

# Modelling soybean prices in a changing policy environment

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Accurate forecasts of commodity prices are an important ingredient in the policy formation process. A commodity price forecasting procedure used routinely by the US Department of Agriculture in their policy and market analysis activities is a simple, linear, reduced-form regression model that predicts season-average farm prices (SAFP) using policy variables and the ratio of total ending stocks to use. This approach is extended to the soybean SAFP to estimate a benchmark model using annual data. Also several specification issues related to this estimation framework are addressed. Evaluation suggests that the standard forecasting procedure may be affected by the fact that the ratio of stocks to use is endogenous to prices. In addition, important structural changes are revealed in these relationships over time. A model is then considered that allows parameters to shift gradually. Improvements in the accuracy of model forecasts allowed by this parameter switching technique are identified and discussed. In addition, the exact nature of the structural shifts is evaluated using dynamic impulse response functions.

## I. Introduction

The oilseed products complex is an important component of the US agricultural sector. In 2000, a record of over 74 million acres were planted in soybeans, representing over 29% of total acreage planted to the eight major programme crops (USDA, 2002). Soybean acreage has increased steadily since 1990, when only 58 million acres were planted. By the end of the decade, harvested soybean acres actually exceeded total harvested corn acres. In 2002, over 72.4 million acres of soybeans were harvested in the USA, as compared to 69.3 million acres of corn.

US agricultural policies over the post-war period included a variety of instruments intended to enhance prices and income while controlling acreage and

output. The production of many crops was constrained through the use of 'base acreage,' which entitled producers to price support payments on a limited acreage. From a historical perspective, soybeans are rather unique among the major government programme crops in that they were not eligible for target-price deficiency payments, nor were they subject to the explicit acreage restrictions of other programme crops. However, the acreage-idling and base-acreage requirements, as well as government stock-holding behaviour of other programme crops, have indirectly affected soybean acreage decisions in the past. Soybeans have been eligible for government price support loans for the past 60 years. In recent years, soybeans have benefited from a high loan rate relative to corn. This, coupled with eligibility

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for government marketing loan gains and loan deficiency payments, has stimulated production of soybeans (Westcott and Price, 2001).

Comprehension of the various factors underlying price determination is essential in order to understand the effects of policy changes and other shifts in market factors. Westcott and Hoffman (1999) considered the effects of market and policy factors using annual models of US farm prices for corn and wheat. Their results confirmed the importance of the stocks-to-use ratio as an indicator of market supply and demand conditions. In addition, they used a number of discrete indicators of changing policy conditions. These indicators confirmed that changes in the policy environment can have important impacts on market prices and may influence the relationship between supply and demand factors and prices.

Such models have an important role in the development and validation of USDA projections of prices. Most analysts evaluating the determinants of soybean prices, even in an informal context, consider the role of the stocks-to-use ratio. Each month, the USDA analyses major agricultural markets and publishes annual supply, demand, and price projections. Simple models relating price to observable supply and demand factors, such as the stocks-to-use ratio, are important tools in assessing predictions of such factors and price forecasts.

The objective of the present analysis is to extend the models of Westcott and Hoffman (1999) by considering factors affecting US soybean prices. It is recognized that a more comprehensive specification of soybean price determination would incorporate the demand for soybean's joint products, meal and oil, in a larger multi-equation framework. However, the goal of this research is to investigate the potential for using the simpler, single-equation stocks-to-use framework as an aid in monthly supply and demand analysis. Following Westcott and Hoffman (1999), the focus is on the stocks-to-use ratio as an indicator of market supply and demand conditions. Policy variables that may have impacted price relationships are also considered. Westcott and Hoffman (1999) focused on the 1975–1996 period. In contrast, a much longer span of data is considered here and explicit attention given to the potential for structural changes in the relationships between prices and market factors.

Focus is also on an issue not previously considered in evaluations of the relationship between the ending stocks-to-use ratio and prices—the potential endogeneity of these variables. It would certainly be expected that prices adjust as supply is realized and as total use changes. However, demand theory suggests that total use will decline as prices increase—suggesting

the potential for simultaneity between total use and prices. Even more likely is the possibility that stock holding behaviour is influenced by prices. Low prices typically serve as an incentive for agents to store a commodity in the hope that future market conditions will result in more favourable prices. Thus, ending stocks will be directly influenced by prices, making them endogenous in models relating prices to the stocks to use ratio. Informal evaluations and empirical models generally do not consider the likelihood that prices, stocks, and total use are jointly determined variables. This may have important implications even for the simple linear regression models that are often used to consider soybean price determinants.

The plan of the paper is as follows. The next section gives a brief review of factors suspected to be relevant to price determination in the US soybean market. Section III presents an empirical analysis of price determinants for soybeans. Structural change and endogeneity tests are discussed. In addition, a gradual switching model is developed that endogenizes the break point and speed of change inherent in the structural break. A more general forecasting model is then considered. In particular, a quarterly vector autoregressive (VAR) model is developed that incorporates the gradual switching considered for the single equation analysis. Improvements in the accuracy of model forecasts allowed by this parameter switching technique are identified and discussed. In addition, the exact nature of the structural shift is evaluated using dynamic impulse response functions. The final section of the paper includes a review of the analysis and offers some concluding remarks.

