Acreage Response Under Farm Programs for Major Southeastern Field Crops

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Abstract

An expected utility model that includes output price and yield uncertainty was used to estimate cotton, corn, and soybean acreage response equations for the Southeast. The model appeared to fit the soybean and corn data well, resulting in own-price elasticity estimates of 0.317 for corn and 0.727 for soybeans. When applied to cotton acreage, however, the model did not yield satisfactory results. When elasticity was allowed to change over time, however, statistical results for the cotton equation improved, yielding an own-price elasticity of 0.915 at data means.

Key Words: Government Programs, Acreage Response, Expected Utility, Time-Varying Parameters

For many crops, the estimation of supply response is complicated by the existence of government programs. Traditionally, farm programs have provided farmers with price guarantees and/or subsidies in exchange for limitations on planting. In this manner, the programs affect both the expected returns and the variance of these returns. Even for crops, such as soybeans, in which direct government involvement is minimal, farm program provisions can strongly affect acreage response in an indirect fashion by making alternative crops either more or less attractive to the producer.

The primary purpose of this paper is to estimate the supply response of three major government program crops grown in the Southeast: cotton, corn and soybeans. Most recent estimates of supply-response elasticities have been at the national level, while little recent empirical work has focused on supply response elasticities for the Southeast. Because the Southeast differs significantly from the rest of the nation in terms of climate and soils, elasticity estimates developed for the country as a whole may not be accurate in this region for predicting acreage response to changing farm program provisions or market conditions.

To the best of our knowledge, no previous empirical work has examined the issue of changing supply response over time. Because of changes in technology and changes in farm program provisions, it is possible that supply schedules have changed over time, becoming either more or less elastic. Accordingly, in this paper, the possibility of time-varying supply response is investigated, as a secondary objective.

Basic Model Formulation

The general behavioral model assumed here is the one hypothesized by Chavas and Holt. Because a full discussion of this model is provided in their article, only a brief summary is provided here.[1] First, it is assumed that the farm household

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has preferences represented by a von Neumann-Morgenstern utility function, and that the household maximizes expected utility subject to a budget constraint in which income is determined both by nonfarm sources (or wealth) and net returns from farming. These assumptions lead to a maximization problem expressed as:

$$(1)\quad \underset{A_i}{\text{Max}} \{ EU[w + \sum \pi_i A_i] \} \text{ s.t.}$$

$$(2)\quad f(A) = 0$$

where EU is expected utility, A_i is the number of acres devoted to the ith crop, A is a vector representing all the A_i, w is normalized initial wealth, π_i is normalized profit per acre of the ith crop, and the constraint serves to limit plantings to acreage available. For normalization, all prices are deflated by a price index.

The per acre profits, π_i, depend on price, yield, and cost. Of these, price and yield are unknown at the time decisions are made. The expectation in (1) therefore must be based on the information available when decisions are made. Further, if the household is not risk neutral, optimal acreage decisions will depend not only on expected normalized profits, but also on higher moments of the profit distributions, so that A^*, the vector of optimal acreage decisions, can be expressed as $A^*(w; \bar{\pi}; \sigma)$, where $\bar{\pi}$ is the expected profits vector, and σ represents higher moments of the profit distribution.

From (1) and (2), and with reference to the work of Sandmo, Chavas and Holt work out the implications for econometric estimations of acreage response. By considering the compensation function, C, defined implicitly by:

$$(3)\quad \underset{A_i}{\text{Max}} \{ EU[w + C + \sum \pi_i A_i] = U^0 \}$$

where C is the compensation (change in wealth) needed to keep utility constant at U^0, the following symmetry restrictions can be derived (Chavas):

$$(4)\quad \partial A^c / \partial \bar{\pi} = \partial A^* / \partial \bar{\pi} - (\partial A^* / \partial w) \cdot A^*$$

where A^c is the wealth-compensated acreage decision found by solving (3). The matrix of compensated effects $\partial A^c/\partial\bar{\pi}$ in expression (4) is symmetric and positive semidefinite (Chavas; Chavas and Holt). If the wealth effect is zero, $\partial A^*/\partial\bar{\pi}$ is symmetric.

