INDUCED INNOVATION IN UNITED STATES AGRICULTURE, 1880–1990: TIME SERIES TESTS AND AN ERROR CORRECTION MODEL

COLIN G. THIRTLE, DAVID E. SCHIMMELPFENNIG, AND ROBERT F. TOWNSEND

An error correction model (ECM) of induced innovation, based on the two-stage CES production function allows direct tests of the inducement hypothesis, which are applied to U.S. data for 1880–1990. The time series properties of the variables include a structural break in 1920, cointegration is established and an ECM constructed, which allows factor substitution to be separated from technological change. Causality tests show that the factor-price ratios and R&D are Granger-prior to the factor-saving biases of technological change. The inducement hypothesis is corroborated, and identified as one factor in the complex development of U.S. agriculture.

Key words: induced innovation, cointegration, error correction.

Hayami and Ruttan’s (1985, p. 4) own description of induced innovation is “technical change that facilitates the substitution of plentiful (hence cheap) factors of production for scarce (hence expensive) factor inputs.” The testable implications are that there should be negative correlations between factor-price ratios and factor ratios, with the causality running from the first to the second, and that this should be the result of technical change rather than factor substitution. The relationship may hold either in the changes or in the levels, as Olmstead and Rhode point out.

Several versions of this relationship between factor ratios and factor prices and numerous empirical studies are reported in Thirtle and Ruttan. In early cases the hypothesis is not clearly stated and the tests amount to no more than establishing a negative correlation between a measure of factor scarcity, such as relative prices and the factor ratios. But, negative relationships between input quantities and input prices may merely mean downward sloping input demand functions, which hardly corroborate the inducement hypothesis, as others have noted. For instance, Binswanger commented that the starting point for his own work on induced innovation was dissatisfaction with Hayami and Ruttan’s tests because they did not distinguish factor substitution from technical change. Hicks (1932) introduced both concepts, but he admitted in his Nobel lecture that the distinction between the two, “was left rather obscure” (reprinted in Hicks 1977, p. 2). This problem remained central in the criticisms of Blaug and of Salter, but Ahmad’s introduction of the innovation possibility curve rehabilitated the induced innovation hypothesis and is the basis of the Hayami and Ruttan formulation.

Whereas almost all the simple, early tests corroborated some aspect of induced innovation, as the tests have become more sophisticated, the empirical results have been far less supportive of the inducement hypothesis. This is especially true of recent tests applied to U.S. agriculture, which all reject the hypothesis.

This study applies time series techniques to 111 years of annual data for the U.S. agricultural sector. While it cannot claim to test the inducement hypothesis in a fully satisfactory manner, it goes further than any previous tests.
and does at least confront it with several hurdles. First, the series in the relationships must have time series properties allowing cointegration. Second, if there is a valid long-run relationship, cointegration must be established. Third, factor substitution must be separated from technical change and there must be some amount of factor ratio change, not accounted for by factor substitution, for induced innovation to explain. Fourth, the correlation between factor/price ratios and the factor ratios must be negative, and fifth, the causality must run from the prices to the ratios. Last, R&D must be one of the variables driving the causal process if it is to be identified as innovation rather than factor substitution.

Background

Hayami and Ruttan’s (1971) original three equation tests of the inducement hypothesis regressed the logarithms of the factor ratios (land/labor, fertilizer/land and machinery/labor) on the logarithms of the factor/price ratios. If the coefficient of the relevant price ratio is negative and significantly different from zero, the result is considered to corroborate the inducement hypothesis. In some cases, the results were poor; Ruttan et al. reported nine tests of the relationship between the land/labor price ratio and the land/labor ratio, in which five outcomes were inconsistent with the hypothesis.\(^1\) As in most early tests, the form of the production relationship is not stated, the distinction between factor substitution and induced innovation is not made, nor is the causality from price ratios to factor ratios tested.

The revised edition of Hayami and Ruttan (1985) corrects many of the defects, estimating the two-stage CES, with time-dependent factor augmentation coefficients, to produce estimates of the Allen elasticities of factor substitution and the biases of technical change. In a second stage, the total changes in factor shares are split into factor substitution and technical change and the share-based bias measures are plotted against relative prices, to test the inducement hypothesis. Similar two-stage CES approaches are used by Kawagoe, Otsuka, and Hayami who also apply the model to the United States and Japan, Thirtle (1985a), who considers U.S. wheat production, and Karagiannis and Furtan who use Canadian data. These studies collectively provide substantial corroboration for the inducement hypothesis, but none even touches on causality.

