# RATIONAL EXPECTATIONS IN AMERICAN AGRICULTURE, 1867–1914

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## I. Introduction

DESPITE the historical importance of the Populist uprising in the late nineteenth century, the causes of the agrarian unrest which culminated in that movement remain obscure. Economic historians have become increasingly dubious of the justifications advanced by the Populists themselves, partly because many of the Populist programs (such as the Free Silver plan) were little more than schemes for an involuntary redistribution of wealth in favor of the farmers, and partly because many of the farmers' stated grievances fail to appear in the aggregate economic statistics of the period (Bowman, 1965; DeCanio, 1974a; Bowman and Keehn, 1974; North, 1974). The most comprehensive recent study concludes that "the agrarian protest of the late nineteenth century was not a simple, straightforward consequence of economic factors as many economic historians have believed" (Klepper, 1974, p. 285). There is some evidence that cyclical economic fluctuations coincided with upsurges of protest (Bowman and Keehn 1974; Klepper 1974), but the agrarian spokesmen of the time (as well as subsequent historians) attempted to identify long-standing structural problems of the agricultural sector as the ultimate explanation of the farmers' distress.

One recurring theme in historical explanations of the agrarian unrest locates the source of the farmers' difficulties in their perceptions of and responses to the economic requirements of the market. Thus, Mayhew suggests that the protests of the Grangers and Alliancemen were a reaction to the *commercialization* of agriculture. This commercialization may have increased farm incomes, but it also made the farmers subject to impersonal market forces (Mayhew, 1972). Farmers' failure to understand the operation of the markets for their products is featured in textbook accounts of the Populist period (Davis, Hughes, and McDougall, 1969, p. 368; Gray and Peterson, 1974, p. 320; and North, 1974, p. 134).

Econometric studies of the price-responsiveness of late nineteenth and early twentieth century American agriculture have shown that the sector as a whole responded properly to market prices in both the choice of crop mix and the choice of technique (Nerlove, 1958; Fisher and Temin, 1970; Hayami and Ruttan, 1971; DeCanio, 1973). These investigations of farmers' responses to output and input prices, however, do not indicate whether the estimated agricultural response parameters were in any sense *optimal*, nor can their fixed-parameter estimation techniques reveal whether the farmers' behavior *changed* in an appropriate manner as underlying market conditions changed.

This paper goes beyond the previous studies by testing directly a *rational expectations hypothesis* for American agriculture during the Populist period. Using a varying-parameter estimation method[1] it is possible to trace changes in the supply response parameters over time, and to compare those parameter variations with the variations implied by a model of rational price expectations. We will show that changes in the farmers' price expectations were indeed consistent with the theory of rational expectations. Since our estimates are based on the same statewide aggregate data used in the previous supply response studies, it is not possible to conclude from our results that *all* farmers formed rational expectations, but the existence of rational expectations in the aggregate leads us to reject the notion that farmers *as a group* were unable to

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[1] Recent developments in the theory of models with varying parameters are discussed in Cooley (1971) Rosenberg (1973) and Cooley and Prescott (1973a, 1973b, 1973c, and 1976).

understand the mechanisms of price formation in the world markets for their agricultural commodities in the late nineteenth and early twentieth centuries.

## II. The Model

The model of rational expectations which forms the basis of our test is the one proposed by Muth (1960). Suppose that the proportion of acres planted to a particular crop depends on the expected relative prices of the crop and its major alternatives:

$$S_t = \alpha_1 + \alpha_2 P_t^e \tag{1}$$

where $S_t$ denotes the share of total acreage devoted to the crop in question and $P_t^e$ is the expected ratio of the price of the crop to an index of the prices of the alternative crops. Then the farmer's decision problem is to predict the relative price so as to maximize net income. Following Muth, we assume that the relative price observed by the price-taking farmer is the sum of a permanent component $P_t^p$ and a transitory component $\eta_t$:

$$P_t = P_t^p + \eta_t. \tag{2}$$

This very general stochastic process is appropriate for describing movements in the price of an agricultural commodity produced and sold in competitive markets. We shall assume that the transitory components are independently and identically distributed with mean zero and variance $\sigma_\eta^2$. The permanent components can be assumed to follow a moving average process

$$P_t^p = P_{t-1}^p + \epsilon_t \tag{3}$$

where the $\epsilon$'s are independently distributed with mean zero and variance $\sigma_\epsilon^2$. The essence of the farmer's decision problem is to forecast the price at time $t$ (the harvest-time price) given the information available up through $t-1$ (the time at which the planting decisions must be made). The price expectation $P_t^e$ which minimizes the error variance $E(P_t - P_t^e)^2$ given the information up to time $t$ is

$$P_t^e = \Sigma_{k=1}^{\infty} (1-\lambda)\lambda^{k-1} P_{t-k} \tag{4}$$

where $\lambda$ depends in a known way on the variances of the permanent and transitory components of the price[2]

$$\lambda = 1 + (1/2)(\sigma_\epsilon^2/\sigma_\eta^2) - (\sigma_\epsilon/\sigma_\eta)[1 + (1/4)(\sigma_\epsilon^2/\sigma_\eta^2)]^{1/2}. \tag{5}$$