## II. Conceptual Issues

Prices are determined by the interaction of supply and demand functions. Thus, a reduced-form expression for prices will relate prices to factors that influence supply and demand. As Westcott and Hoffman (1999) note, these factors are often summarized in the stocks-to-use ratio. Stocks adjust in response to shocks to supply and demand. Stocks will decrease in response to negative production shocks and the concomitant higher prices and will increase when production is high and prices are low. Total use, which includes domestic consumption and exports, is generally more stable and tends to shift gradually over time. Of course, as noted above, both factors may be simultaneously determined along with prices.

Following Labys (1973) and Westcott and Hoffman (1999), an equilibrium model for a storable commodity in a competitive market generally consists of

a supply equation, a demand equation, a stocks equation, and an identity describing equilibrium. Supply ( $S$ ) is a function of price ( $p$ ) (or, more accurately, expected price) and factors ( $z$ ) reflecting production shocks:

$$S_t = s(p_t, z_t) \quad (1)$$

Demand ( $D$ ) is a function of prices and other demand shifters ( $y$ ):

$$D_t = d(p_t, y_t) \quad (2)$$

Stocks ( $K$ ) are influenced by prices and possibly other factors ( $v$ ) reflecting storage costs and capacity constraints:

$$K_t = k(p_t, v_t) \quad (3)$$

Market equilibrium requires  $S_t - D_t - K_t = 0$ . This allows for a price-dependent reduced form expression to be solved that is a function of stocks and supply and demand shifters:

$$p_t = f(K_t, z_t, y_t) \quad (4)$$

Supply and demand shifters will include variables indicating changes in policy regimes as well as factors affecting weather and demand shocks. As noted above, it has become common to consider stocks in terms of their size relative to total usage. Thus, a common specification includes  $K_t/D_t$ , though, as noted earlier in this paper, such a specification does not really represent a reduced form and thus may be subject to simultaneous equation biases. Further, to the extent that stock holdings are influenced by prices,  $K_t$  may also be endogenous to price. In their analysis of corn and wheat prices, Westcott and Hoffman (1999) regressed prices (in logarithmic terms) on the logged ratio of total year-end stocks to use, the ratio of CCC held stocks to use, an interaction term that included a dummy variable representing the years 1978–1985 and loan rate, and a dummy variable for 1986 – a year that was revealed to be an outlier in preliminary analyses. The years 1978–1985 were singled out as a period when government intervention via the Famer-Owned Reserve (FOR) programme, with high release prices and high loan rates relative to market prices, isolated significant amounts of corn and wheat from the market. Their wheat equation also included feed use and corn prices in the summer months, while

excluding the 1986 dummy variable. Their empirical results confirmed a strong inverse relationship between price and the stocks-to-use ratio.

### III. Empirical Analysis

This analysis begins with a simple regression analysis of a form similar to that used by Westcott and Hoffman (1999) in their analysis of corn and wheat prices:

$$P_t = \alpha_0 + \beta_1^* (K_t/U_t) + \beta_2^* \text{LDP} + \beta_3^* \text{Drought} + \beta_4^* \text{Loan rate} + \beta_5^* \text{Loan rate} * D_{78-85} \quad (5)$$

where all continuous variables are in logarithmic terms, LDP is a discrete indicator for the years in which meaningful loan deficiency payments were made (1998–2000), Drought is a discrete indicator variable for drought years (1980, 1983, and 1988), and  $D_{78-85}$  is a discrete indicator representing the period 1978–1985. Westcott and Hoffman (1999) found that government programmes had the most significant effect on prices during this period.

Data were collected from a variety of USDA sources.<sup>1</sup> The data span the period 1942–2000. The soybean price is the season average price received by US farmers.<sup>2</sup> Stocks, denoted in Table 1 as Stocks<sub>4</sub>, are ending (4th quarter) stocks.

Estimates of the Equation 5 (Model 1) are presented in Table 1. Although the results suggest that this simple regression equation explains a considerable proportion of the variation in US soybean prices, there are several reasons to question this specification. These concerns are related to structural shifts that may have occurred during the estimation period, the issue of price deflation, and the endogeneity of stocks-to-use.

For example, one surprising result is that the overall stocks-to-use ratio does not appear to significantly influence soybean prices. The coefficient, though negative, is not statistically significant. For a shorter period of data (1975–1996), Westcott and Hoffman (1999) found a strong negative relationship between the stocks-to-use ratio and price, as would be expected. An examination of the data provides an explanation for this result.

Figure 1 illustrates the relationship between the stocks-to-use ratio and prices. A clear structural

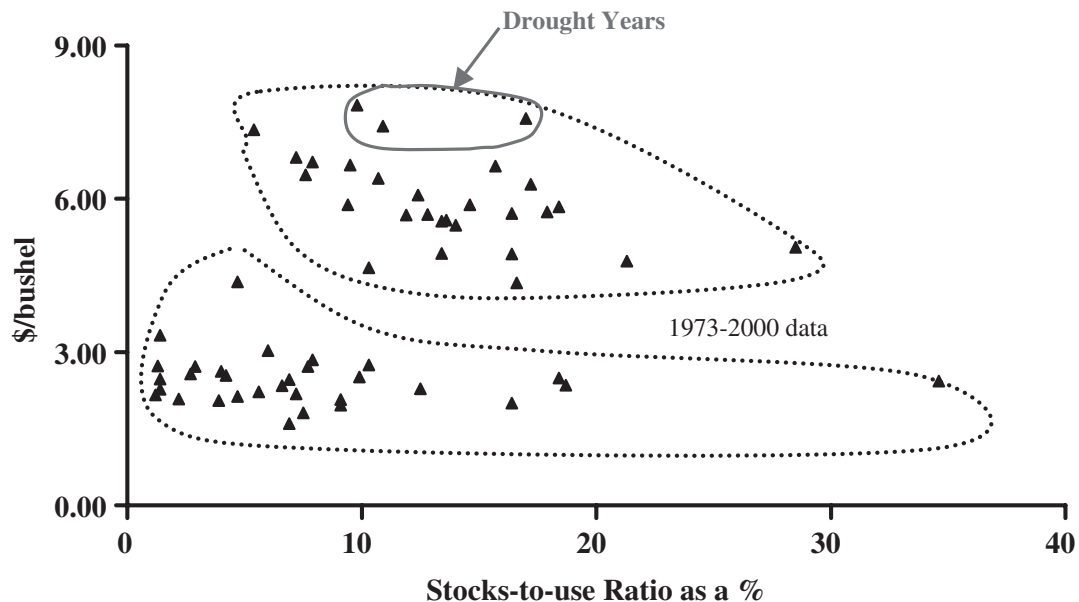
<sup>1</sup> Exact sources as well as the original estimation data are available from the authors on request.