Incorporating Government Policy into the Return Function

To estimate the model posited by Chavas and Holt, a consistent set of policy variables must be developed for Southeastern field crops. While some analysts have preferred to estimate separate supply or acreage response functions for different policy regimes (Morzuch et al.; Lee and Helmberger), others have followed the method outlined by Houck and Subotnik, defining an effective support price (PS^e) in a general enough manner that alternative policy regimes can be represented by one variable. The effective support price depends on both the announced support price and the restrictive conditions (r) required of the farmer for program participation:

$$(5)\quad PS^e = r \cdot PA$$

where PA is the announced government price (loan rate or target price). The adjustment factor, r, embodies planting constraints. When the government price is available without restrictions, r=1. As restrictions become tighter, r moves toward 0. In some instances, r is relatively easy to calculate. Other times, the calculation of r is more difficult. Because soybean programs have involved only a loan program, with no acreage restrictions, the effective support price for soybeans is the loan rate itself.

Chavas and Holt stated that they followed Gallagher's general methods in developing the effective price support for corn, but did not provide details of the series construction; nor, of course, did they develop a series for cotton. A detailed description of the development of an effective support price series for cotton, also based on Gallagher's general methods, is found in Duffy et al. This series was used here, with updates for the years since their series ended. For consistency, the corn effective support price series used here was developed using the guidelines established in Duffy et al.

In the literature, expected market price has been calculated in a number of ways. One method involves simply using a one-period lag (Duffy et al.). Alternatively, a more complicated lag structure may be used (Shumway). Chavas and Holt used one-period lagged price, plus a constant, where the constant was the mean sample difference between current and lagged prices. Implicitly, the Chavas and Holt specification of expectation at time t-1, E_{t-1}, of the normalized market price at time t, P_t, is defined by the equation:

$$(6)\quad E_{t-1}(P_t) = \alpha + \beta P_{t-1}$$

with β constrained to equal 1.

There is no theoretical reason to justify restricting β to 1. Hence, in this study, rather than restrict (6), direct estimation was employed to find the value of β. Additionally, because real prices have trended downwards over time (corresponding to a downward trend in real per unit costs of production), a trend variable (T) was included in the equation to be estimated, so that:

$$(7)\quad E_{t-1}(P_t) = \alpha + \beta P_{t-1} + \gamma T$$

was directly estimated using OLS. Both linear and double-log versions of (7) were estimated. In terms of predictive power as measured by R-square and mean square error, the double-log regression outperformed both the linear regression model and the constrained expression in (6). Thus, the double-log version of (7) was used to generate expected market prices in this study.[2]

Various methods have been used to incorporate the effective support price and the expected market price into one "supply-inducing" price. Shumway, for example, chose the higher of the effective support price or the expected market price. Bailey and Womack, among others, used a weighting scheme based on government program participation. Duffy et al. used an alternative nonlinear weighting scheme first proposed by Romain. Chavas and Holt, however, follow a substantially different scheme. In this paper, the Chavas and Holt method is employed for incorporating government programs into price expectations.

Unlike the ad hoc methods used by others, the Chavas and Holt method is grounded in previous statistical work on the effects of truncating the normal distribution (Johnson and Kotz, Maddala).[3] Because government programs essentially provide a minimum price, they serve to truncate the distribution of expected prices received by farmers. Thus, the truncated distribution has both a different mean and a different variance from those associated with the untruncated distribution.