A Koyck transformation after the substitution of equation (4) into equation (1) leads to an empirical relationship of the form

$$S_t = \alpha_1(1-\lambda) + \alpha_2(1-\lambda)P_{t-1} + \lambda S_{t-1}. \tag{6}$$

This producers' response function has basically the same form as the equations estimated by Fisher and Temin (1970) and DeCanio (1973). It is one version of the dynamic adjustment model developed by Nerlove (1958). The general form of such lagged adjustment models also allows for delays in farmers achieving their desired crop mix, so that the actual acreage share, $S_t$, in equation (1) is replaced by the desired acreage share, $S_t^*$. Adjustment of the actual to the desired share may be described by a general mechanism

$$S_t = \Sigma_{i=1}^{\infty} [\theta_i S_{t-i} + \mu_i(S_t^* - S_{t-i})]. \tag{7}$$

Elimination of the unobserved $S_t^*$ leads to an empirical relationship of the form

$$\begin{aligned} S_t = {} & \alpha_1(\Sigma\, \mu_i)(1-\lambda) + \alpha_2(\Sigma\, \mu_i)(1-\lambda)P_{t-1} \\ & + [\lambda + (\theta_1 - \mu_1)]S_{t-1} \\ & + \Sigma_{i=2}^{\infty} [(\theta_i - \mu_i) - \lambda(\theta_{i-1} - \mu_{i-1})]S_{t-i}. \end{aligned} \tag{8}$$

Theoretical attempts have been made to relate the lagged adjustment parameters to the costs of being out of equilibrium, but these attempts have not been as satisfactory as models of rational expectation formation (Griliches, 1967, pp. 42–43). The data that would be required to chart the course of optimal modifications of the $\theta_i$ and $\mu_i$ parameters (such as data pertaining to the costs of shifting between crops) are not presently available at the state-wide level of

[2] It is not necessary to assume that $\epsilon_t$ and $\eta_t$ are uncorrelated. If $E\epsilon_t\eta_t = \sigma_{\epsilon\eta}$ and $E\epsilon_t\eta_s = 0$ $(t \neq s)$, it is only necessary to replace the ratio $\sigma_\epsilon^2/\sigma_\eta^2$ in equation (5) by $\sigma_\epsilon^2/(\sigma_\eta^2 + \sigma_{\epsilon\eta})$ (Muth, 1960, p. 304).

disaggregation. In addition, there are practical difficulties (chiefly multicollinearity) involved in obtaining precise and economically reasonable estimates of the coefficients of the later lagged terms, beginning with $S_{t-2}$.[3] For these reasons, we will concentrate our attention on the price expectations parameter, $\lambda$, provisionally treating the $\theta_i$ and $\mu_i$ as fixed *within* each state (but not necessarily across states). We also follow previous practice in omitting the lagged terms after $S_{t-1}$ from our empirical equations.

Thus the response functions estimated were of the form

$$S_t = \beta_1 + \beta_2 P_{t-1} + \beta_3 S_{t-1}. \tag{9}$$

A separate equation (with either cotton or wheat acreage in the numerator of $S_t$ and with $P_t$ equal to the ratio of either the cotton or the wheat price to an index of the prices of the major alternative crops) was estimated for each of the ten important cotton-growing states and seventeen wheat-growing states over a sample period beginning soon after the Civil War and ending just before World War I.[4] Our procedure is to assume that the parameters $\beta_i$ are described by a stochastic process and to estimate realizations of the process at specific intervals. Since the sample period spans nearly half a century, there is reason to expect variation in the structural parameters $\alpha_1$ and $\alpha_2$ and hence in $\beta_1$ and $\beta_2$ (David, 1971; Sahi and Craddock, 1974).[5] We identify variations in $\beta_3$ as resulting from changes in the price expectations parameter $\lambda$ (making use of the assumption that $\theta_1$ and $\mu_1$ are fixed). Changes in $\lambda$ should also lead to variations in $\beta_1$ and $\beta_2$, from equation (6) or (8). The results of section IV demonstrate a pattern of parameter variation which is quite consistent with Muth's model of rational expectation formation.