<sup>2</sup> Concerns are frequently raised regarding the possibility that prices may be nonstationary. Standard Dickey–Fuller unit root tests were considered for soybean prices and it was found that prices do appear to be nonstationary. However, when the approach of Perron (1989) was followed and a structural break and the effects of drought years allowed for (as the results below suggest may be relevant) soybean prices did not appear to have a unit root.

**Table 1. OLS estimates of soybean price model**

Variable	Model 1	Model 2	Model 3	Model 4
Intercept	0.1232 (0.1085)	-2.8677 (0.5154)*	0.1058 (1.1439)	0.5694 (1.0117)
Drought	0.2282 (0.1610)	0.2770 (0.1253)*	0.0864 (0.1303)	0.0813 (0.1145)
(Loan rate)* $D_{78-85}$	-0.0281 (0.0735)	0.0316 (0.0580)	0.0339 (0.0534)	0.0473 (0.0471)
LDP	-0.3311 (0.1927)*	-0.2434 (0.1504)	-0.2740 (0.1202)*	-0.2746 (0.1056)*
Loan rate	1.0173 (0.1172)*	-0.2557 (0.2347)	-0.0557 (0.2245)	-0.1440 (0.1985)
Stocks <sub>4</sub> /Use	0.0334 (0.0426)	0.0412 (0.0331)	—	—
Stocks <sub>1</sub> /Use $t_{-1}$	—	—	-0.6342 (0.2092)*	-0.3175 (0.2003)
Corn stocks <sub>1</sub> /Use $t_{-1}$	—	—	—	-0.4369 (0.1100)*
FPPI $_{t-1}$	—	1.0659 (0.1812)*	0.9924 (0.1803)*	1.0449 (0.1589)*
Adjusted $R^2$	0.7424	0.8447	0.8604	0.8922
Wu-Hausman test 1	19.7136	—	—	—
Wu-Hausman test 2	2.5600	—	—	—
Chow test at 1972/73	18.8200	—	—	—

Note: Stocks<sub>1</sub> = 1st quarter stocks; Stocks<sub>4</sub> = ending stocks. Numbers in parentheses are standard errors. Asterisks indicate statistical significance at the  $\alpha = 0.10$  or smaller level.

**Fig. 1. Historical relationship between soybean price and stocks-to-use**

break in this relationship appears to have occurred around 1973. To the extent that this break is ignored, the estimates will suffer from specification biases. A standard Chow test of the significance of this break was applied and found to be very significant, with an  $F$ -value of 18.82, which exceeds the critical values at all conventional levels of significance. It was not possible to test for change in the drought, LDP, and loan rate–dummy variable interaction since these variables are all zero in the early (pre-1973) regime.

The early 1970s was a period of significant changes in world agricultural markets when nearly two decades of fairly stable commodity prices ended with a dramatic spike. This tumultuous period was marked by an unexpected surge in world grain

demand and trade, coupled with poor harvests and rapid, dynamic macroeconomic changes (Riley, 1996). An emergence of international markets from the post-Bretton Woods period enhanced international trade in agricultural commodities. In addition, significant development of soybean production in other competing (Southern Hemisphere) markets occurred during this period. Thus, it is not surprising that structural relationships for soybean prices appear to have shifted during this period.

Another estimation issue involves the fact that nominal prices are the target of the analysis, and yet no adjustments are made for possible movements in the overall price level. The issue of deflating agricultural prices to account for movements in

overall prices is a tricky one. It is widely recognized that real (i.e., deflated) agricultural prices have trended downward over time, although the general levels of nominal (non-deflated) prices have not changed significantly over time.

To account for inflation, an alternative specification (Model 2) was considered that adds an indicator of the overall price level—the farm producer price index (FPPI). The FPPI was lagged one period to obviate any additional endogeneity concerns. This is of minor significance in light of its role as an indicator of long-run aggregate price movements.

This is a flexible alternative to actually deflating the prices since this specification nests a situation of actual deflation (implied by a coefficient value of 1) as well as any other adjustment that may be more suitable. The results would seem to suggest that the loan rate and the FPPI are highly correlated. The loan rate loses its statistical significance in the new specification while the producer price index is significant with a value reasonably close to one. The in-sample explanatory power of the amended specification appears to be considerably higher than the simple specification.

