces three crops in year t denoted $y_{1t}$, $y_{2t}$ and $y_{3t}$, respectively. Crops 1 and 2 are assumed to be produced via linear production functions and a single input land.[1] That is:

$$y_{1t} = f_1 a_{1t}; \text{ and} \quad (7)$$

$$y_{2t} = f_2 a_{2t}; \quad (8)$$

where $a_{1t}$ and $a_{2t}$ are the acreages devoted to $y_{1t}$ and $y_{2t}$, respectively and $f_1$ and $f_2$ are response coefficients to $a_{1t}$ and $a_{2t}$ associated with $y_{1t}$ and $y_{2t}$, respectively. The third crop is assumed to be produced via the quadratic production function

$$y_{3t} = g_1 a_{3t} - \frac{g_{11}}{2} a_{3t}^2$$

where $g_1, g_{11} > 0$ are response coefficients and $a_{3t}$ is the acreage devoted to the third crop. The total acreage that can be allocated to the three crops is limited by the constraint

$$a_{1t} + a_{2t} + a_{3t} = a, \qquad (9)$$

where a is the total amount of land available for crop production. Denote the shadow value associated with land constraint (9) as $\lambda_t$ and assume there is no costly adjustment associated with the third crop. Then the equilibrium condition for the third crop will be

$$p_{3t} (g_1 - g_{11} a_{3t}) = \lambda_t, \qquad (10)$$

where $p_{3t}$ is the price of $y_{3t}$.

Now, the static cost function associated with using $a_{1t}$ and $a_{2t}$ in the production of $y_{1t}$ and $y_{2t}$ would be

$$\begin{aligned} c_t &= \lambda_t (a_{1t} + a_{2t}) \\ &= p_{3t} (g_1 - g_{11} a_{3t}) (a_{1t} + a_{2t}) \\ &= p_{3t} (g_1 a_{1t} + g_1 a_{2t} - g_{11} a_{1t} a_{3t} - g_{11} a_{2t} a_{3t}). \qquad (11) \end{aligned}$$

Using land constraint (9) to substitute $a_{3t}$ out of equation (11), the static cost function becomes

$$c_t = p_{3t} (h'a_t + 1/2 a_t' H a_t), \qquad (12)$$

where $a_t = [a_{1t}, a_{2t}]'$, $h = [g_1 - g_{11}a, g_1 - g_{11} a]'$ and H is a 2x 2 matrix whose elements $h_{ij} = 2g_{11}$ $(i,j =1,2)$. Now choose $p_{3t}$ as the numeraire price and let

$$p_t = [p_{1t}/p_{3t}, p_{2t}/p_{3t}]',$$

where $p_{1t}$ and $p_{2t}$ are prices of $y_{1t}$ and $y_{2t}$ respectively. Finally, assume that there are costs of adjustment associated with changing acreage between $y_{1t}$ and $y_{2t}$ of the form

$$M_t = 1/2\ \Delta a_t'\ M \Delta a_t$$

where M is a symmetric positive definite matrix of cost of adjustment parameters and $\Delta a_t = a_t - a_{t-1}$. Given these assumptions, the optimal control model becomes

$$V_t = \max_{\{a_t\}} E_0 \left[ \sum_{t=0}^{\infty} b^t \left(p_t'\ Fa_t - h'a_t - \frac{1}{2} a_t' H a_t - \frac{1}{2} \Delta a_t'\ M \Delta a_t\right)\right] \quad (13)$$

where $E_0$ is the expectations operator conditioned on information available at time 0, b is the discount factor, and $F = \mathrm{diag}(f_1, f_2)$. Equation (13) is of the form of a standard linear-quadratic model studied by Hansen and Sargent (1981) and therefore has the solution (without the prediction problem solved):

$$a_t = \bar{\Pi}_0 + \bar{\Gamma} a_{t-1} + \delta\left[\left(I - \frac{\bar{\Gamma}}{\lambda_2}\right) M^{-1} F \sum_{i=0}^{\infty} (b\lambda_1)^i E_{t-1} p_{t+i} - \left(I - \frac{\bar{\Gamma}}{\lambda_1}\right) M^{-1} F \sum_{i=0}^{\infty} (b\lambda_2)^i E_{t-1} p_{t+i}\right] \quad (14)$$

where $\bar{\Gamma}$ is chosen to satisfy the restrictions

$$\bar{\Gamma} + (b\bar{\Gamma})^{-1} = (1 + 1/b) I + (bM)^{-1} H, \quad (15)$$

$\lambda_1$ and $\lambda_2$ are the eigenvalues of the feedback matrix, $\bar{\Gamma}$, $\bar{\Pi}_0$ is a vector of intercept terms and $\delta = \lambda_1 \lambda_2 / (\lambda_1 - \lambda_2)$ (Hansen and Sargent (1981) and Clark (1987)).