## III. Estimation Method

The estimation method used in this study has been developed in Cooley (1971) and Cooley and Prescott (1973c, 1976). The coefficients are assumed to be subject to both permanent and transitory changes over time:

$$\begin{aligned} \beta_t &= \beta_t^p + u_t \\ \beta_t^p &= \beta_{t-1}^p + w_t. \end{aligned} \tag{10}$$

The vector $\beta_t^p$ represents the permanent component of the parameters at time $t$. The $u_t$ and $w_t$ are independent and identically distributed random variables with zero mean vectors and covariance matrices which are specified as

$$\begin{aligned} \operatorname{cov}(u_t) &= (1-\gamma)\sigma^2 \textstyle\sum_u \\ \operatorname{cov}(w_t) &= \gamma\sigma^2 \textstyle\sum_w . \end{aligned} \tag{11}$$

[3] Fisher and Temin (1970) and DeCanio (1974b) were unable to obtain coefficients of $S_{t-2}$ which were significantly different from zero, and the actual estimated values of the coefficients of $S_{t-1}$ and $S_{t-2}$ together implied economically unreasonable values of the structural parameters.

[4] Here and throughout, all variables are in natural logarithms. The data and price indices we use are those of Fisher and Temin (1970) and DeCanio (1973 and 1974b), and the sources and reliability of the data are discussed in those references (see also footnote 5 below). The only exception is that the cotton price series were extended backward from 1882 to 1870 by substituting an average U.S. cotton price (United States Bureau of the Census, 1960, Series K303, p. 302) in place of the state cotton prices used after 1882. The correlation between the U.S. price and the state prices was quite high after 1882 because of the competitiveness of the national cotton market, and since this market was well developed before 1882 it is unlikely that any substantial error is introduced by use of the national price for the early years of the sample.

[5] Our varying-parameter specification also possesses an advantage in dealing with "structural change" in the model's parameters arising from systematic errors and inaccuracies in the data. Fisher and Temin have been criticized for ignoring the fact that the timing of winter wheat and spring wheat planting decisions is different. While the timing of the model they employed is appropriate for spring wheat, winter wheat was an increasingly important share of the total U.S. wheat crop during the 1866-1914 period (Higgs, 1971; Page, 1974). Fisher and Temin have replied, adequately in our view, that the close relationship between winter and spring wheat prices largely obviates this difficulty (Fisher and Temin, 1971, 1974). The December 1 price in calendar year $t$ will be closely correlated with the season average price of the wheat crop harvested in year $t$, and this harvest is largely complete at the time the winter wheat is planted in year $t$. (Harvest and planting dates are given in Nerlove (1958, pp. 137-138).) If the winter wheat problem did necessitate alterations over time of the price elasticity parameter $\beta_2$, our technique could take account of those changes while Fisher and Temin's fixed parameter model would not. Fisher and Temin's original article also argues that some inaccuracies in both the trend and level of the wheat acreage data may be present (1970, pp. 136-137). If the errors were unchanging, they would present no problem except in the accuracy of the constant term. However, if the errors of measurement changed systematically over time, our varying constant term would absorb the changes while a fixed constant (even augmented by a trend variable) could not. Cooley and Prescott (1973a, p. 254) report the results of Monte Carlo experiments showing the potential significance of a trend variable in models undergoing structural change which are estimated by fixed-parameter methods.

The matrices $\sum_u$ and $\sum_w$ specify the relative magnitudes of the parameter changes and are assumed known up to a scale factor. In the current analysis we choose

$$\textstyle\sum_u = \sum_w = I. \tag{12}$$

Extensive experiments were carried out with alternative specifications of $\sum_u$ and $\sum_w$, and the parameter histories traced out with these alternative covariance specifications were all very similar, with extremely high correlations between both the values of the parameters at different base periods (see below) and changes in the parameter values from one base period to the next. Comparisons of the Bayesian posterior odds did not indicate the superiority of any particular specification of the $\sum$ matrices (Zellner, 1971, pp. 291-302). Thus, the analysis presented below is quite robust with respect to alternative $\sum$ specifications.

The parameter $\gamma$ (which is constrained to lie between zero and one) specifies the relative variance of the permanent and transitory components of the changes in the $\beta$'s. If $\gamma$ is significantly different from zero, then the $\beta$'s are subject to permanent changes over time. Since we are interested in the permanent component of the parameter process and, in particular, specific realizations of the process, we normalize the equation around a specific time period. If we let $\tau$ represent such a period and $X_t'$ represent the row vector of independent variables $(1, P_{t-1}, S_{t-1})$, we can rewrite the producers' response equation as

$$S_t = X_t'\beta_\tau^p + \pi_t. \tag{13}$$

The error vector $\pi_t$ is distributed normally with mean zero and covariance matrix

$$\text{cov}(\pi) = \sigma^2[(1-\gamma)R + \gamma Q] \equiv \sigma^2\Omega(\gamma). \tag{14}$$

The matrix $R$ is a diagonal matrix which depends on $\sum_u$ and $X$, while $Q$ is a matrix which depends on $X$, $\sum_w$ and the period on which the parameter process is normalized.

The object of the estimation procedure is to obtain a consistent estimate of $\gamma$ which will yield the asymptotically efficient estimates of the $\beta$'s. The formal details of the estimation technique and the asymptotic properties of the estimates are developed fully in Cooley (1971), Cooley and Wall (1975), and Cooley and Prescott (1976). In this study we calculated the Bayesian estimates of the parameters, assuming priors which are sufficiently diffuse so that the sample information dominates. The parameter $\gamma$ is only estimated once for each state, and estimates of $\beta_\tau^p$ conditional on this $\gamma$ were computed at five-year intervals over the entire sample period.

