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THE STRUCTURE OF U.S. AGRICULTURAL TECHNOLOGY, 1910-1978

by

John M. Antle

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The Structure of U.S. Agricultural Technology, 1910-1978

American agriculture of the twentieth century is an important example of successful agricultural development. It displayed a remarkable transformation, a high rate of productivity growth, and dramatic changes in factor proportions and production scale (Cochrane). Understanding this transformation is important for both theorists and policy analysts.

Hayami and Ruttan hypothesized that relative resource prices in an economy determine the direction of innovation followed by both private entrepreneurs and public institutions. Hayami and Ruttan argue convincingly that scarce labor and abundant land in the United States plus the scientific advances of the twentieth century led to innovations which provided substitutes for scarce labor and complements for abundant land, notably farm machinery, chemical fertilizers, and high-yielding plant varieties. While Hayami and Ruttan (Ch. 6) provided evidence that U.S. factor proportions were highly correlated with relative prices over the past one hundred years, their analysis does not explain how the structure of technology changed to make new factor proportions possible.

Binswanger suggests that relative price trends in the United States induced technology to be biased towards chemical and mechanical processes and against labor. His study of aggregate agricultural technology, based on a homothetic translog cost function, measured a bias in technical change consistent with the induced innovation theory. It showed technical change to be biased towards mechancial and chemical technology and against land and labor in the 1912-1968 period. Several critical assumptions underlying his analysis remain untested. Thus, the validity of his findings may be questioned.

The assumption of a homothetic technology made by Binswanger should be subjected to empirical test. Berndt and Khaled noted that, with homotheticity, cost shares are invariant to production scale. With nonhomothetic technology, cost shares change with scale. If the technology was nonhomothetic and a substantial scale change occurred during the 1912-1968 period, as the literature suggests, then Binswanger's estimates could be incorrect. One issue which needs to be resolved, therefore, is whether U.S. agricultural technology is homothetic or not. If not, to what degree does the homotheticity assumption affect the measured bias in technical change? This question is especially relevant since a recent study by Lopez showed that Canadian agricultural technology is nonhomothetic and that the measured technology bias is affected by homotheticity restrictions.

Another of Binswanger's assumptions is that the parameters of the aggregate cost function for the 1949-1964 period are valid for the entire 1912-1968 period. Considering the massive structural and technical changes that occurred in this period, parameter stability is an important assumption which needs to be tested.

Kislev and Peterson offer another explanation of how relative prices induced technical change in U.S. agriculture. They argue that all agricultural productivity growth cannot be attributed to technical change on the farm. They differentiate innovations within agriculture (primarily biological agricultural research and human capital accumulation) from innovations in manufacturing which improved farm machinery. This distinction between "internal" and "external" technical change means that input data, especially for machinery, must be adjusted for quality change to reflect external technical change. With quality-adjusted data, the resulting estimates of biased technical change can be attributed to internal innovations. Kislev and Peterson criticize Binswanger's use of USDA time series which are not adequately adjusted for quality change.

This paper utilizes 1910-1978 data and an aggregate translog profit function to measure the structure of U.S. aggregate agricultural technology and to test the hypotheses discussed above. Estimates of aggregate agricultural technology for 1910-1946 and 1947-1978 indicate that, within these periods, the technologies are different and nonhomothetic. Both biased technical change and scale changes explain the observed trends in aggregate factor use. These technological differences are consistent with the theory of induced innovation and can be rationalized by different relative prices during the two periods. The inappropriate assumption of homotheticity biases estimates of the technology structure.

Nonhomotheticity, Biased Technical Change, and the Translog Profit Function

In this section, the empirical model is defined and used to measure biased technical change and nonhomotheticity and to construct tests of hypotheses about the structure of technology. The theoretical framework for the normalized profit function is provided by Lau. The restricted normalized profit function is

(1) $\pi = G[p, Z]$

where G is a convex function of p, $p = (p_1,\ldots,p_n)$ is a vector of normalized input prices ($p_i=w_i/P$, where w_i is the ith input price and P is aggregate output price) and $Z = (Z_1,\ldots,Z_m)$ is a vector of exogenous variables. The Z_i are a time variable, interpreted as an index of technical change, and time dummy variables. I use the translog profit function

$$\text{(2)}\quad \ln G[p, Z] = \alpha_o + \sum_i^n \alpha_i \ln p_i + 1/2 \sum_i^n \sum_j \alpha_{ij} \ln p_i \ln p_j + \sum_i^m \beta_i \ln Z_i$$

$$+ \sum_i^n \sum_j^m \beta_{ij} \ln p_i \ln Z_j + \sum_i^m \gamma_i (\ln Z_i)^2.$$