Finally, in addition to possible misspecification concerns regarding structural change and movements in aggregate prices, the aforementioned issues relating to the possible endogeneity of the stocks-to-use ratio are relevant to an evaluation of the simple specification. As already noted, conceptual and intuitive considerations lead to the suspicion that the ending stocks-to-use ratio may be jointly determined with prices. To evaluate this possibility, standard Wu–Hausman tests of endogeneity was considered. It was assumed that the ratio of the 1st-quarter stocks (December of the September–August crop year) to the preceding year's use (referred to as  $\text{Stocks}_1/\text{Use}_{t-1}$  in Table 1) is exogenous to farm prices received. This is used as an instrument for ending stocks and conduct the Wu–Hausman test for endogeneity. The results are somewhat startling—the Wu–Hausman test strongly confirms the significance of endogeneity. The test statistic is 19.7, which exceeds the chi-square critical value at conventional levels of significance. When the ending stocks-to-use ratio is replaced by this instrument (Model 3), the stocks-to-use ratio reveals strong statistical significance and the expected negative effect on prices.

Of course, it might be also suspected that the ratio of first-quarter stocks to lagged use might also be endogenous. To evaluate this possibility, the Wu–Hausman test was repeated for first-quarter stocks to lagged use, using the preceding year's ending stocks as an instrument. In this case, the test statistic has a value of 2.56, which is less than the

chi-square critical value at conventional levels of significance. Thus, a model that utilizes first-quarter stocks rather than ending stocks, as is common, reveals the expected negative relationship between relative stocks and prices. However, the results imply that the conventional approach of including ending stocks in a price dependent reduced form model may result in estimates that suffer from simultaneity biases.

It is also of interest to consider the relationship between corn stocks and soybean prices. Corn and soybeans are competing crops that are often grown in the same areas. It would be suspected that corn stocks are likely to be highly correlated with soybean stocks in light of the fact they are grown in common geographic areas and are thus likely to be similarly affected by weather shocks. To a lesser extent, soybeans and corn are also substitutes in consumption as both provide oil and feed ingredients. Also, corn stocks may indirectly affect soybean supply and use via their influence on relative corn–soybean prices and associated producer behaviour.

The model was repeated including the ratio of ending corn stocks to use (model 4). The results are quite similar, though the soybean relative stocks variable loses much of its significance. This likely reflects suspicions that corn and soybean stocks are highly correlated and thus likely convey similar information to market participants.

In summary, the results raise important concerns about the simple specification that uses ending stocks to use and ignores structural change. This is not to say that earlier papers (e.g., Westcott and Hoffman, 1999) necessarily ignored structural change. On the contrary, their focus on later periods of data for analysis reflects recognition of the structural change issue. An analysis of shifts in the relationship between the stocks-to-use ratio and prices confirms a structural break that appears to have occurred in 1973. In addition, the intuition that the ending stocks-to-use ratio may be jointly determined with price is confirmed, suggesting the potential for biases in empirical results that ignore this issue.

#### IV. A Switching Model of Structural Change

A variety of methods for modelling structural change have been proposed in the literature. Almost all such methods entail a shift or break in parameters over time. The simplest case involves the standard Chow test, in which a break at a predetermined point in the data is assumed. Of course, a problem associated with such an approach is that the timing of such a break must be known a priori. Alternatives to specifying the

break prior to the test involve searching for the most significant break over a range of possible dates. Research by Andrews (1993) has demonstrated that conventional inference procedures are not applicable in such cases. In particular, the resulting  $F$ -statistic is a *supremum* value over the range defined by the search space. The distribution of a  $\sup(F)$  is not the same as a standard  $F$  and thus alternative inferential procedures are needed.

In addition to the issues associated with searching for a break point, conventional methods for modelling structural change are limited by the fact that they typically assume that such change occurs instantaneously. Although abrupt structural shifts are certainly possible, it would be expected that gradual structural change is more likely to occur in economic relationships. Thus, a method that allows the data to choose the break point and the speed of adjustment between regimes is desirable. In this vein, a gradual switching regression method is utilized.

Gradual switching regressions were introduced by Tsurumi *et al.* (1986). In contrast to their approach, a smooth transition function is utilized to represent the speed and timing of a structural shift between regimes. The use of transition functions as a means for modelling structural shifts was introduced by Bacon and Watts (1971). In the present analysis, the shift is allowed to occur gradually and the timing and speed of the shift identified using the estimation data. In particular, structural change is represented in terms of a shift in the parameter set from  $\beta^{(I)}$  to  $\beta^{(II)}$ . A mixing term  $\lambda_t$ , that is constrained by construction to lie in the open interval  $(0, 1)$ , is used to represent shifting between regimes. The specification of the mixing problem allows the simple regression relationship considered above  $y = X\beta$  to be rewritten as:

$$y_t = (1 - \lambda_t) X_t \beta^{(I)} + \lambda_t X_t \beta^{(II)} + e_t \quad (6)$$

The mixing term  $\lambda_t$  is given by:

$$\lambda_t = \Phi((t - \mu)/\sigma) \quad t = 1, \dots, N \quad (7)$$

where  $t$  is an observation counter ( $t = 1, 2, \dots, T$ ),  $\Phi$  is the normal cumulative distribution function (cdf) and  $\mu$  and  $\sigma$  are parameters to be estimated. The smooth transition function approach has much in common with the smooth threshold modelling techniques of Terasvirta (1994). A similar approach to specification and estimation is undertaken there, though in that case observations may switch between

regimes more than once. In the present approach, the regime switch is permanent.