## IV. Results

Estimation of (9) and examination of the estimates at five-year intervals reveal the existence of substantial parameter variation. Realizations of the response function at each point of normalization (1874, 1879, ..., 1914) are very similar in general appearance to the functions obtained previously by Fisher and Temin (1970) and DeCanio (1973). The $b_i$ estimates[6] are generally large relative to their standard errors ($t$-ratios larger than 2 in absolute value) and of the correct sign in most cases. The direction of drift of the varying constant term is, with only a few exceptions, the same as the sign on the pure trend included in the Fisher and Temin, and DeCanio specifications. One interesting difference between the fixed-parameter and varying-parameter estimates is that the varying-parameter estimates of $\beta_3$ are usually smaller than the $\beta_3$ coefficients in the fixed-parameter specifications incorporating both a trend and an autocorrelated disturbance, which in turn are smaller than estimates of $\beta_3$ derived from a naive model without a trend and estimated by ordinary least squares. Misspecification of the explanatory variables can lead to biased estimates of the coefficient of a lagged dependent variable if the omitted factors exert a persistent influence. The varying-parameter technique is effectively a particular kind of unobservable component procedure and a more general alternative to the first-order autocorrelated disturbance process (Cooley and Prescott, 1973c, p. 468), and therefore should be less susceptible to bias in the coefficient of the lagged share.[7]

[6] Here and throughout, Roman letters will be used to designate the estimated values of parameters designated by Greek letters.

[7] Tables of these comparisons, as well as the realizations of the response functions at the endpoints of the sample period, are available on request from the authors and are also given in Cooley and DeCanio (1974).

Our present interest is centered, however, not on the gross outlines of the response functions (which have already been determined by the fixed-parameter estimates), but on the fine structure of the parameter *variation* over time. The first observation to be made regarding this variation is that it is substantial and significant in most states. Table 1 contains estimates of the measure of parameter variation $\gamma$ and values of the ratio of the estimate $g$ to its standard error $s_g$ for each state. The small sample distribution of $g$ is not known, but $g/s_g$ is asymptotically normal (Cooley and Prescott, 1976). On the basis of the asymptotic distribution, the null hypothesis $\gamma = 0$ can be rejected at the 5% level (one-tailed test) for five cotton states and nine wheat states; at the 10% level (one-tailed test) it can be rejected for nine cotton states and fourteen wheat states.

TABLE 1. — ESTIMATES OF $\gamma$ FOR STATE RESPONSE FUNCTIONS

| Cotton States | $g$ [$g/s_g$] | Wheat States | $g$ [$g/s_g$] |
|---|---|---|---|
| North Carolina | .513 [2.389] | Iowa | .537 [2.418] |
| South Carolina | .258 [1.278] | California | .321 [1.457] |
| Georgia | .346 [1.430] | Kansas | .339 [1.256] |
| Florida | .327 [1.792] | Nebraska | .569 [2.493] |
| Tennessee | .380 [1.683] | Minnesota | .656 [3.227] |
| Alabama | .310 [1.463] | Illinois | .381 [1.945] |
| Mississippi | .525 [2.373] | Maryland | .355 [1.553] |
| Arkansas | .283 [1.324] | Michigan | .267 [1.404] |
| Louisiana | .683 [3.330] | Missouri | .500 [1.954] |
| Texas | .322 [1.425] | Wisconsin | .184 [1.185] |
| | | Indiana | .219 [1.568] |
| | | Virginia | .450 [1.892] |
| | | Pennsylvania | .398 [1.438] |
| | | New York | .415 [1.533] |
| | | Ohio | .213 [1.067] |
| | | North Dakota | .399 [1.669] |
| | | South Dakota | .418 [1.674] |

Notes: Sample period for cotton states is 1870–1914; for wheat states 1367–1914, except California 1869–1914, Minnesota 1868–1914, North and South Dakota 1883–1914.