Applying Hotelling's Lemma to (2), one obtains input demand and output supply functions. For parameter estimation I utilize the following equations derived from input demand equations:

$$\text{(3)}\quad \pi_i = \frac{\partial \ln G}{\partial \ln p_i} = \alpha_i + \sum_j^n \alpha_{ij} \ln p_j + \sum_j^m \beta_{ij} \ln Z_j,\ i=1,\ldots,n,$$

where $\pi_i \equiv -x_i p_i/\pi$ and x_i is the profit maximizing quantity of the ith input. The output supply function is

$$\text{(4)}\quad Q = G[p, Z] + \sum_{i=1}^n p_i x_i = G[p,Z]\ (1 - \sum_{i=1}^n \partial \ln G/\partial \ln p_i)$$

where the latter equality occurs because (3) implies $x_i = -\pi\ \pi_i/p_i$. Input demand elasticities can be calculated from (3) (Sidhu and Baanante).[1] For the own price, cross price, and output price they are

$$\text{(5)}\quad \eta_{ii} = \alpha_{ii}/\pi_i + \pi_i - 1$$

$$\text{(6)}\quad \eta_{ij} = \alpha_{ij}/\pi_i + \pi_j\ ,\ i \neq j$$

$$\text{(7)}\quad \eta_i = -\sum_{j=1}^n \eta_{ij}.$$

Output supply elasticities are derived from (4). The elasticity of supply with respect to output price P is

$$(8)\quad \nu_o = -\sum_i^n \pi_i + \sum_i^n \sum_j^n \alpha_{ij}/(1 - \sum_i^n \pi_i)$$

and the elasticity of supply with respect to the ith input price is

$$(9)\quad \nu_i = \pi_i - \sum_j^n \alpha_{ij}/(1 - \sum_i^n \pi_i),\ i = 1,\ldots,n.$$

Thus $\nu_o = -\sum_{i=1}^n \nu_i$. With the homotheticity restrictions discussed below, $\sum_j \alpha_{ij} = 0$ so that $\nu_o = -\sum_i^n \pi_i$ and $\nu_i = \pi_i$.

Testing for Homotheticity

The translog profit function (2) is nonhomothetic in the p_i. This implies that the corresponding production function is nonhomothetic in the inputs x_i (Lau). Testable homotheticity restrictions are obtained by noting that the profit function is homogeneous of degree $\sum_i^n \alpha_i$ if and only if

$$(10)\quad \begin{aligned} &\sum_j^n \alpha_{ij} = 0\ ,\ i=1,\ldots,n \\ &\sum_i^n \beta_{ij} = 0\ ,\ j=1,\ldots,m \end{aligned}$$

It is nonhomothetic otherwise. With (10) we have

$$(11)\ \ln G[\lambda p,Z] = \ln G[p,Z] + \ln \lambda \sum_i^n \alpha_i$$

Hence, the restrictions force the profit function to be homogeneous of degree $\sum_{i}^{n} \alpha_i$, and the dual production function is homogeneous of degree $\sum_{i}^{n} \alpha_i / (\sum_{i}^{n} \alpha_i - 1)$. Equation (10) allows homotheticity to be tested under the assumption of a translog technology. However, it is possible that if (10) is rejected by a statistical test, some other homothetic, nonhomogeneous function might not be rejected by the same data.[2]

Measuring Biased Technical Change

Several methods of measuring biased technical change have been used. The Hicksian measure based on marginal rates of technical substitution has the drawback that biases must be measured between every input pair (Belinfante). Binswanger has proposed that changes in factor cost shares, with constant factor prices, be used to measure biased technical change. Changes in the ith factor cost share, c_i, can be decomposed into price effects, output effects, and the effects of exogenous variables. Let t represent the effects of exogenous factors over time. Then

$$(12) \quad \frac{d\ln c_i}{dt} = \sum_{j}^{n} \frac{\partial \ln c_i}{\partial \ln p_j} \frac{\partial \ln p_j}{\partial t} + \frac{\partial \ln c_i}{\partial \ln Q} \frac{\partial \ln Q}{\partial t} + \frac{\partial \ln c_i}{\partial t}$$

Under homotheticity, $\partial \ln c_i / \partial \ln Q = 0$. Then biased technical change can be measured as a residual, as Binswanger did, by calculating the difference between actual cost share changes and estimated cost share changes due to factor price changes. This measure is valid if all residual changes, represented by $\partial \ln c_i / \partial t$, are caused by technology differences over time.