The variance of the untruncated normalized prices, P_t, was defined following Chavas and Holt:

$$(8)\quad \mathrm{VAR}(P_{it}) = \sigma_{pii,t} = \sum_{j=1}^{3} \lambda_j \, [P_{i,t-j} - E_{t-j-1}\,(P_{i,t-j})]^2$$

where the weights, λ_j, are .5, .33, and .17, and t is a time subscript. This result, along with (7), was then used to find the mean and variance of the truncated price distribution. Letting $\bar{\mathrm{p}}_i$ represent the expected mean price of the ith crop from the truncated distribution and $\bar{\sigma}_{pii}$ represent the variance of the truncated distribution:

$$(9\text{a})\quad \bar{\mathrm{p}}_{it} = PS_{it}^{e}\,\Phi(h_i) + \sigma_{pii,t}^{1/2}\phi(h_{it}) + P_{it}[1 - \Phi(h_{it})]$$

$$\begin{aligned}(9\text{b})\quad \bar{\sigma}_{pii,t} = {} & (PS_{it}^{e})^2\Phi(h_{it}) + \sigma_{pii,t}h_{it}\phi(h_{it}) \\ & + 2P_{it}\sigma_{pii,t}^{1/2}\phi(\mathrm{h}_{it}) + (P_{it}^{2} \\ & + \sigma_{pii,t})*(1 - \Phi(h_{it})) - \bar{\mathrm{p}}_{it}^{2}\end{aligned}$$

with:

$$(9\text{c})\quad h_{it} = (PS_{it}^{e} - P_{it})/\sigma_{pii,t}^{1/2}$$

where $\Phi(\cdot)$ and $\phi(\cdot)$ are the standard normal density function and the distribution function respectively. The derivation of (9) is provided by Chavas and Holt.[4] Because expectations are not static across time, (9) must be computed for each year of the estimation period. The formula for covariance is also presented in Chavas and Holt, but not reproduced here.

Once the mean and variance of the truncated price distribution are calculated, an expression for expected profits can be derived. To get expected yields, Chavas and Holt used a regression on trend. In this study, expected yields were calculated in the same manner as expected prices, with the natural log of yield regressed on the natural log of lagged yield and a time trend. This formulation was found to give more accurate predictions of expected yield for the Southeast (as measured by R-squares and mean square errors) than were generated by a simple trend.

Following Chavas and Holt, expected profit of the ith crop in year t, $\bar{\pi}_{it}$, is defined as:

$$(10)\quad \bar{\pi}_{it} = E_{t-1}\{\, P_{it} \cdot Y_{jt} - C_{it} \mid P_{it} \geq PS_t^e \,\}$$

where, Y_{it} is yield, and C_{it} is normalized per acre costs (assumed known at planting). Because of covariance between yields and prices, (10) is calculated using:

$$(11)\quad \bar{\pi}_{it} = P_{it} \cdot Y_{it}$$

$$+ ((1 - \Phi(h_{it})) \cdot \rho_{yi,pi} \cdot \sigma_{yi} 1/2 \cdot \bar{\sigma}_{pi,t}{}^{1/2})) - C_{it}$$

where σ_{yit} is the standard deviation of yield, $\bar{\sigma}_{pi\,t}$ is the variance of the truncated price distribution, and $\rho_{yi,pi}$ is the correlation between yield and price.

Data and Estimation

Given the economic hypotheses in (1) and the formula for expected profit in (11), the optimal acreage equations $A^*(w, \bar{\pi}, \sigma)$ were then specified from a Taylor-series expansion of an arbitrary functional form. After substitutions, following Chavas and Holt, the form of the model to be estimated is:[5]

$$(12)\quad A_{it} = a_i + \alpha_i(w_{t-1} + \sum_{j=1}^{2} A_j\bar{\pi}_{jt})$$

$$+ \sum_{j=1}^{2} \beta_{ij}\bar{\pi}_{jt} + \sum_{j=1}^{2}\ \sum_{k=1}^{2} \gamma_{ijk}\bar{\sigma}_{pjk,t} + \tau T + u_{it}$$

Because the β_{ij} represent the compensated slopes, $\partial A_i^c/\partial\bar{\pi}_{ij}$, then the symmetry condition in (4) requires that $\beta_{ij} = \beta_{ji}$, for $i \neq j$. In the Chavas and Holt paper, only two crops, soybeans and corn were examined. In the Southeast, however, cotton, soybeans and corn are all widely grown. While the Chavas and Holt model can theoretically be extended to cover multi-crop situations, a practical difficulty lies in estimation. To extend the model to three crops would require an additional four parameters in each equation, exceeding the practical limit on the ratio of parameters to observations suggested by Belsley et al. In our study, therefore, three sets of equations were estimated: corn and soybeans, cotton and corn, and cotton and soybeans.