Note that  $\mu$  represents the observation lying one-half way between regimes I and II (i.e., for which  $\lambda_t = 0.50$ ). The bandwidth parameter  $\sigma$  represents the speed of adjustment between regimes, with larger values of  $\sigma$  corresponding to more gradual adjustments between regimes. Note that  $\lim_{x \rightarrow \infty} \Phi(x) = 1$  and  $\lim_{x \rightarrow -\infty} \Phi(x) = 0$ .<sup>3</sup>

Estimation of the switching regression model may pose challenges. Though estimation follows standard nonlinear regression methods, identification issues may arise as the break point  $\mu$  nears either end of the data and as the speed of adjustment becomes very fast (i.e., as  $\sigma$  approaches zero). The following estimation approach is adopted in this analysis. First a standard grid search over possible values of  $\mu$  and  $\sigma$  is considered. The values that minimize a sum of squared error criterion (or, equivalently, that maximize a  $F$ -test of the specification against one without structural shifts) are selected. The optimal values of  $\mu$  and  $\sigma$  are then used as starting values in a standard non-linear regression model.

Estimates of the gradual switching regression models are presented in Table 2. Three alternative specifications are considered. The first includes only loan rates and the stocks-to-use ratio (using the ratio of first-quarter stocks to last year's use). The second includes dummy variables representing drought years and the LDP as well as the producer price index.<sup>4</sup> The final specification also includes the corn stocks-to-use ratio.

In all three cases, the  $\mu$  estimates indicate a strong and immediate structural break centred at observation number 31, corresponding to 1972. Furthermore, the  $\sigma$  estimates are quite small (from 0.86 to 0.95) suggesting a very rapid adjustment phase of approximately 2–3 years. Thus, the results are consistent with the Chow tests reported earlier as well as with earlier research that has argued in favour of structural breaks at this point in time. The speed and timing of the structural shift in the two single-equation models is illustrated in Fig. 2.

The gradual switching model allows not only the timing and speed of structural shifts to be identified but also the nature of the shifts to be characterized. In both models, the results suggest that the negative influence of the stocks-to-use ratio is considerably stronger in the latter period. In Model 1, the

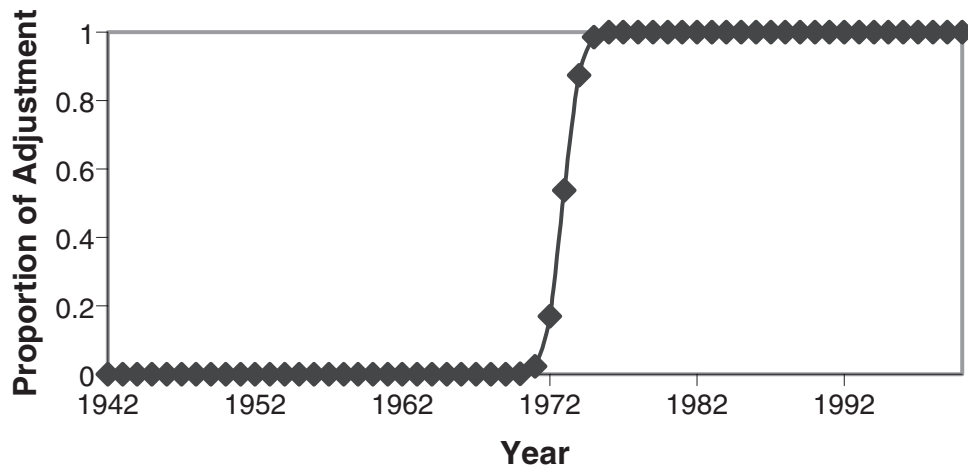
<sup>3</sup> In reality, all observations fall between regimes given the asymptotic nature of the transition function, which never actually reaches zero from above or one from below.

<sup>4</sup> Note that the parameter on the producer price index is not allowed to shift. Estimates of such a specification were numerically unstable.

**Table 2. Estimates of gradual switching soybean price model**

Variable	Model 1	Model 2	Model 3
$\mu$ (mid-point observation)	30.9078 (0.2384)*	30.8547 (0.1914)*	30.7426 (0.1965)*
$\sigma$ (speed of adjustment)	0.9525 (0.3243)*	0.8646 (0.2655)*	0.8784 (0.2738)*
Intercept <sup>I</sup>	2.0395	2.0885	2.1815
Loan rate <sup>I</sup>	0.8285	0.8349	0.6911
(Stocks <sub>1</sub> /Use <sub><i>t</i>-1</sub> ) <sup>I</sup>	-0.4109	-0.4183	-0.2933
Corn (Stocks <sub>1</sub> /Use <sub><i>t</i>-1</sub> ) <sup>I</sup>	—	—	-0.2545
Intercept <sup>II</sup>	4.9449	4.5554	3.7974
Loan rate <sup>II</sup>	-0.0851 (0.1079)	-0.0057 (0.1106)	-0.0907 (0.1058)
(Stocks <sub>1</sub> /Use <sub><i>t</i>-1</sub> ) <sup>II</sup>	-0.6829 (0.1469)*	-0.6119 (0.1570)*	-0.4659 (0.1694)*
Corn (Stocks <sub>1</sub> /Use <sub><i>t</i>-1</sub> ) <sup>II</sup>	—	—	-0.1207 (0.0979)
Drought	—	0.0788 (0.0695)	0.1031 (0.0623)*
LDP	—	-0.2595	-0.2574 (0.0493)*
FPPI <sub><i>t</i>-1</sub>	—	-0.0055 (0.1160)	0.1649 (0.1247)
Adjusted $R^2$	0.9538	0.9697	0.9757

Note: Regime I data are for 1942–1972; Regime II data is for 1973–2000. Stocks<sub>1</sub> = 1st quarter (1 December) stocks. Numbers in parentheses are standard errors. Asterisks indicate statistical significance at the  $\alpha = 0.10$  or smaller level.