In this study, an alternative measure of biased technical change was developed. It involves production elasticities of the dual production function. This measure of biased technical change is equivalent to both the Hicksian measure and that based on cost shares. For production function $f(X,Z)$ let the ith production elasticity be

$$\varepsilon_i \equiv f_i x_i / f,$$

where f_i is the marginal product of x_i, and define

$$\varepsilon \equiv \sum_i^n \varepsilon_i.$$

Now use the production elasticity share $\varepsilon_i/\varepsilon$ to define biased technical change as follows. Technical change is biased towards (against) input i as

$$(13) \quad B_i \equiv \partial \ln (\varepsilon_i/\varepsilon)/\partial \ln t$$

is greater (less) than zero. Technical change is neutral with respect to input i when $B_i = 0$.

This measure of biased technical change is introduced for several reasons. First, in equilibrium $\varepsilon_i/\varepsilon = c_i$, where c_i is the ith factor cost share. Hence, B_i is equivalent to the cost share measure of biased technical change defined in (12). Using the duality result that $p_i = \partial f/\partial x_i$,

$$\varepsilon_i/\varepsilon = f_i x_i / f \quad \sum_i^n f_i x_i / f$$

$$= p_i x_i / \sum_i^n p_i x_i = c_i.$$

Second, the B_i are eqivalent to the the Hicksian measure of technical change. Technical change is Hicks-neutral for all factors if and only if $B_i = 0$ for all i. To prove this, note that technical change is Hicks-neutral with respect to two inputs x_i and x_j if $\partial(f_i/f_j)/\partial t = 0$, or if

$$(14)\quad \frac{\partial f_i}{\partial t}\frac{f_j}{f_i} = \frac{\partial f_j}{\partial t}$$

at given input and output levels. From (13) and the definitions of ε_i and ε we have

$$(15)\quad B_i = \frac{\partial f_i}{\partial t}\frac{t}{f_i} - \sum_j \frac{\partial f_j}{\partial t}\frac{x_j}{f}\frac{t}{\varepsilon}.$$

Substituting (14) into (15) it follows that Hicks-neutral technical change for all inputs is a sufficient condition for $B_i = 0$ for all i. In addition, $B_i = 0$ for all i is a sufficient condition for Hicks-neutral technical change. From (13) and (15), $B_i = 0$ implies

$$\frac{\partial f_i}{\partial t}\frac{t}{f_i} = \frac{\partial \ln \varepsilon}{\partial t} \text{ for all i,}$$

and therefore

$$\frac{\partial f_i}{\partial t}\frac{t}{f_i} = \frac{\partial f_j}{\partial t}\frac{t}{f_j}, \text{ for all i, j.}$$

This condition is identical to the definition of Hicks-neutral technical change in equation (14). These results show that using B_i to measure technical change provides a test for Hicks-neutral technical change versus biased technical change. If $B_i = 0$ for all i, we cannot reject the hypothesis of Hicks-neutral technical change. Otherwise, technical change is biased.

The third rationale for defining biased technical change with the B_i is that they can be expressed as linear functions of profit function parameters. Thus, the B_i and their standard errors can be computed and used to test the Hicks-neutral hypothesis. To express the B_i in terms of the profit function parameters, consider that

$$\frac{\partial \ln G(p,Z)}{\partial \ln p_i} = - \frac{\varepsilon_i}{1-\varepsilon},$$

as shown by Lau. Therefore,

$$\varepsilon/1-\varepsilon = - \sum_{i=1}^{n} \frac{\partial \ln G}{\partial \ln p_i}$$

and

$$\frac{\varepsilon_i}{\varepsilon} = - \frac{\partial \ln G}{\partial \ln p_i} \Bigg/ \sum_{i=1}^{n} - \frac{\partial \ln G}{\partial \ln p_i}.$$

Letting $Z_1 \equiv t$ we have

$$(16) \quad B_i = \frac{\partial \ln (\varepsilon_i/\varepsilon)}{\partial \ln t} = \frac{\partial \ln \left[- \dfrac{\partial \ln G}{\partial \ln p_i}\right]}{\partial \ln t} - \frac{\partial \ln \left[\displaystyle\sum_{i=1}^{n} - \dfrac{\partial \ln G}{\partial \ln p_i}\right]}{\partial \ln t}$$

$$= [\beta_{i1}/\pi_i] - [\sum_{i=1}^{n} \beta_{i1}/ \sum_{i=1}^{n} \pi_i].$$

Since $\sum_{i=1}^{n} \varepsilon_i/\varepsilon = 1$, a weighted average of the B_i is zero:

$$\sum_{i=1}^{n} \pi_i B_i / \sum_{i=1}^{n} \pi_i = 0.$$

When the homotheticity restrictions in (10) are imposed, $B_i = \beta_{i1}/\pi_i$.