The problem remains of relating this parameter variation to farmers' behavior. Consider the weight $\lambda$ in the price expectation equation (4). Nerlove has argued that the "coefficient of expectation" ($1 - \lambda$ in our notation) should vary inversely with the variance of the actual price series, provided increases in observed variance are associated with a greater than proportional increase in the variance of the transitory component of the price (1967, p. 145). This follows because if changes in the price are due largely to "noise" rather than to permanent changes, then farmers' optimal forecast of future prices will give nearly equal weights to all past observations ($\lambda$ closer to unity) so that transitory components tend to cancel each other out. Nerlove found such an inverse relationship between coefficients of expectation and the variability of the prices of various crops in his original study (1958, p. 221). In our relative price series disaggregated to the state level, however, it is not at all obvious whether the transitory component dominates, particularly for the wheat states. Table 2 contains estimates of the $\gamma$ measure of the ratio of the variance of the permanent component to the variance of the transitory component for each state's relative price series. These estimates were generated by performing a varying-parameter regression of the price on a constant, specifying $\sum_u = \sum_w = [1]$. This procedure yields estimates of $\gamma = \rho/(1 + \rho)$, where $\rho =$ the variance ratio $\sigma_\epsilon^2/\sigma_\eta^2$. The $\gamma$ measures estimated for each state's price series should not be confused with the $g$'s of table 1 measuring total parameter variation in the state response functions. To avoid confusion, we shall denote the "$\gamma$" for the price series by the upper case $\Gamma$. The $G$ of table 2 show that while the variance of the transitory component dominates the variance of the relative cotton price series (low $\Gamma$), the same cannot be said

TABLE 2. — ESTIMATES OF THE RELATIVE VARIANCE OF PERMANENT AND TRANSITORY COMPONENTS OF THE OBSERVED RELATIVE PRICE

| Cotton States | $G$ [$G/S_G$] | Wheat States | $G$ [$G/S_G$] |
|---|---|---|---|
| North Carolina | .152 [1.613] | Iowa | .200 [ .908] |
| South Carolina | .040 [ .700] | California | .079 [ .698] |
| Georgia | .098 [1.212] | Kansas | .131 [ .925] |
| Florida | .129 [1.279] | Nebraska | .214 [1.084] |
| Tennessee | .153 [1.609] | Minnesota | .254 [1.114] |
| Alabama | .081 [1.069] | Illinois | .412 [1.359] |
| Mississippi | .050 [ .802] | Maryland | .597 [2.465] |
| Arkansas | .064 [ .985] | Michigan | .421 [1.448] |
| Louisiana | .084 [1.080] | Missouri | .147 [ .908] |
| Texas | .048 [ .751] | Wisconsin | .589 [1.972] |
| | | Indiana | .584 [2.093] |
| | | Virginia | .644 [2.896] |
| | | Pennsylvania | .792 [4.786] |
| | | New York | .764 [4.361] |
| | | Ohio | .573 [2.079] |
| | | North Dakota | .168 [ .820] |
| | | South Dakota | .203 [ .869] |
| Unweighted Average | .090 | Unweighted Average | .398 |

Notes: (1) The equation $P_t = \beta_0$ was estimated by the varying parameter technique with $\Sigma_u = \Sigma_w = [1]$, yielding the $G$ estimates. (2) Sample period for each state is the same as in table 1.

with certainty for the relative wheat price series. Although most of the statewide cotton price series exhibit $G/S_G$ ratios suggestive of some permanent variation, the average $G$ for the wheat states is more than four times the average $G$ for the cotton states, and the $t$-ratios are generally larger as well.

To test the relationship between the expectations parameter $\lambda$ and price variability, changes in $\lambda$ over time were correlated with changes in the variance of prices within each state. When this is done, increases in price variability are found to be associated with increases in $\lambda$ (hence decreases in $1 - \lambda$) for the cotton states but with *decreases* in $\lambda$ for the wheat states. This pattern is consistent with Nerlove's hypothesis, for only in the cotton states does the variance of the transitory component clearly dominate the total price variance. The test was performed by defining $v_t$ as the sample standard deviation of the price variable computed over the five-year interval ending in $t$. Correlating first differences (to eliminate potentially spurious correlations due to common trends) of the $v_t$ with first differences of the $b_3$ calculated at five-year intervals yields 8 (of 10 possible) positive correlations for the cotton states, but 14 (of 17 possible) negative correlations for the wheat states.

This test is inconclusive because there is no unique way of measuring changes in the variability of the price series over time, and more importantly, because it does not distinguish between variance arising from the permanent and transitory components of the price. On the other hand, decomposition of the price series into permanent and transitory components allows a *direct* test of the rational expectations hypothesis. Recalling equation (5), the price expectations parameter $\lambda$ is a function of the variance ratio $\rho = \sigma_\epsilon^2/\sigma_\eta^2$, with $d\lambda/d\rho < 0$. Therefore, changes in $\lambda$ should be inversely related to changes in $\rho$, assuming some finite memory span of the decision-makers. In fact, just such a negative relationship holds for *both* the cotton and wheat states. Let $G_t$ be the estimate of $\Gamma$ for a state's relative price series computed over the five-year interval ending in $t$, and define $r_t$ as the "temporary relative variance" of the state's relative price series for the same five-year period. ($\Gamma = \rho/(1 + \rho)$, so $r_t = G_t/(1 - G_t)$.) Correlations of the temporary relative variance with successive realizations of $\beta_3$ at five-year intervals are given in table 3. When first differences are correlated, there is only one exception to the predicted sign pattern out of the ten cotton states, and five exceptions out of the seventeen wheat states. The undifferenced correlations show two exceptions in the cotton states and five in the wheat states.