**Fig. 2. Estimated transition function for single equation model**

coefficient changes from  $-0.41$  in the early regime to  $-0.68$  in the latter regime. Likewise, in Models 2 and 3, the shifts are from  $-0.42$  to  $-0.61$  and  $-0.29$  to  $-0.46$ , respectively.

The effect of loan rates on soybean prices also appears to vary from period to period. In the first regime, the coefficient on loan rates is statistically significant with a value of about 0.70 to 0.83. In the second regime, loan rates do not appear to have influenced prices. The addition of discrete indicators for drought and the marketing loan programme and the inclusion of the producers' price index as an indicator of general price movements do not appear to significantly alter these the results. When local market prices fall below the loan rate, the marketing loan programme allows producers to capture the price difference as a payment from the government. Prior to implementation of the marketing loan

programme, when market prices fell below the loan rate farmers would cede their crops to the government in return for the loan rate. Thus, the marketing loan programme prevents the loan rate from acting as a floor for market prices. This negative effect on average market prices is captured by the LDP variable. Finally, relative corn stocks have a negative effect on soybean prices, though the effect is only significant in the first regime. Again, this may reflect the considerable degree of correlation between corn and soybean stocks.

In summary, the results are largely consistent with the findings of earlier research. A structural shift does indeed appear to have characterized market price relationships in the reduced form model of soybean farm prices. The shift appears to have occurred at about 1972–1973 and appears to have been very rapid.

## V. A Quarterly Dynamic Model

The preceding analysis provides valuable inferences regarding the presence, timing, and speed of structural adjustments in soybean annual market price relationships. It is, however, also of interest to consider how such changes may have affected dynamic relationships in the short run. To this end, a dynamic vector autoregressive (VAR) model for prices and the stocks-to-use ratio that incorporates the gradual switching methods described above is considered. Goodwin (1992) undertook a similar analysis of commodity prices. The model is applied to quarterly data covering the period from 1964 to 2000.

If  $Y_t$  is defined as a  $2 \times T$  matrix containing prices and the stocks-to-use ratio, a standard vector autoregressive model is given by:

$$Y_t = A_0 Z_t + A_1 Y_{t-1} + \dots + A_k Y_{t-k} \quad (8)$$

where  $Z_t$  represents any deterministic components of the model, including an additive intercept. If the VAR model is expressed as  $Y_t = \Gamma(Z_t, \dots, Y_{t-k})$ , a switching version of the model developed using the same approach as that applied above is given by:

$$Y_t = (1 - \lambda_t)\Gamma_1(Z_t, \dots, Y_{t-k}) + \lambda_t\Gamma_2(Z_t, \dots, Y_{t-k}) \quad (9)$$

Again, the mixing parameter  $\lambda$  reflects the adjustment between regimes. In contrast to the single equation model evaluated in the preceding section, the VAR model is comprised of two equations—one for prices and another for the stocks-to-use ratio. It is assumed that the entire VAR model (i.e., both equations) shifts simultaneously. (Note, to the extent that the VAR models represent reduced form expressions of dynamic supply and demand conditions, it would

be expected for structural shifts to affect each equation in the same fashion, i.e., at the same point and rate of adjustment.)

NASS quarterly stocks data and quarterly average cash soybean prices (at Central Illinois) are utilized taken from USDA sources. A complication relates to the fact that total use statistics are only available on an annual basis. To construct a quarterly use series, cubic spline interpolation is utilized which smooths and interpolates the annual data to obtain quarterly series. Of course, such an approach does not capture quarterly shocks that may reflect seasonal consumption. Since the goal is to represent longer run changes in a relatively stable series—total use—such a smooth series is preferred. Quarterly dummy variables are included to capture seasonal influences on prices and the stocks-to-use ratio. In addition, the loan rate is included as an exogenous variable. The loan rate is constant across a marketing year and thus the same value is used for each appropriate quarter of a single marketing year. All continuous variables were considered in logarithmic forms.

The quarterly VAR model was estimated using the grid search technique described above. The results indicated a very rapid break at observation number 43, representing the third quarter of 1974. The estimated transition function is illustrated in Fig. 3. Thus, the switching, dynamic, quarterly VAR model estimates are largely consistent with the results obtained for the annual data above.

An evaluation of the nature of dynamic relationships inherent in estimates of a VAR model is best pursued using impulse response analysis. The data were divided at observation 43 and VAR models reestimated for each regime. This approach is justified by the very rapid nature of the structural break