This methodology is attractive because it provides a measure of biased technical change with a nonhomothetic technology. However, it should be noted that this parametric approach, using the time variable to represent the effects of technical change, may pick up other effects that occur over time. For example, there were probably different rates of technical change in various crops which led to a change in the composition of aggregate output and aggregate factor proportions. Consequently, neutral technical change in crops with different factor proportions could be measured as biased technical change in an aggregate model. The results of the empirical analysis must be interpreted subject to the qualification that there may be aggregation bias.[3]

Measuring Nonhomotheticity

If the technology is nonhomothetic, changes in factor shares stem from output changes, scale changes, input price changes, or biased technical change. A scale change is defined as a change in output due to a proportional change in all input prices. Therefore, the effects of scale change on production elasticities and factor cost shares can be used to measure the technology's nonhomotheticity. In equilibrium, this can be translated into the profit function by dividing input prices by λ. Thus, the relative effect of a scale change on cost shares is

$$(17)\quad N_i \equiv \frac{\partial \ln (\varepsilon_i/\varepsilon)}{\partial \ln \lambda} = \frac{\partial \ln\left[-\dfrac{\partial \ln G[p/\lambda, Z]}{\partial \ln p_i}\right]}{\partial \ln \lambda} - \frac{\partial \ln\left[-\sum\limits_{i}^{n} \dfrac{\partial \ln G[p/\lambda, Z]}{\partial \ln p_i}\right]}{\partial \ln \lambda}$$

$$= -\sum_{j}^{n} \alpha_{ij}/\pi_i + \sum_{i}^{n}\sum_{j}^{n} \alpha_{ij}/\sum_{i}^{n} \pi_i$$

From the homotheticity restrictions in (10), it follows that $N_i = 0$ for all i if the technology is homothetic. Also, a weighted average of the N_i equals zero, with weights $(\pi_i/\sum_i^n \pi_i)$. This shows that some N_i must be positive and some N_i must be negative unless all equal zero.

Data, Estimation, and Testing

The production data are based on U.S. Department of Agriculture (1980a) and Department of Commerce aggregate annual time series for the period 1910-1978, excluding 1919-1921 and 1929-1934 which are extreme outlying observations. The variables are prices and quantities of total production, farm labor (L), farm real estate (T), mechanical power and machinery (M), and agricultural chemicals (C). A trend variable is included to represent the effects of technical change and time dummy variables are included to measure intercept shifts.[4]

The input data must be measured in constant efficiency units to obtain correct parameter estimates. Since the machinery price and quantity indices do not adequately reflect quality change (U.S. Department of Agriculture 1980b, Kislev and Peterson), the adjustment method used by Hayami and Ruttan (Appendix C-2, pp. 336-337) to adjust for quality change was adopted for 1910-1960 and extended through 1978. Farm labor quality change also is accounted for by using a ten year moving average of the percentage of school age population enrolled in public elementary and secondary schools to calculate a quality adjustment factor. This factor is $q_t = E_t/E_{1910}$, where E_t is the averaged enrollment ratio. The labor quantity and price series are $L_t^* = L_t q_t$ and $p_{Lt}^* = p_{Lt}/q_t$. The chemicals and land data were judged to be adequately adjusted for quality change.

To translate the theoretical profit function into an econometric model, several assumptions must be made. The approach used here, as in other studies

(Berndt and Christensen, Berndt and Wood, Berndt and Khaled, Lopez), is to derive input demand relations from the first order conditions of profit maximization and estimate them as a system of equations. Implicit is the assumption that a systematic static aggregate relationship exists between input quantities and prices which reflects individual firm behavior. In addition, it must be assumed that input demand function parameters are being identified rather than input supply parameters. This is justified if shifts in input supply functions over time are large relative to demand shifts. A final assumption is that normalized aggregate resource prices are exogenous to agricultural input markets. Lopez observed that this last assumption is reasonable for all inputs except land. It is reasonable to argue that while these assumptions are likely to be violated to some degree, their overall validity can be judged by the results of various statistical tests of the model.