TABLE 3. — SIMPLE CORRELATIONS BETWEEN PARAMETER ESTIMATES AND THE TEMPORARY RELATIVE VARIANCE OF THE PRICE

| Cotton States | Corr $(\Delta b_3, \Delta r_t)$ | Corr $(b_3, r_t)$ | Wheat States | Corr $(\Delta b_3, \Delta r_t)$ | Corr $(b_3, r_t)$ |
|---|---|---|---|---|---|
| North Carolina | −.355 | .079 | Iowa | −.212 | −.317 |
| South Carolina | −.118 | −.403 | California | −.691 | −.096 |
| Georgia | −.120 | −.298 | Kansas | .138 | −.195 |
| Florida | −.257 | −.348 | Nebraska | −.202 | −.447 |
| Tennessee | −.079 | −.095 | Minnesota | −.363 | .025 |
| Alabama | −.419 | −.425 | Illinois | −.046 | −.431 |
| Mississippi | −.389 | −.376 | Maryland | −.051 | .441 |
| Arkansas | −.129 | −.265 | Michigan | −.216 | −.079 |
| Louisiana | .880 | .843 | Missouri | .635 | .460 |
| Texas | −.607 | −.433 | Wisconsin | .158 | −.168 |
| | | | Indiana | −.061 | −.340 |
| | | | Virginia | −.027 | −.235 |
| | | | Pennsylvania | .354 | .027 |
| | | | New York | .106 | −.167 |
| | | | Ohio | −.175 | −.163 |
| | | | North Dakota | −.511 | −.037 |
| | | | South Dakota | −.007 | .433 |

Notes: (1) $\Delta$ is the first-differencing operator; $r_t$ is defined in the text. (2) Total sample period for each state is the same as in table 1; $b_3$ estimates computed at five-year intervals beginning with 1874 and ending with 1914; $r_t$ computed for successive five-year periods ending in $t = 1874, 1879, \ldots, 1914$, except North and South Dakota, which begin with 1889. (3) Number of first differences correlated for each state = 8 (5 for North and South Dakota); number of undifferenced estimates correlated for each state = 9 (6 for North and South Dakota).

It should be pointed out that the negative association between $b_3$ and $r$ also holds weakly *across* states. The simple correlation between the full-period price series' $r$'s (computed from the $G$'s of table 2) and the 1914 $b_3$'s is −0.219 for the cotton states and −0.081 for the wheat states. For 10 and 17 observations, respectively, these across-state correlations are not significant even at the 25% level, but it is not surprising that they are weaker than the typical within-state correlations over time. From equation (8), $\beta_3$ may vary across states with $(\theta_1 - \mu_1)$ as well as with $\lambda$. The parameters $\theta_1$ and $\mu_1$ are likely to depend on the overall crop mix, geography, and other factors varying more across states than within states. Our test of the rational expectations hypothesis over time is equivalent to controlling for the structural differences existing across states.

On the other hand, table 3 provides strong support for the rational expectations hypothesis. If all 27 states are considered as a single sample of independent observations of the correlation, the two-tailed binomial probability that the sign pattern of correlations would be as unbalanced as it is (or more unbalanced) is either 0.006 or 0.019, depending on whether the signs of the differenced or undifferenced correlations are counted.[8] Aside from this purely probabilistic argument, the results of table 3 eliminate the need for ad hoc explanations of the parameter variations in the response functions.[9] The aggregate behavior of both the cotton and wheat farmers is consistent with the same model of price prediction, a model based on a natural optimal decision rule. It is obvious, of course, that nineteenth-century farmers did not mathematically decompose the time series of relative crop prices into permanent and transitory components. They *are* likely to have been aware of whether relative price changes tended to persist or fluctuate randomly. A perspicacious farmer must surely have been aware of more of the information contained in the relevant crop price history than merely the value of the previous year's price.

In the case of cotton, the relative dominance of the transitory component of the price variance may have been due to the fact that permanent shifts in the world demand for cotton tended to be offset by corresponding shifts in supply brought about by American cotton farmers' price responsiveness, so that relative cotton *price* movements would not be permanent. That southern agrarian spokesmen were aware of the South's monopoly position in the world cotton market[10] is argued convincingly by Wright and Kunreuther (1975), and a realization of the adverse macro effect of each individual cotton farmer's micro rationality could account for Southerners' preoccupation with the issue of cotton "overproduction." In contrast, U.S. wheat production amounted to only around one quarter of total world output (United States Department of Agriculture, 1937, p. 18; Malenbaum, 1953, pp. 238-239) so that price responsiveness on the part of U.S. farmers in adjusting to permanent shifts in demand would not necessarily be price offsetting if producers in other parts of the world were less responsive. The "boom or bust" pattern of wheat prices has been commented on before, but the conclusion has usually been drawn that wheat farmers were confused by the course of these price movements. But while the precise *causes* of permanent or transitory wheat and cotton relative price fluctuations may not have been known to the farmers (and is still an interesting question for further research), the *history* of the price fluctuations was known to them, and our results indicate that a substantial portion of the farmers acted on this knowledge. Enough cotton and wheat farmers discounted recent price information (increased $\lambda$) when the variance of the transitory component of the pertinent relative price was increasing in comparison to the variance of the permanent component to produce measurable variations in the parameters of the state response functions.