The prediction aspect of equation (14) can be solved using Hansen and Sargent's (1980) optimal prediction formula once the

VAR generating prices given by equation (4) is known.[2] This results in the reduced form VAR:

$$a_t = \bar{\Pi}_0 + \bar{\Gamma} a_{t-1} + \bar{\Pi}_1(L)p_{t-1} + \bar{\Pi}_2(L)x_{t-1}. \tag{16}$$

Note that equation (16) is observationally equivalent to equation (5). It also has the property that det $(I-\tilde{\Gamma}) = 0$, or it contains a unit root. This is due to the fact that det H=0 in equation (13).[3] As in the Nerlovian multiproduct partial adjustment model, the unit root in equation (16) is the result of the underlying structural model and is independent of the solution to the prediction problem embedded in equation (14).

The existence of a unit root in the Nerlovian model is the result of an entirely different structural aspect of the problem than the existence of a unit root in the linear-quadratic model. Recall that the Nerlovian multiproduct partial adjustment model will have a unit root if det D=0, that is, if the partial adjustment matrix has a zero root. This arises as a special case of the <u>short-run</u> principle of partial adjustment. In contrast, the existence of a unit root in the linear-quadratic model arises because det H=0. This is due to a <u>long-run</u> singular quadratic cost function, the long-run in the case of the linear-quadratic model being when $\Delta a_t = 0$ in equation (13) so that costs of adjustment no longer affect acreage response. Therefore, although the existence of a unit root in a system of acreage response functions can be justified in either a Nerlovian or a linear-quadratic framework, the justification is fundamentally different in each case.

## 3. EMPIRICAL CONSIDERATIONS

The reduced form multivariate system of acreage responses that can be derived from either a Nerlovian multiproduct partial adjustment model or a linear-quadratic model can be written as:

$$a_t = \beta_0 + \beta_1 a_{t-1} + \beta_2(L) p_{t-1} + \beta_3(L) x_{t-1} + \epsilon_t \qquad (17)$$

where $\beta_i$ $(i = 0,\ldots,3)$ are parameter matrices to be estimated. From our previous discussion, the matrix $\beta_1$ will have a unit root if either det D=0 in the Nerlovian model or det H=0 in the linear-quadratic model. The empirical problem is to test for the existence of a unit root within the framework of equation (17). To do this, we use a method based on Dickey and Fountis (1989).

Care must be taken when applying the Dickey-Fountis procedure since the critical values of their test are sensitive to the time-series properties of the forcing variables, $p_t$ and $X_t$. It can be shown (see Clark, 1990) that the Dickey-Fountis test applied to (17) is only appropriate if $a_t$ does not Granger-cause the forcing variables and if there is no contemporaneous correlation between the error terms of (17) and the error terms of VAR's generating the forcing variables. One way to circumvent the problem is to include VAR's generating the forcing variables as part of an expanded VAR system. This system would include (17) as well as the VAR's for the forcing variables. The Dickey-Fountis test would then be applied to this expanded VAR system. The disadvantage of this approach is that it greatly increases the size of the system that needs to be estimated. An alternative, possibly simpler approach would involve carrying out preliminary tests on the null hypotheses that $a_t$ does not Granger-cause the forcing variables. If any of the null hypotheses are rejected then the Dickey-Fountis test must be applied to the expanded VAR system. If, however, the null hypotheses are not rejected then the Dickey-Fountis test may be applied simply to (17). For this alternative approach it is necessary to assume no contemporaneous correlation between the errors of (17) and those of the VAR's generating the forcing variables. For our analysis, the preliminary tests were carried out and the null hypotheses could not be rejected at the 10 percent significance level. (See Section 4 for the test results.) Hence we continue discussion of the Dickey-Fountis test applied only to (17).