The Southeastern states considered in this analysis were Alabama, Georgia, North Carolina and South Carolina. Cotton, soybeans, and corn annual time-series data were used in the estimations. Because of the need for lagged information for some of the independent variables, data were collected for the period 1955-1988. In actual estimation, however, the dependent variables span the period 1958-1988. Acreages planted to each crop and the prices received by farmers were obtained from various issues of U.S. Department of Agriculture (USDA publications). The costs of production used in this study were variable costs of production, originally reported by Gallagher and Green and updated by Taylor. Yields per acre were obtained from USDA publications. The consumer price index, used to normalize all prices, is reported by the Bureau of Labor Statistics. Initial wealth, w_{t-1}, was measured by farmers' equity as reported in various issues of *Economic Indicators of the Farm Sector* weighted by the ratio of the state's acreage in the crops of interest to the national acreage.[6]

In estimation, the version of equation (12) used for acreage supply was modified to include a dummy variable for PIK programs (1983) and a measure of the effective diversion payments for cotton and corn. A similar modification was made in Chavas and Holt. As in Chavas and Holt, aggregated data are used, with all the attendant problems, although the extent of the aggregation problems may be lessened by using regional rather than national data.

Empirical Results

The parameter estimates for the acreage supply equations, with symmetry imposed, are reported in tables 1 through 3. Equations were estimated as SUR systems, with symmetry of cross-revenue effects imposed. Data were corrected for autocorrelation problems before final estimation, if

Table 1. Estimated Southeastern Corn and Soybeans Acreage Equations with Symmetry Imposed, 1958-1988

Notation	Independent Variables	Parameter Estimate Corn	Parameter Estimate Soybean
a_i	Intercept	7777944** (9.38)	-4210446** (-4.99)
π_1($/acre)	Corn expected revenue	12239.51** (2.35)	-13697.1** (-3.90)
π_2($/acre)	Soybeans expected revenue	-13697.1** (-3.90)	16298.3** (3.74)
$\bar{\sigma}_{11}$	Corn price variance	-1494889 (-0.84)	2518811 (1.39)
$\bar{\sigma}_{22}$	Soybeans price variance	177383.1** (5.15)	-121093** (-3.04)
$\bar{\sigma}_{12}$	Corn-Soyb. price covariance	54159.4 (0.10)	209856.9 (0.37)
$w + \Sigma_j A_j \pi_j$	Adjusted wealth	.0000655** (5.43)	.0001626** (14.10)
DPCN	Corn diversion payment	-1366694** (-4.36)	
DUM	PIK Dummy variable	-897953** (-2.16)	-818799* (-1.80)
T	Trend variable	-200934** (-11.51)	164007.6** (8.59)
Number of	observations:	31	31
	R-Square.	0.95	0.98
	D.W.	2.54	2.36

[a] *t*-statistics are in brackets below the parameters estimates. *R*-Square and Durbin-Watson are from single equation OLS estimates. Final equations estimated as SUR system with symmetry imposed. Dependent variables are acres of corn and soybeans in the Southeast (Alabama, Georgia, North Carolina, and South Carolina).

** Significant at the 95 percent (* 90 percent) level of confidence.

necessary. *R*-square and Durbin-Watson values reported in the tables are from the first-stage OLS estimates.