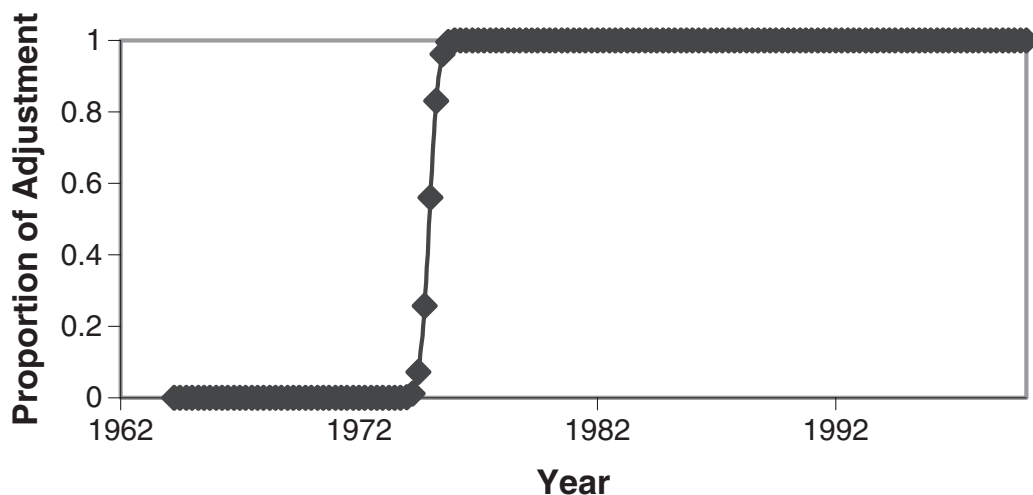
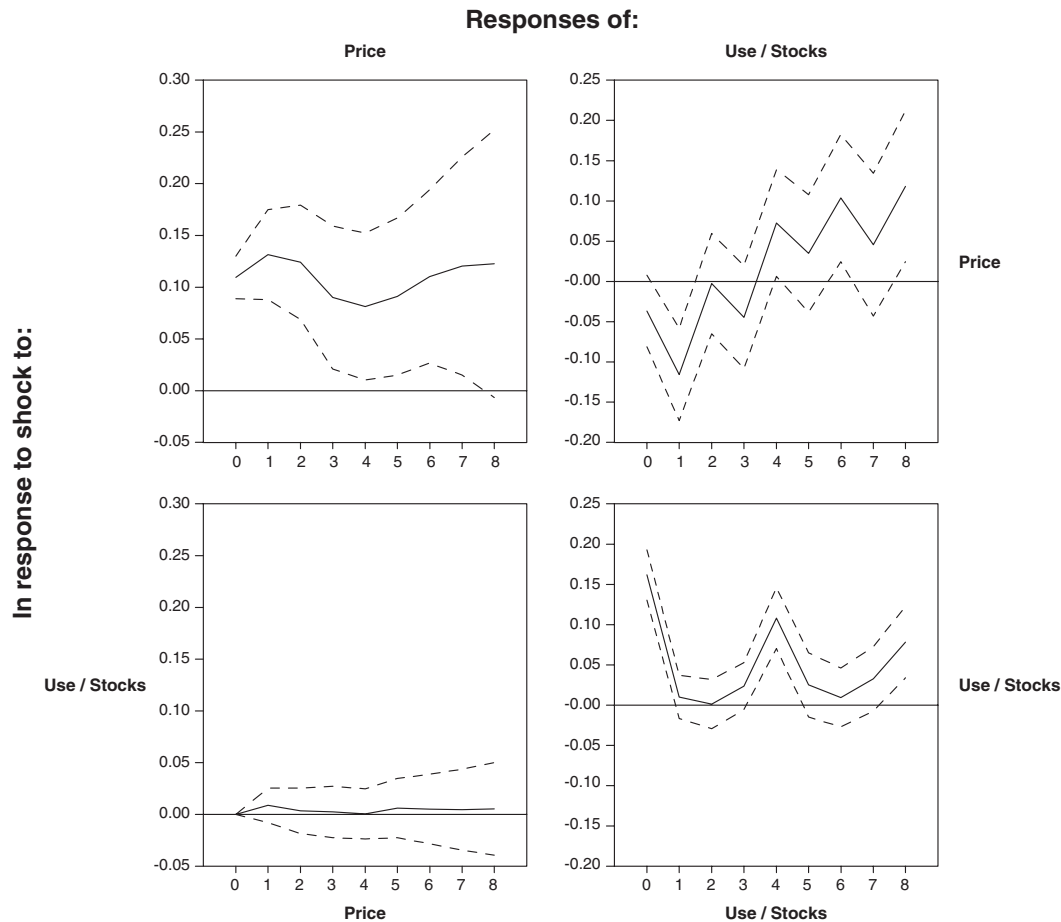


Fig. 3. Estimated transition function for VAR model





**Fig. 4. Regime I impulse responses**

at this observation (i.e., the estimate of  $\sigma$  was 1 in the VAR model). Parameters for each alternative model were used to consider orthogonalized impulse responses to one-standard-deviation shocks to prices and the stocks-to-use ratio.