Consider estimating (17) as a system of seemingly unrelated regressions (SUR), where $E(\epsilon_t \epsilon_t') = \Omega$. Let the unrestricted estimator be

$$a_t = \hat{\beta}_0 + \hat{\beta}_1 a_{t-1} + \hat{\beta}_2(L) p_{t-1} + \hat{\beta}_3(L) x_{t-1} + u_t. \tag{18}$$

Consider finding the canonical form of the matrix $(I - \hat{\beta}_1)$:

$$I - \hat{\beta}_1 = C \Lambda C^{-1} \tag{19}$$

where $\Lambda$ is the diagonal matrix of eigenvalues of $(I-\hat{\beta}_1)$ and C is the associated matrix of eigenvectors. Equation (19) can be rewritten

$$\hat{\beta}_1 = I - C \Lambda C^{-1}. \tag{20}$$

Substituting (20) into (18), we have

$$a_t = \hat{\beta}_0 + (I - C \Lambda C^{-1}) a_{t-1} + \hat{\beta}_2(L) p_{t-1} + \hat{\beta}_3(L) x_{t-1} + \epsilon_t. \tag{21}$$

Premultiplying (21) by $C^{-1}$ and setting $z_t = C^{-1} a_t$ gives

$$\Delta z_t = - \Lambda z_{t-1} + C^{-1}(\hat{\beta}_0 + \hat{\beta}_2(L)p_{t-1} + \hat{\beta}_3(L)x_{t-1}) + C^{-1}\epsilon_t. \qquad (22)$$

where $\Delta z_t = z_t - z_{t-1}$. Dickey and Fountis show that if (22) is estimated by generalized least squares regression, then the highest eigenvalue of the estimated $\hat{\beta}_1$ matrix ($\rho$) can be tested for the existence of a unit root. Given the number of observations (n), the procedure is to compare $n(\rho - 1)$ against the appropriate critical value from the $\rho_\mu$ Table in Fuller (1976, p.371).

Our estimator is slightly different from that of Dickey and Fountis because we allow for contemporaneous correlation between the error terms in (22). Therefore the covariance matrix used in the GLS estimation of (22) is

$$\bar{\Omega} = (C^{-1}\Omega C^{-1\prime}). \qquad (23)$$

If the null hypothesis of a unit root is not rejected, (22) is re-estimated with the restriction that a unit root exists. This is accomplished by restricting the relevant diagonal element of the eigenvalue matrix ($\Lambda$) to zero.

Let the restricted estimator of $\beta_1$ be defined as $\bar{\beta}_1$. This can be recovered from the restricted estimator of (22) by the transformation

$$\bar{\beta}_1 = I - C \bar{\Lambda} C^{-1}$$

where $\bar{\Lambda}$ is the restricted diagonal matrix of eigenvalues. The resulting $\bar{\beta}_1$ matrix will have precisely one unit root.

## 4. EMPIRICAL EXAMPLE

As an illustration of the above test procedure and the implications of a unit root, consider the example of the wheat/barley acreage response system for the prairie region of Canada. These two crops account for about 75 percent of the cropped acreage in the region. The model to be estimated is:

$$a_t = \beta_0 + \beta_1 a_{t-1} + (\beta_{21}+\beta_{22}L)p_{t-1} + (\beta_{31}+\beta_{32}\,L)x_{t-1} + \beta_4 w_t + \epsilon_t. \quad (24)$$

This is similar to the form of (17) with first-order polynomials in the lag operator. The difference is in the addition of $w_t$, a vector of variables not included in our theoretical model but which are deemed to be helpful in explaining prairie wheat/barley acreages. In this model:

$$a_t = \begin{vmatrix} AW_t \\ AB_t \end{vmatrix}; \qquad p_{t-1} = \begin{vmatrix} PW_{t-1} \\ PB_{t-1} \end{vmatrix};$$

$$x_{t-1} = \begin{vmatrix} IPW_t \\ IPB_t \end{vmatrix}; \qquad w_t = \begin{vmatrix} WS_{t-1} \\ LIFT \end{vmatrix};$$

where:

$AW_t$ = planted wheat area in Saskatchewan, Manitoba and Alberta in year t, thousand acres [Source: Statistics Canada, #22-002].

$AB_t$ = planted barley area in Saskatchewan, Manitoba and Alberta in year t, thousand acres [Source: Statistics Canada, #22-002].

$PW_t$ = average price received by Saskatchewan, Manitoba and Alberta farmers for wheat in year t, \$/tonne [Source:

Statistics Canada, #22-002].

$PB_t$ = average price received by Saskatchewan, Manitoba and Alberta farmers for barley in year t, $/tonne [Source: Statistics Canada, #22-002].