Binswanger’s (1974, 1978) pioneering applications of duality provided further support, by fitting a translog cost function, and using relative factor prices to explain the residual factor shares, net of factor substitution. The numerous duality-based cost and profit function approaches to technical change in agriculture that have followed, are conveniently summarized by Evenson and Pray. The “technology variables,” such as R&D and extension expenditures, that shift the flexible functional form over time, are included in the specification of the “meta-profit” function. Applications such as Huffman and Evenson (1989) study the bias effects of R&D and comment that the fertilizer-using and labor-saving biases found are “consistent with the induced innovation hypothesis,” though there are no actual tests. This is typical, as most of the dual studies address induced innovation only tangentially.

In contrast, Olmstead and Rhode concentrate on the inducement hypothesis, beginning with the observation that the relationship between factor ratios and factor/price ratios may be expressed in the levels or in terms of changes. They show that Hayami and Ruttan’s stylized facts do not match the data. Particularly, before 1910 the land price rose relative to wages and from 1910 to 1940, the price of land rose relative to the price of fertilizer. They stress the many endogenous factors affecting the land price, point out the very considerable regional differences that are hidden by the aggregation, and conclude that in their regressions, the inducement hypothesis is wrong about as often as it is right.

The hypothesis has also been increasingly challenged as time series tests have been developed. An early dynamic model of U.S. agriculture by Antle used aspects of time series modeling and concluded that the evidence on induced innovation was mixed, with two out of seven tests producing results that were contrary to the implications of the hypothesis. More recently, there have been two time series studies of induced innovation in U.S. agriculture. Machado fits the factor share equations from the dual cost function to Capalbo and Vo’s data for 1948–83. He estimates the biases of technical change and comes to the far

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\(^1\) Thirtle (1985b, 1985c) shows that the land/labor ratio tests fail both because the land price is endogenous and the land/labor ratio is determined by the relative rates of biological and mechanical technical change.
clearer conclusion that for the preferred model there is no cointegration between the factor biases and the factor prices, so the hypothesis is soundly rejected. In the second study, Tiffin and Dawson use Hayami and Ruttan’s (1985) data to find that the factor and price ratios are not of the same order of integration, so the hypothesis cannot hold either in the levels or in the changes. Instead, the factor/price ratios must explain changes in the factor ratios, but still, cointegration cannot be firmly established. Thus, all the recent tests for the United States are, at the very least, highly critical of the induced innovation hypothesis.

Other time series studies have reached the opposite conclusion. Khatri, Thirtle, and Townsend (1998) combine time series analysis with a third-order cost function fitted to British agriculture. They test induced innovation explicitly and find clear cointegrating relationships and that input prices are both negatively related and causally prior to input biases. Similarly, Thirtle, Townsend, and van Zyl applied cointegration and causality tests to an ECM based on the two-stage CES and found that all aspects of their tests supported the inducement hypothesis in the case of South African commercial agriculture. This model is developed in the next section.

A Model for Testing the Hypothesis

The inducement hypothesis implies a long-run relationship between the direction of technical change and a measure of factor scarcities, such as relative prices. The variables should not diverge too much in the long run. While there may be shorter-run deviations, there should be some equilibrating mechanism bringing them back together eventually (Granger 1986). Thus, cointegration techniques allow formal testing of the inducement hypothesis. Specifically, the time series properties of the series can be established, to ensure that there can be a non-spurious relationship between the variables (Yule). Then, if a cointegrating vector exists, an error correction model (ECM) can be constructed to determine the long-run relationships and the direction of causality can also be established.