The most efficient estimators of the translog profit function are obtained by jointly estimating equations (2) and (3). However, using time dummy variables in the model yields insufficient degrees of freedom for direct estimation of the profit function (2). Estimation of the profit "share" equations (3) is possible. However, the β_i and γ_i parameters do not appear in these equations and cannot be estimated. Using the share equation estimates it is possible to measure all parameters needed both to test for homotheticity and to calculate input demand elasticities, output supply elasticities, factor biases, and scale effects. Without estimates of β_i and γ_i, it is not possible to measure the total effects of technical change on aggregate supply.

Equations (3) are assumed subject to random disturbances which may be contemporaneously correlated across equations and time. With these assumptions, tests can be performed to check for the validity of the translog specification. These include tests for the error specification, parameter stability over time, and appropriateness of the translog specification.

The unrestricted model was estimated using ordinary least squares to test for first-order positive autocorrelation of the residuals for the periods 1910-1946, 1947-1978, and 1910-1978. As table 1 shows, only the chemicals equations for the 1910-1978 and 1947-1978 samples have Durbin-Watson statistics in the inconclusive region. All other equations exhibit Durbin-Watson statistics greater than 2. An adjustment for autocorrelation of the residuals was deemed unnecessary. Without the time dummy variables, very small Durbin-Watson statistics were obtained, suggesting misspecification without time dummy variables.

Since the model spans the 1910-1978 period, tests were performed for parameter stability over time with the ordinary least squares estimates. The periods were 1910-1946, 1947-1978, and 1910-1978. This breakdown was justified by a careful study of the aggregate time series. The prewar data do not show a strong trend indicative of structural change. Yet the postwar period shows clear trends in both factor quantities and prices, table 2. Table 1 shows that the hypothesis of constant parameters is rejected for all equations at the 5 percent significance level.

Another test for model validity concerns the symmetry of the α_{ij} parameters. The test statistics in table 1 show symmetry is not rejected at the 1 percent level for the 1910-1946 and 1910-1978 samples and not rejected at 5 percent for the 1947-1978 sample. Hence, symmetry of the α_{ij} parameters across equations was maintained in the rest of the analysis. The estimated profit function was checked at the point of approximation for each sample with symmetry restrictions imposed and was found to be convex.[5] These findings suggest that the translog profit function is a satisfactory approximation to the true function for this analysis.

The Structure of U.S. Agricultural Technology

The parameters of the translog profit function are different for the sample periods 1910-1946 and 1947-1978. Therefore, I estimated and compared the structure of agricultural technology for these two periods. The system of equation (3) was estimated iteratively with the seemingly unrelated regression (SUR) estimator and the symmetry restrictions imposed for the α_{ij}. Magnus shows that these estimates are consistent and asymptotically normal, and equal to maximum likelihood estimates if the disturbances are normally distributed. These estimates and their standard errors are presented in table 3. Many parameters are precisely estimated, with only six parameters having t-statistics less than 1.9 in 1910-1946 and none in 1947-1978.

Test statistics for the homotheticity hypothesis, based on equation (10) and the parameter estimates of table 3, are 29.08 (7 and 86 degrees of freedom) for 1910-1946 and 45.70 (8 and 98 degrees of freedom) for 1947-1978. Since the corresponding 1 percent critical values of the F distribution are 2.85 and 2.69, the homotheticity restrictions are strongly rejected for both periods.

Estimates and test statistics are presented in table 4 for the B_i, equation (14), which measure internal biased technical change, and the N_i, equation (15), which measure nonhomotheticity as scale change effects on factor cost shares. These estimates are calculated at the sample means of the π_i. The 1910-1946 estimates show that the hypothesis of neutral (internal) technical change is rejected and that the bias is primarily towards machinery and against land. However, the biases are not large, and they contradict Binswanger's findings of a substantial bias towards chemicals during the prewar period. The N_i show that a scale increase would result in reduced chemical inputs, with small and statistically insignificant effects on labor, machinery, and land. The limited use of chemicals in the prewar period suggests that, overall, scale changes would not have significantly affected factor cost shares.

The 1947-1978 estimates, however, show a dramatic change in the structure of technology, both in the magnitude and direction of biased technical change and in nonhomotheticity. The bias effects for the 1947-1978 period are statistically significant and consistent with the induced innovation theory. The estimates show that internal technical change was biased mostly towards chemical inputs and against labor. This finding is consistent with the Kislev-Peterson analysis which suggests that internal productivity growth came primarily from biological innovations, whereas improvements in mechanical technology were caused by external innovations. The positive and significant value of N_i for land suggests that increased U.S. farm size in the postwar period was in part a result of the changing scale of U.S. agricultural technology.