$IPW_t$ = initial payment made by the Canadian Wheat Board (CWB) on No. 1 CWRS wheat in year t, $/tonne [Source: CWB Annual Reports].

$IPB_t$ = initial payment made by the CWB on No. 1 feed barley in year t, $/tonne [Source: CWB Annaul Reports].

$WS_{t-1}$ = July 31 Canadian wheat stocks in all positions, at the end of year t-1, million bushels [Source: CWB Annual Reports].

LIFT = zero-one variable to represent the impact of the federal government's LIFT (lower inventories for tomorrow) program in 1970. The variable takes on a value of 1 in 1970 and 0 otherwise.

In this model, all of the variables in $p_{t-1}$ and $x_{t-1}$ were deflated by the aggregate input price index for Western Canada [Source: Statistics Canada, #62-002 and #62-004]. The variables included in $w_t$ were lagged wheat stocks and a zero-one dummy variable. Lagged wheat stocks was included to represent the restrictive effect on wheat acreage of the CWB's delivery quota policy. The zero-one variable was included to represent the restrictive effect on wheat acreage of the federal government's LIFT program instituted in 1970 for just the one year. These variables have been frequently used in other studies that have attempted to estimate Canadian prairie acreage response functions (e.g., Meilke, 1976).

To determine whether a straightforward application of the Dickey-Fountis test is appropriate, preliminary tests were carried out on the null hypotheses that the components of $a_t$ do not Granger-cause (denoted "$\nrightarrow$" below) the components of $p_t$ or $x_t$. Allowing for lag responses of up to two years, the relevant F test statistics are as follows:

$F\ (a_t \nrightarrow PW_t) = 1.72;$

$F\ (a_t \nrightarrow PB_t) = 1.57;$

$F\ (a_t \nrightarrow IW_t) = 0.35;$

$F\ (a_t \nrightarrow IB_t) = 1.02.$

These values may be compared with the critical value from an F(4,24) distribution. At the 10 percent significance level we fail to reject the null hypotheses that the components of $a_t$ do not Granger-cause the forcing variables. Hence we can proceed to the straightforward application of the Dickey-Fountis test.

In Table 1 are presented the estimation results. In columns (1) and (2) are presented the coefficient estimates for the unrestricted model. The largest eigenvalue of the feedback matrix ($\hat{\beta}_1$) is 0.91. This is close enough to 1 to suggest the possibility of a unit root. The appropriate test statistic is $n(\rho - 1)$ which in this case is -3.33 ($n = 37$, $\rho = 0.91$). This is compared against the appropriate critical value from the $\rho_\mu$ Table in Fuller (1976). Using a five percent significance level, example critical values are -12.5 for n = 25 and -13.3 for n =

50. Hence we fail to reject the null hypothesis of a unit root in $\beta_1$. The model was re-estimated, imposing the restriction that the largest eigenvalue of $\beta_1$ is one. The results of the restricted regression are presented in columns (3) and (4) of Table 1. Note that the coefficients in the restricted model are only marginally different from those in the unrestricted model. However, as we shall see below the prediction implications of the two models are very different.

In what follows, we consider an illustration of the prediction implications of the two models given a one-time exogenous shock to the system of supply equations. The shock is provided by the 1987 Special Canadian Grains Payment (SCGP).

The SCGP was introduced in December 1986 by the Canadian Government. It was designed to offset the impact on Canadian grain producers of the subsidy war between the United States and the European Community. The program provided for a payout of $1 billion in 1987, of which $860 million was provided to Western Canada. This transfer amounted to $17.82 per tonne on wheat and $11.77 per tonne on coarse grains (including barley). Many observers have argued that this policy did not distort production since the payment was based on historical yields. However, since the policy was introduced to offset a price decline attributable to the subsidy war between the U.S. and the EC, producers may have interpreted the subsidy as a price support which would have a production effect. Suppose that, in determining production plans for period T, producers interpret the expected price vector to include both price and the SCGP in period T-1. Then, with respect to (24), the output price component can be factored as follows:

$$\beta_2 \text{ (L) } p_{T-1} = \beta_{21} \ (p_{T-1} + s_{T-1}) + \beta_{22} \ p_{T-2}$$

where,

$p_t$ = output price vector, year t and

$s_t$ = vector of SCGP's per unit of output, year t.

In this way the SCGP for 1987 enters the supply model as a one-period price shock. Given the estimated equations in Table 1, the purpose of this illustration is to compare the effects of this temporary shock on the supply model when the model has a unit root and when it does not.