## V. Conclusions

The results of the previous section support the rational expectations hypothesis in explaining farmers' response to price. Farmers were neither unresponsive to price changes nor insensitive to the history of fluctuations in the prices of their agricultural products. Enough farmers behaved optimally in the Muth sense to enable their reactions to be detected at the state-wide level of aggregation. Of course, the rational behavior of a substantial number of farmers

[8] Two-tailed binomial tests of the null hypothesis of equiprobability of positive and negative correlations yield low probability-values for the sign counts of the individual columns of table 3 as well.

[9] If variations in $\lambda$ were the only source of parameter variation, equation (6) suggests that the correlation between changes in $b_2$ and changes in $b_3$ would be $-1$. If there is independent variation in the long-run elasticity $\alpha_2$, the correlation between $b_2$ and $b_3$ will not be perfect. In actuality, the correlations between $b_2$ and $r_t$ are generally opposite in sign to the correlations between $b_3$ and $r_t$, although the association between the temporary relative variance and $b_2$ is somewhat weaker than its association with $b_3$ in both the cotton and the wheat states. Thus, variations in $b_2$ are also consistent with the rational expectations hypothesis.

[10] U.S. cotton comprised approximately two-thirds of the total world output of cotton during our sample period (United States Congress, Senate, 1895, vol. I, pp. 501-506; United States Department of Agriculture, 1937, p. 92).

does not preclude bewilderment or sub-optimal reactions on the part of many other farmers. But in a larger sense, even the farmers who were fully aware of their situation may not have been immune from economic distress.

For the wheat producers, the existence of permanent changes in the relative price of their main crop presented them with unique problems of response. Permanent shifts of the relative wheat price opened the possibility of gain by quick action in the event of a price change. Since price increases could be permanent, a "bonanza" approach to expansion of wheat acreage might pay large dividends. On the other hand, not *all* relative wheat price fluctuations were permanent, so farmers who rapidly revised their price expectations in response to *some* of the fluctuations must have been disappointed. Even if rapid expansion were temporarily successful, a period of greater-than-average prices could be followed by a period of less-than-average prices with distressing suddenness. Bad years associated with world weather or business cycles might induce acreage contraction without any subsequent price increase following the reduction in American supply. An awareness of these possibilities and attempts to adapt to them would not necessarily have guaranteed even the most intelligent farmers the security they desired. Cotton farmers faced a somewhat different situation since they provided a classic example of competitors unable to collude to jointly maximize profits. Awareness of their position and of its consequences for world cotton prices could improve decision-making at the micro level, but could not overcome the fundamental "problem of overproduction."

Our results suggest that at least a portion of the farmers displayed a remarkable degree of sophistication in their evaluation of the historical market data available to them. The perception of and adaptation to risks or special market opportunities does not guarantee satisfaction, however, especially when practicable alternatives are severely limited. Furthermore, if the pattern of parameter variation actually found was in part due to the survival of farms whose operators' expectations conformed to economic reality, the failure of the less-than-rational farming enterprises would not have been any easier for their owners to accept. But whether many farmers reacted optimally to the market from the onset of agricultural commercialization or microeconomic adjustment was accomplished through the failure of those who were slow to optimize, a significant number of farmers throughout the late nineteenth and early twentieth centuries responded rationally to the market information embodied in relative crop price movements.

## REFERENCES


Bowman, John D., "An Economic Analysis of Midwestern Farm Land Values and Farm Land Income, 1860 to 1900," *Yale Economic Essays* 5 (Fall 1965), 317-352.

Bowman, John D., and Richard H. Keehn, "Agricultural Terms of Trade in Four Midwestern States, 1870-1900," *Journal of Economic History* 34 (Sept. 1974), 592-609.

Cooley, Thomas F., *Estimation in the Presence of Sequential Parameter Variation*, unpublished Ph.D. dissertation, University of Pennsylvania, 1971.

Cooley, Thomas F., and Stephen J. DeCanio, "Varying-Parameter Supply Functions and the Sources of Economic Distress in American Agriculture, 1866–1914," National Bureau of Economic Research Working Paper No. 57, Cambridge, Mass. (Sept. 1974).

Cooley, Thomas F., and Edward Prescott, "Tests of an Adaptive Regression Model," this REVIEW 55 (May 1973a), 248-256.

———, "An Adaptive Regression Model," *International Economic Review* 14 (June 1973b), 364-371.

———, "Varying Parameter Regression: A Theory and Some Applications," *Annals of Economic and Social Measurement* 2 (Oct. 1973c), 463-473.