Table 5 presents input demand and output supply elasticities based on the parameter estimates in table 3 and equations (5), (6), (7), (8), and (9) calculated at the sample means of the data. All own-price demand elasticities are negative as theory predicts, and most elasticities are absolutely less than one. Some substantial differences appear between the two periods. Most notable is the labor demand elasticity which is greater than one in the prewar sample but near zero in the postwar sample. The cross-elasticities of demand show most inputs exhibit complementarity in the 1910-1946 sample. In the 1947-1978 sample, labor substitutes for machinery and chemicals, and chemicals complement land. The estimated supply elasticities virtually all show a marked reduction in the postwar period as compared to the prewar period.

The absolute reduction in both demand and supply elasticities indicates a fundamental change in the technology structure. One explanation for the reduced price-responsiveness is that the postwar technology was more capital intensive. A related explanation is suggested by table 4. It shows that changes in factor use were more a function of biased technical change and scale change in the postwar period than in the prewar period. Thus, it is reasonable that input use in the postwar period was less price responsive than in the prewar period. One means of testing this explanation is to investigate the effects that homotheticity restrictions have on the price responsiveness of the technology, since this is equivalent to forcing the scale effects to be zero.

To determine how the imposition of homotheticity affects the estimates of the model, the restrictions (10) were imposed on the model along with symmetry restrictions and it was reestimated using iterated SUR. The values of the B_i, N_i, and η_{ii} derived from these parameter estimates are presented in table 6 with their standard errors. Since the homotheticity restrictions are strongly rejected by the data, the values in table 6 should be biased. They are indeed different from the corresponding values in tables 4 and 5. The B_i for 1947-1978 in table 6 underestimate the bias against labor and the bias toward chemicals. A rather large negative value was obtained for land as compared to the positive value in table 4. Binswanger obtained a very similar bias pattern under the assumption of homotheticity. The input demand elasticities of table 6 are uniformly greater in absolute value than those in table 5, and similar to those obtained by Binswanger.[6] The supply elasticities are also greater under the homotheticity restrictions. For 1910-1946 and 1947-1978, the supply elasticities under the restrictions are 2.485 and 1.440 as compared to 1.349 and 0.427 without the restrictions. Thus, it is evident that imposing homotheticity does make the technology appear more price-responsive. This is evidence that price inelasticity is a structural characteristic of the postwar technology.

Because the induced innovation hypothesis is important to the theory of economic development, we need to know why such different parameter estimates are obtained for the prewar and postwar periods. Do these findings contradict the induced innovation hypothesis? Several facts suggest that these seemingly contradictory findings can be explained.

Data showing how the prewar and postwar periods differed are presented in table 2. The exponential rates of change show that whereas machinery, chemicals, and land prices declined on average over the 1910-1978 period, the 1925-1940 and 1941-1978 price trends were in the opposite direction for machinery and chemicals and roughly zero for land. Considering the relative price trends in the prewar period, the estimates of the structure of the prewar technology do not clearly contradict the Hayami-Ruttan induced innovation hypothesis. Indeed, with the decline in the relative price of labor, a bias towards labor is consistent with induced innovation.

Another factor reconciling these results is that the induced innovation theory describes how long-run trends in factor prices affect the direction of technical change. There is likely to be a long lag between changes in relative price trends and the creation of new innovations. Therefore, when price trends change direction in relatively short time periods, as they did in the 1910-1946 period, it is reasonable to find a small average bias in the technology.

Table 2 also presents exponential rates of change in input indexes for the prewar and postwar periods. The prewar period showed relatively constant factor proportions indicative of little structural or technological change. The postwar period involved a dramatic technological revolution in mechanical and chemical innovation. The estimates of biased technical change and scale effects for the postwar period bear this out.

Conclusions

In this paper, a statistical methodology was developed for estimating input demand and output supply elasticities and testing hypotheses of homotheticity, parameter stability, and biased technical change. Measures of the effects on cost shares of both biased technical change and scale change were devised based on the translog profit function. The results of the empirical analysis can be summarized as follows:

(i) The hypotheses of homotheticity and parameter stability are rejected. Different, nonhomothetic aggregate technologies characterize the pre- and post-World War II periods.