Statistical results for the corn and soybean model (table 1) were generally strong in terms of statistical significance of parameters and the overall model fit. Corn own-revenue elasticity (compensated) at the mean is 0.095, 40 percent higher than the value of 0.068 found by Chavas and Holt. Soybeans own-revenue elasticity (compensated) is 0.560, approximately double the value of 0.279 for national acreage computed by Chavas and Holt. Compensated own-price elasticities are 0.317 and 0.727 for corn and soybeans, respectively. The higher elasticities for the Southeastern region probably reflect the greater number of crop options available in the region, including cotton, and in some areas peanuts, tobacco, and horticultural crops. The wider availability of production substitutes for the Southeast, as opposed to the other regions of the country, would make producers more responsive to changes in profitability. In addition, given the number of alternative crops, soybeans are not as important in rotational considerations here as they are in the Midwest (see Mims et al.).

For the corn-soybean model, a test of the symmetry restriction yielded a borderline $F(1,43)=5.87$. Given the evidence of the Chavas-Holt paper in which the F value for this test was

Table 2. Estimated Southeastern Cotton and Soybeans Acreage Equations with Symmetry Imposed, 1958-1988

Notation	Independent Variables	Parameter Estimate Cotton	Parameter Estimate Soybeans
a_i	Intercept	2405586** (3.07)[a]	-2619675* (-1.99)
π_1($/acre)	Cotton expected revenue	2123.31 (1.21)	-2672.1 (-1.11)
π_2($/acre)	Soybeans expected revenue	-2672.1 (-1.08)	13316.44** (2.16)
$\bar{\sigma}_{11}$	Cotton price variance	762071.9 (0.12)	8182745 (0.76)
$\bar{\sigma}_{22}$	Soybeans price variance	-59358.5 (-1.59)	53394.7 (0.86)
$\bar{\sigma}_{12}$	Cotton-Soy. price covariance	1196737 (0.62)	-5903751* (-1.93)
$w + \Sigma_j A_j \pi_j$	Adjusted wealth	-0.0000268* (-2.01)	0.000255** (13.11)
DPC	Cotton diversion payment	-45404.10 (-1.71)	
DUM	PIK Dummy variable	-92393.2 (-0.32)	-1360257** (-2.37)
T	Trend variable	-36597.1 (-1.72)	90104.13** (2.57)
Number of	observations:	31	31
	R-Square.	0.88	0.97
	D.W.	0.97	1.38

[a] *t*-statistics are in brackets below the parameters estimates. *R*-Square and Durbin-Watson are from single equation OLS estimates. Final equations estimated as SUR system with symmetry imposed. Cotton equation corrected for serial correlation before final estimation.

** Significant at the 95 percent (* 90 percent) level of confidence.

highly significant, the restrictions were maintained here. With symmetry imposed, a test for risk neutrality was conducted through an *F*-test for all $\gamma_{ijk} = 0$ and $\alpha_i = 0$, yielding an $F(8,43) = 56.71$. Risk neutrality can thus be rejected at the 0.0001 level of significance. In both equations, the soybean-price variance is significant, while the corn-price variance is not. Because corn is protected by more extensive program provisions than soybeans, these results are not surprising. In both equations the wealth variable is positive and significant, indicating that corn and soybean farmers in the Southeast exhibit decreasing absolute risk aversion (DARA), the same finding that Chavas and Holt report at the national level.

The second model pairs cotton and soybeans (table 2). Statistical results are less satisfactory than in the corn-soybean case in that the cotton own-revenue parameter is not statistically significant at the usual levels of confidence. Nor are the cross-revenue effects significant. The wealth effect is positive and significant in the soybean equation, but negative and insignificant in the cotton equation, a finding contrary to the DARA hypothesis. The low Durbin-Watson for the cotton equation can indicate a problem with autocorrelation, which was corrected before final estimation, but may alternatively indicate improper specification, further evidence that the Chavas and Holt model does not explain acreage response for cotton.