———, "Estimation in the Presence of Stochastic Parameter Variation," *Econometrica* 44 (Jan. 1976), 167-184.

Cooley, Thomas F., and Kent D. Wall, "On the Identification of Time Varying Structures," National Bureau of Economic Research Working Paper No. 85, Cambridge, Mass. (May 1975).

David, Paul A., "Comment" in Michael D. Intriligator (ed.), *Frontiers of Quantitative Economics* (Amsterdam: North-Holland Publishing Co., 1971), 459-467.

Davis, Lance E., Jonathan R. T. Hughes, and Duncan M. McDougall, *American Economic History: The Development of a National Economy*, third edition (Homewood, Illinois: Richard D. Irwin, Inc., 1969).

DeCanio, Stephen J., "Cotton 'Overproduction' in Late Nineteenth-Century Southern Agriculture," *Journal of Economic History* 33 (Sept. 1973), 608-633.

———, "Productivity and Income Distribution in the Post-Bellum South," *Journal of Economic History* 34 (June 1974a), 422-446.

———, *Agriculture in the Postbellum South: The Eco-*

*nomics of Production and Supply* (Cambridge, Mass.: The M.I.T. Press, 1974b).

Fisher, Franklin M., and Peter Temin, "Regional Specialization and the Supply of Wheat in the United States, 1867-1914," this REVIEW 52 (May 1970), 134-149.

———, "Regional Specialization and the Supply of Wheat in the United States, 1867-1914: A Reply," this REVIEW 53 (Feb. 1971), 102-103.

———, "Regional Specialization and the Supply of Wheat in the United States, 1867-1914: A Reply," this REVIEW 56 (Feb. 1974), 114-115.

Gray, Ralph, and John M. Peterson, *Economic Development of the United States* (Homewood, Illinois: Richard D. Irwin, Inc., 1974).

Griliches, Zvi, "Distributed Lags: A Survey," *Econometrica* 35 (Jan. 1967), 16-49.

Hayami, Yujiro, and Vernon W. Ruttan, *Agricultural Development: An International Perspective* (Baltimore: The Johns Hopkins Press, 1971).

Higgs, Robert, "Regional Specialization and the Supply of Wheat in the United States, 1867-1914: A Comment," this REVIEW 53 (Feb. 1971), 101-102.

Klepper, Robert, "The Economic Basis for Agrarian Protest Movements in the United States, 1870-1900," *Journal of Economic History* 34 (Mar. 1974), 283-285.

Malenbaum, Wilfred, *The World Wheat Economy* (Cambridge, Mass.: Harvard University Press, 1953).

Mayhew, Anne, "A Reappraisal of the Causes of Farm Protest in the United States, 1870-1900," *Journal of Economic History* 32 (June 1972), 464-475.

Muth, John F., "Optimal Properties of Exponentially Weighted Forecasts," *Journal of the American Statistical Association* 55 (June 1960), 299-306.

Nerlove, Marc, *The Dynamics of Supply: Estimation of Farmers' Response to Price* (Baltimore: The Johns Hopkins Press, 1958).

———, "Distributed Lags and Unobserved Components in Economic Time Series," in *Ten Economic Studies in the Tradition of Irving Fisher* (New York: John Wiley and Sons, 1967).

North, Douglass C., *Growth and Welfare in the American Past*, second edition (Englewood Cliffs, N.J.: Prentice-Hall, Inc., 1974).

Page, Walter P., "Wheat Culture and Productivity Trends in Wheat Production in the United States, 1867-1914: A Comment." this REVIEW 56 (Feb. 1974), 110-114.

Rosenberg, Barr, "The Analysis of a Cross Section of Time Series by Stochastically Convergent Parameter Regression," *Annals of Economic and Social Measurement* 2 (Oct. 1973), 399-428.

Sahi, R. K., and W. J. Craddock, "Estimation of Flexibility Coefficients for Recursive Programming Models—Alternative Approaches," *American Journal of Agricultural Economics* 56 (May 1974), 344-350.

United States Bureau of the Census, *Historical Statistics of the United States, Colonial Times to 1957* (Washington, D.C.: U.S. Government Printing Office, 1960).

United States Congress, Senate, *Report of the Committee on Agriculture and Forestry on Condition of Cotton Growers in the United States, the Present Prices of Cotton, and the Remedy; and on Cotton Consumption and Production*, 53d Congress, 3d Session, Report 986, 2 vols. (Washington, D.C.: U.S. Government Printing Office, 1895).

United States Department of Agriculture, *Agricultural Statistics 1937* (Washington, D.C.: U.S. Government Printing Office, 1937).

Wright, Gavin, and Howard Kunreuther, "Cotton, Corn and Risk in the Nineteenth Century," *Journal of Economic History* 35 (Sept. 1975), 526-551.

Zellner, Arnold, *An Introduction to Bayesian Inference in Econometrics* (New York: John Wiley and Sons, Inc., 1971).