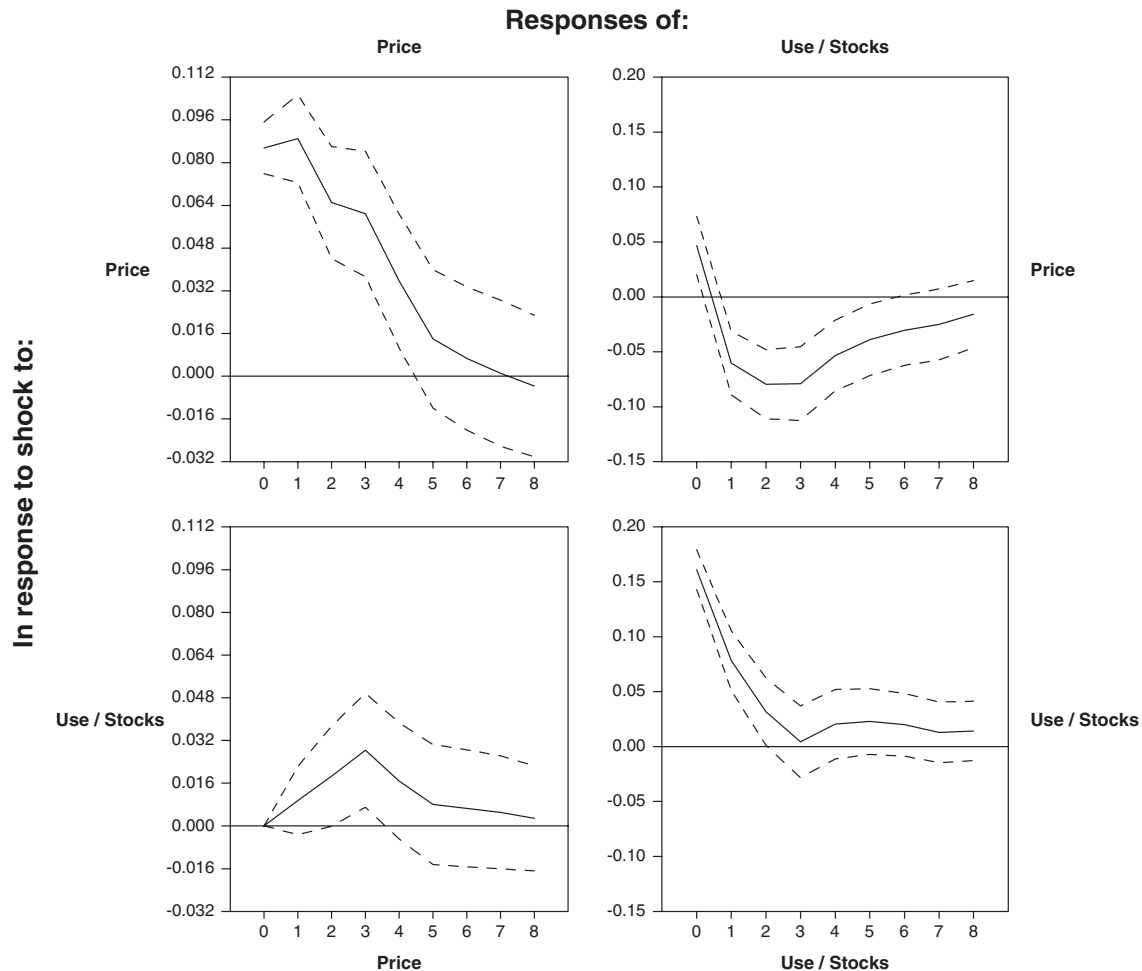
Figure 4 presents impulse responses (with 90% confidence bands) to shocks to soybean prices and the ratio of total stocks-to-use, respectively, in the early regime.<sup>5</sup> The results indicate that soybean prices do not appear to be significantly affected by shocks to stocks, though stocks do appear to react negatively in the period immediately following a positive shock to prices. Estimates for the first regime may reflect a degree of instability in the responses. This may reflect the limited number of observations available for estimating the VAR parameters for the first regime.

Figure 5 illustrates impulse responses in the second regime. In this case, the responses are much different. Figure 5 indicates that a positive shock to prices

significantly lowers the stocks-to-use ratio though this negative effect is preceded by a small positive response. The effects of this shock persist for several months. Conversely, prices appear to temporarily increase in response to a shock in the stocks-to-use ratio. The response is significant only in the second and third quarters following the shock. This may reflect seasonal patterns that are not fully captured by the quarterly dummy variables that are included in the VAR models.

In all, the results for the quarterly VAR model are largely consistent with those reported for the annual models. A strong and abrupt structural break appears to have affected quarterly reduced form relationships between prices and the stocks-to-use ratio. The timing of the break is very similar to what was revealed for the annual data, suggested a break around 1975. Other research has identified the multiple-year period from 1972 through 1976 to represent a distinct break

<sup>5</sup> Monte Carlo integration was used to provide confidence bands, which are given by the response plus/minus 1.64 times the estimated standard error.



from previous price variability patterns (Schnepf and Goodwin, 1999).

## VI. Concluding Remarks

An understanding of fundamental reduced form relationships among variables important to supply and demand and market prices is important to commodity and policy analysts. This paper reports on an analysis of such market relationships for soybeans. Following earlier research, a simple regression model was considered for annual soybean prices that included the stocks-to-use ratio, the loan rate, and a number of discrete indicators of policy. Two distinct issues were pursued in the consideration of this relationship.

The first involves explicit modelling of structural change. A primary focus of the analysis involved the identification and characterization of structural shifts. Models of discrete structural breaks were utilized as well as an alternative gradual switching regression

approach that permits change to occur gradually. The results confirm the significance of an abrupt structural break that occurred at about 1973–1974. The timing and speed of the adjustment were robust over a number of alternative specifications. The results suggest that soybean prices have become more sensitive to relative stocks (i.e., the ratio of stocks to use).

A second focus of the analysis involves the potential endogeneity of the stocks-to-use ratio and prices. Theoretical considerations of stockholding behaviour suggest that stocks will be affected by prices. Likewise, total use should be negatively influenced by prices. Explicit tests of this endogeneity have been conducted and it was confirmed that significant biases may arise if the endogeneity of the stocks-to-use ratio is ignored in a reduced form price equation.

A dynamic VAR analysis of quarterly soybean prices was also considered. The results of this analysis are largely consistent with those obtained with the annual data. A significant structural break appears to have occurred in 1974. Again, the structural break was very quick.

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