Table 3. Estimated Southeastern Cotton and Corn Acreage Equations with Symmetry Imposed, 1958-1988

Notation	Independent Variables	Parameter Estimate Cotton	Parameter Estimate Corn
a_i	Intercept	2584156** [2.95][a]	2395257 [1.46]
π_1(\$/acre)	Cotton expected revenue	2846.43 [1.63]	1302.33 [0.54]
π_2(\$/acre)	Corn expected revenue	1302.33 [0.54]	8510.54 [1.18]
$\bar{\sigma}_{11}$	Cotton price variance	-7262616* [-1.85]	2720814 [0.41]
$\bar{\sigma}_{22}$	Corn price variance	-2730467 [-1.67]	-1647866 [-0.56]
$\bar{\sigma}_{12}$	Cotton-Corn price covariance	-2105924 [-0.97]	3387524 [0.87]
$w + \sum_i A_i \pi_i$	Adjusted wealth	-0.000046 [-1.49]	0.000157** [2.75]
DPCN	Corn diversion payment		-1540155** [-3.07]
DPC	Cotton diversion payment	-56443.2** [-2.31]	
DUM	PIK Dummy variable	-296596 [-0.86]	107372.9 [0.17]
T	Trend variable	-54357.8** [-2.68]	-47157 [-1.32]
Number of	observations:	31	31
	R-Square:	0.88	0.89
	D.W.	1.38	1.72

[a] *t*-statistics are in brackets below the parameters estimates. *R*-Square and Durbin-Watson from single equation OLS estimates. Final equations estimated as SUR system with symmetry imposed. Dependent variables are acres of cotton and corn in the Southeast (Alabama, Georiga, North Carolina, and South Carolina).

** Significant at the 95 percent (* 90 percent) level of confidence.

The test for symmetry in the cotton-soybean model yielded an $F(1,42) = 0.190$, indicating that symmetry cannot be rejected. As the cross-revenue effect is not significantly different from zero, this result is not surprising. The test for risk neutrality led to an $F(8,42)=23.71$, indicating that risk neutrality cannot be rejected.

The cotton and corn model (table 3) also yielded disappointing results in terms of *t*-values on important parameters. The test for symmetry in the cotton-corn model yielded an $F(1,43) = 2.94$, indicating that symmetry should not be rejected; however, given that the cross-revenue effects are not significantly different from zero, this test is not very meaningful. The test for risk neutrality yielded an $F(8,43) = 2.82$, a borderline value that contrasts sharply to the high *F*-values for the other models. Here, cotton and corn, two crops extensively covered by government programs, are paired. If the government programs are effective in reducing price-risk, it is not surprising that the acreages of these commodities would show little response to these price risk variables.

In general, results for the Southeast indicate that the expected utility model fits the corn-soybeans data fairly well, but not the cotton data.

This statistical lack of response may reflect a genuine characteristic of cotton producers, but it might alternatively be explained by the diversity of crops in the area. Given the many possible substitutes, it is difficult to isolate one alternative crop with a significant influence on cotton acreage in the aggregate. In Georgia, for example, cotton is most often grown on farms where soybeans, corn, and peanuts are also produced. Cotton can be used as a rotation crop with peanuts; thus, decisions regarding cotton planting may be affected by factors outside the cotton market. Because of the nature of the farm program provisions for peanuts, peanuts could not be easily modeled in the framework used here. Further work, probably involving farm-level decision models, is needed to trace the interaction of cotton and peanuts.

Another possible explanation for the lack of revenue response in the cotton equation involves the possibility of a response that has changed over time. A static parameter may test as insignificant because the supply schedule has rotated over time, becoming either more or less elastic. Some economists have argued that supply should be more elastic now than previously because of greater reliance on purchased inputs (Tomek and Robinson, p. 362); alternatively, mechanization and the reliance on crop-specific equipment could make the supply schedule less elastic.

Time-Varying Parameters

Time-varying parameter models can allow for both systematic and stochastic changes in parameter values. Systematic changes involve nonrandom changes in parameter values, while stochastic changes can take place either around a stationary or a nonstationary (time-varying) mean parameter value. Singh et al. developed a model in which regression coefficients are specified as stochastic functions of calendar time, so that:

$$\beta_{it} = \beta_i^* + f_i(t)'\varphi_i^* + \zeta_{it} \tag{13}$$

where β^* represents the "base" value of the parameter, $f_i(t)$ is a function vector containing time (t), φ_i^* is a parameter vector, and ζ_t is an error term. The Singh et al. model is intuitively appealing because time is used to represent unobservable or unmeasurable factors that systematically alter parameter values.