The results of the dynamic simulation over a 50-year time horizon beginning in 1988 are presented in Figure 1. This figure shows that in the short run there is very little difference between the two models. Both models suggest that in the first year following the announcement there would be an expansion in wheat area and a contraction in barley area. This situation is reversed in the second year when barley area expands while wheat area contracts. However, in the longer run, Figure 1 reveals that the two models yield very different response paths. In the case of the stationary supply model, the effect of the price shock in period 0 (1987) declines to zero. However, when the supply model is assumed to have a unit root, the effect of this price shock does not decline to zero. There is a permanent effect on wheat and barley acreage. Wheat area stabilizes at approximately 300,000 acres below the levels that would have been achieved without the 1987 price shock while barley area stabilizes at approximately 350,000 acres above the levels that would have been achieved without the 1987 price shock. The magnitude of the difference between the stationary model and the unit root model suggests that the determination of whether or not a supply model has a unit root is not insignificant.

One interesting implication of the feedback model is that the presence of one unit root generally implies that a temporary shock has a permanent effect on <u>all</u> acreage variables. Only in the special case where $\Gamma$ is a diagonal matrix will one unit root permanently affect only one acreage variable. An example of this would be a system of a single product feedback equations, such as a set of acreage equations involving Nerlove's single-product partial adjustment principle. A unit root would be the limiting case in which the Nerlovian partial adjustment coefficient would be equal to zero.

## 5. SUMMARY AND CONCLUSIONS

This paper has been concerned with the prediction implications of a multiproduct acreage response model with a feedback component. Two alternative but observationally equivalent theoretical structures were discussed which yield such a model. They were the Nerlovian multiproduct partial adjustment model and the supply model with linear-quadratic costs.

For prediction purposes, a critical feature of such a model is whether or not the feedback matrix contains a unit root. A computational procedure developed by Dickey and Fountis was used to determine the presence of a unit root.

To illustrate the prediction implications of a unit root, we estimated a simple two-equation acreage response model for the Canadian prairies. We found that the null hypothesis of one unit root could not be rejected at the five percent significance level. In a restricted version of the model, we imposed a unit root on the feedback matrix. The resulting coefficients were only marginally different from those of the unrestricted model. The prediction implications of the unrestricted and restricted models were then compared with reference to a single one-period shock. The shock was assumed to arise from the 1987 Special Canadian Grains Payment. In the first three years the responses predicted by the unrestricted and restricted models were very similar. However, thereafter the responses diverged markedly. The response in the unrestricted model tapered off to zero, while that of the restricted model tapered off to some non-zero value. This latter result implies that the presence of a unit root leads to a permanent response to a temporary shock.

Consequently, the time-series containing a unit root implies a permanent maladjustment within the wheat/barley acreage response system due to the special Canadian grains payment instituted in 1986 by the Canadian government. At first blush, this policy would seem to be rather innocuous when compared to the size of the subsidy programs introduced in the U. S. and the

E. C. . However, our results show that in fact this policy could have been far from harmless, leading to a permanent reduction of approximately 300,000 acres planted to wheat and a permanent increase of approximately 350,000 acres planted to barley. If the results of this paper are robust in the sense that acreage response systems in other countries also contain unit roots, then one can speculate that the permanent misallocation of resources caused by governments will continue to blight world grain markets for years to come. This last comment applies even if the present subsidy programs to grain producers that are in place around the world are removed.

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Table 1: Seemingly Unrelated Regression Results for Wheat and Barley Acreage Response System (000's acres, 1949-1986)*