(ii) The hypothesis of neutral technical change is rejected for both prewar and postwar periods. Moreover, the direction of bias is substantially different during these two periods. In contrast to Binswanger's estimates of biased technical change, the prewar technology is biased towards labor and mechanical technology and against land. These findings are not consistent with Binswanger's estimates or with the description of the United States as labor scarce and land abundant. However, the postwar estimates are more in agreement with Binswanger's findings. They show a strong and statistically significant bias against labor and towards mechanical and chemical technology. The estimates are consistent with the Kislev-Peterson theory of internal and external induced technical change.

(iii) The effects of scale change on factor cost shares caused by nonhomotheticity are weak in the prewar period but consistent with the bias in technical change during the postwar period. Hence, both scale change and biased technical change contributed to the observed trends in factor proportions. When homotheticity is imposed on the model, the estimated biases in technical change are both qualitatively and quantitatively different for the postwar period.

(iv) Estimates of input demand and output supply elasticities are biased upwards with the homotheticity restrictions. Thus, the misspecification of agricultural technology as homothetic when it is really nonhomothetic may produce biased estimates of technical change and other structural characteristics of technology.

Table 1

Test Statistics for the Unrestricted Model

Test	Equation L	M	C	T
Autocorrelation				
1910-1946	2.88	2.49	3.10	2.82
1947-1978	2.68	2.44	1.66	2.23
1910-1978	1.99	2.07	1.68	2.28
Parameter Change	11.43[a]	6.32[a]	2.97[b]	5.30[a]
	(6,42)	(6,42)	(6,42)	(6,42)

	Sample Period 1910-1946	1947-1978	1910-1978
Symmetry	2.26[b]	1.85	2.52[b]
	(6,80)	(6,92)	(6,192)

Note: [a]Significant at 1 percent level for the F-distribution.

[b]Significant at 5 percent level for the F-distribution.

Degrees of freedom for F-distribution in parenthesis.

Table 2

Exponential Rates of Change in Relative Prices and Factor Indexes

	1910-1978	1925-1940	1941-1978
Relative Prices			
Machinery/Labor	-0.017	0.030	-0.011
	(0.001)	(0.009)	(0.002)
Chemicals/Labor	-0.031	0.019	-0.031
	(0.001)	(0.007)	(0.002)
Land/Labor	-0.010	-0.002	0.012
	(0.001)	(0.005)	(0.002)
Factor Indexes			
Labor	-0.022	-0.009	-0.039
	(0.001)	(0.002)	(0.001)
Machinery	0.035	0.017	0.025
	(0.001)	(0.005)	(0.002)
Chemicals	0.056	0.023	0.066
	(0.001)	(0.012)	(0.001)
Land	-0.0001	0.0003	-0.001
	(0.0001)	(0.001)	(0.0003)

Note: Trend estimates based on regression of the logarithm of the variable on time.

Standard errors in parentheses.

Table 3

Symmetry Restricted, Iterated SUR Estimates

Of the Translog Profit Function

Parameter	1910-1946	1947-1978
α_L	-6.878 (0.824)	-11.812 (1.246)
α_M	-1.491 (0.182)	-5.891 (1.435)
α_C	-0.317 (0.047)	-0.401 (0.140)
α_T	-4.198 (0.536)	-5.304 (1.299)
α_{LL}	-1.309 (0.156)	-0.666 (0.099)
α_{MM}	-0.383 (0.049)	-0.610 (0.056)
α_{CC}	-0.041 (0.017)	-0.092 (0.022)
α_{TT}	-0.876 (0.097)	-0.462 (0.051)
α_{LM}	-0.282 (0.043)	-0.253 (0.047)
α_{LC}	-0.055 (0.011)	-0.102 (0.037)
α_{LT}	-0.381 (0.103)	-0.089 (0.044)
α_{MC}	-0.006 (0.017)	0.054 (0.027)
α_{MT}	0.108 (0.056)	-0.151 (0.047)
α_{CT}	0.004 (0.015)	0.069 (0.023)
β_{L1}	0.065 (0.111)	2.158 (0.260)
β_{M1}	0.064 (0.025)	0.831 (0.293)
β_{C1}	0.004 (0.006)	0.069 (0.023)
β_{T1}	0.212 (0.073)	0.895 (0.266)

Note: Standard errors in parentheses.

Parameters of time dummy variables not shown.

Table 4

Average Effects of Technical Change and Scale Change on Factor Cost Shares

	Equation			
	L	M	C	T
1910–1946				
B_i	0.039	0.273	0.014	-0.193
	(0.016)	(0.067)	(0.093)	(0.035)
N_i	0.032	-0.040	-0.480	0.048
	(0.031)	(0.142)	(0.217)	(0.067)
1947–1978				
B_i	-2.302	0.708	6.116	0.077
	(0.468)	(0.194)	(1.069)	(0.273)
N_i	-0.505	-0.035	1.257	0.266
	(0.181)	(0.072)	(0.389)	(0.103)

Note: B_i and N_i computed according to equations (13) and (14) at the sample means of the π_i.