The model developed by de Janvry, Sadoulet, and Fafchamps exploits the tractability of the two-stage CES by incorporating transaction costs and collective action as determinants of the factor-saving biases of technological change. Frisvold uses the model in a straightforward manner to test the induced innovation hypothesis. These two approaches are combined here, as in Thirtle, Townsend, and van Zyl.

\[
\begin{align*}
\text{Ln}\left[ \frac{F}{A} \right] &= \sigma_2 \text{Ln}\left[ \frac{1 - \alpha}{\alpha} \right] + (\sigma_2 - 1) \\
&\quad \times \text{Ln}E_f - \sigma_2 \text{Ln}\left[ \frac{P_f}{P_a} \right].
\end{align*}
\]

Rearranging the logarithms of the first-order conditions from the profit maximization problem and assuming equilibrium, so that the marginal products are equal to factor prices, gives these estimating equations for the first stage of the two-stage CES

\[
\begin{align*}
\text{Ln}\left[ \frac{M}{L} \right] &= \sigma_1 \text{Ln}\left[ \frac{1 - \beta}{\beta} \right] + (\sigma_1 - 1) \\
&\quad \times \text{Ln}E_m - \sigma_1 \text{Ln}\left[ \frac{P_m}{P_l} \right].
\end{align*}
\]

where \( M/L \) is the machinery-labor ratio, which is explained by a constant term, a term representing efficiency, \( E_m \), and the own price ratio \( P_m/P_l \). Similarly, the fertilizer/labor ratio, \( F/A \), is explained by a constant, an efficiency term, \( E_f \), and the own price ratio, \( P_f/P_a \). The \( \sigma \)'s represent the usual substitution parameters and the \( \beta \)'s are used to construct the factor share parameters of the CES function. The direct partial elasticity of substitution of labor for machinery is \( \sigma_1 \) and that for fertilizer and land is \( \sigma_2 \) (Kawagoe, Otsuka, and Hayami).

Thus, the two-stage CES approach gives a theoretical basis for direct tests of the inducement hypothesis, since the factor ratios are functions of the price ratios and the efficiency parameters. If the current factor-price ratios are significant in explaining factor substitution, the coefficients of these terms can be interpreted as direct partial elasticities of substitution. Frisvold suggests that if lagged price ratios explain the factor ratios, then the inducement hypothesis is corroborated.

The treatment of the efficiency parameters follows de Janvry, Sadoulet, and Fafchamps and Frisvold who assume that the efficiencies are functions of research activities,

\[
E_{f,m} = E_{f,m}[\theta, B, R, t]
\]

so where \( \theta \) is a vector of shares of past public sector research budgets, \( B \) allocated to land-saving technical change (\( E_f \)) and labor-saving technical change (\( E_m \)), \( R \) is a vector of past
private sector research expenditures, and \( t \) is a time trend representing exogenous change in scientific knowledge. The vector of share parameters, \( \mathbf{\theta} \), allocating the research budget between activities, itself depends on expected relative factor prices and assumptions regarding the government and private research budget allocation, so that

\[
0^* = \left( \frac{P^*_m}{P^*_f}, \frac{P^*_f}{P^*_a}, \frac{P^*_a}{P^*_l}, B, \mathbf{R} \right)
\]

where the price ratios are all expectations. de Janvry, Sadoulet, and Fafchamps also include transaction costs to explain research allocations. As the transaction costs for labor (supervising, negotiating, information costs) increase with farm size there will be an increasing bias in research toward labor-saving technology if large farmers’ demands prevail. Conversely, the transaction costs for land decrease with farm size because the fixed cost in land transactions implies that the price of land declines with farm size. This effect decreases the bias toward land-saving technology if small farmers’ demands prevail. de Janvry, Sadoulet, and Fafchamps substitute (3) and (4) into (1) and (2) so that the determinants of the optimal technical changes and factor ratios appear in their reduced form equations.2

The same approach is followed here, but the model is applied to over 100 years of historical data. The factor ratios are assumed to be functions of factor prices and the past public and private R&D expenditures, that generated the technologies and extension expenditures that transmitted the results to the farmers, thereby diffusing the technology. Farm size is included as it is expected to be a cause of factor-saving biases and electrification is considered as a proxy for rural infrastructure.

Data

Relatively few series are required to estimate this version of the induced innovation model. From 1910 onward the USDA has annual data for the quantities and prices of land, labor, fertilizer, and power, so the main difficulty is to decide how the series should be defined. Prior to 1910, the basic source is Hayami and Ruttan (1985) and in most cases their quinquennial data have been interpolated and then spliced to the USDA indices from 1910 onward. These series are defined and referenced in the data appendix and are available upon request.

There are several differences from the series constructed by Hayami and Ruttan (1985), the most significant being the treatment of animal and mechanical power. Rather than using horsepower, the series used here are the service flows from the capital stocks of mechanical power and machinery and horses and mules. The service flows are depreciation and running costs and the Tornqvist-Theil index is used in aggregation to give a single series. The corresponding price series are similarly a Tornqvist-Theil share-weighted average, whereas Hayami and Ruttan ignored the prices of the draft animals, which were the major source of power in the early part of the period. Olmstead and Rhode point out that draft animal power needs to be included and that the price fell far more slowly than for machinery.