| Regressor | Unrestricted Model Wheat Acreage (1) | Unrestricted Model Barley Acreage (2) | Restricted Model Wheat Acreage (3) | Restricted Model Barley Acreage (4) |
|---|---|---|---|---|
| Intercept | 30505.75 (5.12) | 5189.06 (1.49) | 30963.83 (6.34) | 3447.33 (1.60) |
| $AW_{t-1}$ | .356 (2.89) | -.179 (-3.28) | .388 | -.197 |
| $AB_{t-1}$ | -.439 (-2.92) | .770 (7.15) | -.483 | .845 |
| $PW_{t-1}$ | 18.56 (1.81) | -22.93 (-3.03) | 19.21 (1.90) | -23.89 (-3.20) |
| $PW_{t-2}$ | -33.61 (-2.10) | 37.72 (3.32) | -34.93 (-2.26) | 40.64 (3.79) |
| $PB_{t-1}$ | 2.06 (.15) | 17.55 (1.97) | -1.84 (-.16) | 20.96 (2.69) |
| $PB_{t-2}$ | -10.68 (-.57) | -15.80 (-1.59) | -5.56 (-.36) | -18.93 (-2.07) |
| $IPW_t$ | 67.58 (2.58) | -18.43 (-.95) | 64.76 (2.57) | -13.14 (-.72) |
| $IPW_{t-1}$ | -68.66 (-2.28) | .70 (.03) | -64.37 (-2.26) | -6.30 (-.31) |
| $IPB_t$ | -71.63 (-1.93) | 68.94 (2.75) | -82.96 (-2.90) | 80.67 (3.96) |
| $IPB_{t-1}$ | 58.10 (1.66) | -76.84 (-2.99) | 66.42 (2.23) | -87.64 (-4.02) |
| $WS_{t-1}$ | -8.08 (-2.33) | | -7.30 (-2.49) | |
| LIFT | -12555.98 (-7.57) | | -12870.10 (-8.32) | |
| $R^2$ | .90 | .80 | .90 | .79 |
| Durbin-M | -1.10 | .94 | -1.13 | .98 |

* t-values in parentheses.

## FOOTNOTES

1 The reader is referred to Tegene, Huffman and Miranowski (1988) for the assumptions required to make output a function of the single input land.

2 This statement is not quite correct because the vector autoregression generating the vector $x_t$ also needs to be estimated before a reduced form VAR can be derived from this system. To do this, we will change the notation slightly from that used in the main body of the text. Let $x_{1t} = [p_t, x_t]'$ and redefine equation (3) as $x_{1t} = \theta_0 + \theta(L)x_{1t-1} + \mu_t$ where the $\theta_0$ and $\theta(L)$ matrices are now of appropriate dimension for the $x_{1t}$ vector. Using the above notation, the reduced form autoregression for the linear-quadratic model can be derived using Hansen and Sargent's (1980) optimal Wiener-Kolmogorov prediction formula given by;

$$\sum_{i=0}^{\infty} (b\lambda_e)^i E_{t-1}p_{t+i} = U\,\theta(b\lambda_e)^{-1} \; [[I + \sum_{j=1}^{r-1} \sum_{k=j+1}^{r} (b\lambda_e)^{j-k}\,\theta_j L^k]$$

$$-(b\lambda_e)^{-1}I]x_{1t-1}, \qquad e = 1,2$$

where $U = [I,0]$, $\theta(b\lambda_e) = [I - b\lambda_e\,\theta_1 - (b\lambda_e)^2\,\theta_2 \ldots - (b\lambda_e)^r\theta_r]$

and r is the length of the autoregression associated with $x_{1t}$. The above equation is slightly different from that given in Hansen and Sargent (1980) because it has been modified to account for the fact that $p_t$ is a vector rather than a scalar.

3 This can be shown as follows. The long-run or steady state is defined in the linear-quadratic model when $a_t = a^*$ for all t, and $E_{t-1}\,p_t = p^*$ for all t. Imposing these restrictions on the linear-quadratic objective function given in equation (12) would yield a long-run acreage response system of

$$a^* = H^{-1}\,F\,p^*.$$

Since det H=0, this system of acreage responses do not satisfy the second order conditions for a maximum and an infinite steady-state response is implied. Imposing the same steady-state restrictions on the system of acreage responses without prediction solved given in equation (13) would yield the solution

$$a^* = (I - \tilde{\Gamma})^{-1}\,R\,M^{-1}\,F\,p^*,$$

where $R = \delta[(I - \frac{\bar{\Gamma}}{\lambda_2})/(1-b\lambda_1) - (I - \frac{\bar{\Gamma}}{\lambda_1})/(1-b\lambda_2)]$.

Equating the above two steady state systems

$$H = M\, R^{-1}\, (I - \bar{\Gamma})$$

and therefore

$$\det H = \det M \det R^{-1} \det (I-\bar{\Gamma}).$$

If either $\det M = 0$ or $\det R^{-1} = 0$ then, in view of equation (13), a finite solution for acreage in the short-run would be violated. Therefore

$$\det H=0 \text{ implies } \det (I-\bar{\Gamma}) =0.$$

Figure 1: Acreage Responses to a Temporary Shock in the Restricted and Unrestricted Models