Standard errors in parentheses.

Table 5

Estimated Input Demand and Output Supply Elasticities

Variable		Elasticity with respect to price of: L	M	C	T	Q
Labor:	1910-1946	-1.311	-0.135	-0.006	-0.459	1.911
		(0.121)	(0.033)	(0.008)	(0.079)	(0.199)
	1947-1978	-0.008	0.059	0.114	-0.190	0.025
		(0.216)	(0.102)	(0.081)	(0.096)	(0.214)
Machinery:	1910-1946	-0.504	-0.253	-0.031	-1.059	1.847
		(0.123)	(0.140)	(0.049)	(0.160)	(0.157)
	1947-1978	0.092	-0.252	-0.218	-0.077	0.455
		(0.095)	(0.114)	(0.055)	(0.096)	(0.169)
Chemicals:	1910-1946	-0.164	-0.225	-0.194	-0.667	1.250
		(0.229)	(0.354)	(0.354)	(0.312)	(0.287)
	1947-1978	0.485	0.008	-0.254	-1.023	0.784
		(0.343)	(0.250)	(0.204)	(0.213)	(0.522)
Land:	1910-1946	-0.802	-0.894	0.053	-0.582	2.225
		(0.137)	(0.075)	(0.020)	(0.129)	(0.224)
	1947-1978	-0.066	0.099	-0.288	-0.181	0.436
		(0.115)	(0.122)	(0.060)	(0.133)	(0.319)
Output:	1910-1946	-0.725	-0.187	-0.020	-0.419	1.349
		(0.076)	(0.016)	(0.004)	(0.049)	(0.136)
	1947-1978	-0.004	-0.222	-0.079	-0.125	0.427
		(0.040)	(0.034)	(0.023)	(0.045)	(0.119)

Note: Standard errors in parentheses.

Elasticities calculated according to equations (5)-(9) at the sample means of the π_i.

Table 6

Estimates of B_i, N_i, and η_{ii} With Homotheticity Restrictions

	Equation			
	L	M	C	T
1910-1946				
B_i	0.038	0.269	0.036	-0.194
	(0.015)	(0.061)	(0.087)	(0.033)
N_i	0	0	0	0
η_{ii}	-2.131	-0.610	-0.236	-1.082
	(0.022)	(0.140)	(0.250)	(0.112)
1947-1978				
B_i	-0.908	0.707	3.034	-0.670
	(0.214)	(0.090)	(0.503)	(0.142)
N_i	0	0	0	0
η_{ii}	-1.346	-0.923	-0.589	-0.637
	(0.168)	(0.067)	(0.204)	(0.094)

Note: Standard errors in parentheses.

B_i, N_i, and η_{ii} defined in equations (14), (15), and (5), computed at sample means.

Homotheticity restrictions defined in equation (10).

Footnotes

[1]The careful reader who compares these equations to Sidhu and Baanante should be aware of a typographical error in their equation (2), where a minus sign was omitted before S_i.

[2]Lopez used a generalized Leontief cost function which allows technology to be either constant returns to scale (CRS) or nonhomothetic. Lopez's test against homotheticity is actually a test against CRS; the rejection of CRS necessarily implies the technology is nonhomothetic only if a generalized Leontief cost function is correct.

[3]I am indebted to an anonymous referee for bringing this issue to my attention.

[4]For the 1910-1978 model, dummy variables are for 1930-1939, 1940-1949, 1950-1959, 1960-1972, 1973-1974, and 1975-1978; for the 1910-1946 model, dummy variables are for 1930-1939 and 1940-1946; for the 1947-1978 model dummy variables are for 1960-1972, 1972-1974, and 1975-1978. These periods were selected on the basis of major political and economic events that affected U.S. agriculture.

[5]At the approximation point of the translog profit function, where $\ln p_i = 0$, for all i, it can be shown that the profit function is convex if the symmetric matrix with kth diagonal element $\alpha_k^2 - \alpha_k + \alpha_{kk}$ and (k,ℓ) element $\alpha_{k\ell}+\alpha_k\alpha_\ell$ is positive definite (see Denny and Fuss).

[6]Binswanger's corresponding estimates are -0.911 for labor, -.089 for machinery, -0.945 for fertilizer, and -0.336 for land.

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