Incorporating (13) into the supply response equations, and assuming $f_i(t) = t$,[7] the equations to be estimated are:

$$\begin{aligned}(14)\quad A_{it} = {} & a_i + \alpha_i(w_{t-1} + \sum_{j=1}^{2} A_j\bar{\pi}_{jt}) \\ & + \sum_{j=1}^{2}(\beta_{ij}^* + \varphi_{ij}T)\bar{\pi}_{jt} + \sum_{j=1}^{2}\sum_{k=1}^{2}\gamma_{ijk}\bar{\sigma}_{pjk,t} \\ & + \tau T + \psi_{it}\end{aligned}$$

where $\psi_{it} = \zeta_{it}(\bar{\pi}_{1t} + \bar{\pi}_{2t}) + u_{it}$ is a heteroscedastic disturbance term. If the disturbance term is not, in fact, heteroscedastic the model reduces to Stone's dynamic model.

The model in (14) was estimated and the error terms were tested for heteroscedasticity of three possible forms using the Glesjer and the Breusch-Pagan tests. Results of the Glesjer test indicated no heteroscedasticity in any cases, while one of the three versions of the Breusch-Pagan indicated possible heteroscedasticity problems in cotton. Because the Glesjer test is more powerful, and because the sample size is small, we assumed homoscedasticity for the final estimation. (Greene indicates that correcting for heteroscedasticity can be more harmful than helpful when sample size is small.)

All three sets of equations were re-estimated allowing for time variance on the revenue effects. The most dramatic change in results occurred in the cotton-corn pair, table 4. Here, the own-revenue parameter for cotton was significant, as was the cross-revenue effect. Results indicate that over time acreage response has become more inelastic for cotton, perhaps because investment in machinery has increased. A compensated own-revenue supply elasticity for cotton of 0.570 at the mean (with t = 16) was computed using the time-varying parameter model results. The compensated own-price elasticity from this model was 0.915.[8] The Durbin-Watson statistic for the cotton equation, which had previously been low, is now

Table 4. Estimated Southeastern Cotton and Corn Acreage with Symmetry Imposed and Time-Varying Parameters, 1958-1988

Notation	Independent Variables	Parameter Estimate Cotton	Parameter Estimate Corn
a_i	Intercept	2079705**	4204737
		[3.53]	[2.51]
π_1(\$/acre)	Cotton expected revenue	6728.49**	-9302.82**
		[4.21]	[-2.68]
π_2(\$/acre)	Corn expected revenue	-9032.82**	22719.76
		[-2.68]	[1.43]
$\bar{\sigma}_{11}$	Cotton price variance	-4303940	395481.5
		[-1.67]	[0.06]
$\bar{\sigma}_{22}$	Corn price variance	-1174110	-3773084
		[-1.08]	[-1.21]
$\bar{\sigma}_{12}$	Cot-Corn price covariance	-820849	1821185
		[-0.62]	[0.47]
$w + \sum_i A_i \pi_i$	Adjusted wealth	-.000051**	.000190**
		[-2.68]	[3.31]
DPC	Cotton diversion payment	-51890.7**	
		[-3.58]	
DPCN	Corn diversion payment		-1332514**
			[-2.41]
DUM	PIK Dummy variable	-241711	-87479
		[-1.10]	[-0.12]
$\pi_1{}^*T$	First interaction term	-225.04**	408.03**
		[-3.13]	[2.49]
$\pi_2{}^*T$	Second interaction term	408.03**	-569.00
		[2.49]	[-0.70]
T	Trend variable	-26114.8	-128489**
		[-1.42]	[-2.53]
Number of	observations:	31	31
	R-Square:	0.96	0.90
	D.W.	2.05	1.81

[a] *t*-statistics are in brackets below the parameters estimates. *R*-Square and D.W. from OLS estimates. Equations estimated as SUR system with symmetry imposed.

** Significant at the 95% (* 90%) level of confidence.

approximately 2.00. If this statistic was, in fact, signifying an underlying problem with specification rather than autocorrelation, it would appear that the problem has been corrected through allowing time-varying elasticities of supply.

Incorporating time-varying parameters into the corn-soybean model yielded very poor results in terms of significance on the revenue parameters and the interaction terms, although the expected signs were largely maintained. The cotton-soybean pairing under time-varying parameters for the revenue variables was also somewhat disappointing, although the own-revenue parameter for cotton was positive and significant.

Conclusions

The present paper focused on corn, soybeans, and cotton acreage decisions in the Southeast. Systems of acreage equations under expected utility maximization were developed following a general model proposed by Chavas and Holt. The application of the model to a three crop system was limited by the problem of overparameterization, so that the model had to be fitted in"piecewise" two by two fashion.

Overall, our study results indicate that the model fits the corn-soybean data fairly well, but that cotton acreage cannot be adequately modeled in this framework. The regional corn-soybeans model generally mirrored results found by Chavas et al. at the national level, but the estimated elasticities were considerably higher, indicating that Southeastern farmers are more responsive to changes in profitability.

Neither the cotton-corn or cotton-soybeans models gave satisfactory results in terms of significant revenue parameters. In addition, the cotton-corn model showed little evidence of risk aversion on the part of producers. Given the extensive government program provisions to reduce price risk for these commodities, these results are not surprising.

When the hypothesis of changing supply response over time was tested, the cotton equation yielded evidence that elasticity of supply had been decreasing over time. For corn and soybeans, no indication of time changes in parameter values was found. One explanation for this phenomenon is that machinery for cotton has become highly specialized, while corn and soybean equipment has remained largely interchangeable. Thus, one would expect over time to see a reduction in acreage response for cotton, but not for the other crops.

Overall, our results indicate that risk variability of soybeans appears to affect acreage of soybeans and corn, and possibly cotton, but that price variability in corn and cotton has little effect on planting decisions. Because the extensive farm program provisions for corn and cotton are largely designed to mitigate against the effects of market price volatility, these results indicate that the programs are working to that end.

Further research in time-varying supply elasticity for other crops and other regions of the country is warranted given our results for cotton. In particular, crops such as cotton in which machinery complements have become more specialized should be investigated.

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Endnotes

1. The authors gratefully acknowledge the assistance of Matt Holt, who provided us with copies of the computer programs used in estimating the Chavas and Holt model. These programs, along with additional information he provided, were extremely helpful and saved us a great deal of effort in estimating the models presented here.

2. Expectation functions for prices and yields were "well-behaved" in that they had reasonable *R*-square values and exhibited no problems with nonstationarity, as tested by Durbin-*h* test and Durbin's alternative test.

3. Normality of prices could not be rejected using the Kolmogorov-Smirnov test.

4. The formulas presented here for expected prices and variance are not found in exactly this form in Chavas and Holt, who presented a more general version of the effects of truncation. These formulas were obtained by private communications with Matt Holt.

5. We note that this specification of wealth involves use of the dependent variables, leading to possible simultaneous equation bias. An alternative model, using instrumental variables to calculate the wealth variable, was also estimated. Results were not affected to any noticeable agree. To keep our results most readily comparable to those of Chavas and Holt, we retain the original specification for reporting purposes.

6. This specification follows the Chavas and Holt method of scaling wealth by acreage.

7. A simple linear function was selected because more complicated functions involving multiple terms would reduce the degrees of freedom in the equations to a point where results are not reliable.

8. For the time-varying parameter estimates, the own-revenue response becomes negative in the last two years of the data period (t = 30 and t = 31). The problem, here, is likely related to the linear specification of the time function.