The R&D expenditures of the public and private sectors, from 1888 to 1990, are available in Huffman and Evenson (1993), who also have public extension expenditures from 1915. Prior to 1956, their private R&D series are constructed from the results of regressions of private R&D on patenting. Prior to 1915, the extension expenditures are assumed to be collinear with the funding going to the Office of Experiment Stations (Huffman and Evenson 1993).

The dotted line in figure 1 shows that the ratio of animal and mechanical power to labor was increasing from 1880 to 1920, as more land was brought into cultivation. The solid line is the same factor ratio, net of factor substitution, which follows a similar path.3 During this period farmers received land at little or no cost and there was a high rate of capital formation (Cochrane). After 1920, the area of agricultural land remained fairly stable, labor in agriculture declined and mechanization increased.4 Although mechanization in agriculture began in the 1930s, figure 1 shows that the machinery/labor ratio actually fell during the Great Depression as the decline in the farm population was temporarily reversed. In the 1930s farmers lacked financial resources, savings, income or credit to make the outlays

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2 Unfortunately, it has not been possible to incorporate more formal dynamic adjustment costs (Stefanou, Luh and Stefanou) in this time series approach.

3 The elasticity of substitution estimates are reported later.

4 Employment opportunities in the cities increased due to the wartime and post-war boom. Tests follow later to determine whether labor was pushed off the land by mechanization, or pulled off by better paying jobs in the cities.
required to adopt new technologies, or to substitute capital for labor, but by the Second World War these financial constraints had lifted (Cochrane). The factor ratio then increased more rapidly from after the Second World War until it stabilized in the 1970s.

The third (notched) line plotted in figure 1 shows that the labor/machinery price ratio rose over the period, like the factor ratios, and that the series often have the same turning points, particularly for the wars and the depression periods. However, there is actually a negative correlation from about 1898 to 1915. As the acreage expanded, the numbers of horses and mules grew in response to demand, despite their increasing relative price, as Olmstead and Rhode have noted. This is contrary to the induced innovation hypothesis, which fares poorly during periods when the agricultural land area is increasing. Thirtle, Townsend, and van Zyl similarly show that in the period after the Second World War, the arable area in South Africa expanded, and more animal and mechanical power was used, although the price of labor was falling rapidly relative to the prices of these inputs.

Figure 2 shows that the land/fertilizer price ratio rises throughout the period, except during the First World War and the Depression, just as Olmstead and Rhode have pointed out. But, the fertilizer/land factor ratio, either in gross terms, or net of factor substitution, also increased for the entire period, except during the First World War and the Great Depression. Thus, Olmstead and Rhode’s observation that the price trends in these periods did not fit Hayami and Ruttan’s stylized facts is reflected in the behavior of the factor ratios, rather than being at odds with the induced innovation hypothesis.

To the extent that the factor ratio recovered slightly while the price ratio was still falling in the Depression can be explained by the need to counter soil exhaustion in the era of the Oklahoma dust bowl and the significant advances in biological technologies, which made fertilizer more productive. From the 1940s there is a fairly rapid rise in the factor and price ratios, until both respond with blips, to the oil price shock in the early 1970s. Then the factor ratio ceases growing in the 1980s, following the turn in the price ratio, just as the inducement hypothesis predicts. However, the fact that the series are all non-stationary or trended does mean that the correlations may be spurious and that time series tests are required to establish the properties of the series and then to establish cointegration.

**Figure 1. Correlation of the machinery/labor factor ratio and the labor/machinery price ratio**

**Time Series Tests**

If the variables are cointegrated, then there is a valid relationship and divergence from a stable equilibrium state must be stochastically bounded (Banerjee et al., p. 136). In simple cases, two conditions must be satisfied for variables to be cointegrated. First, the series for the individual variables should be integrated
of the same order, which they are, since all the variables are non-stationary in the levels, but stationary in first differences. Second, a linear combination must exist that is integrated of an order one less than the original variables. That is, the error terms from the cointegrating regressions must be stationary. This methodology was used in the analysis of induced innovation in South Africa by Thirtle, Townsend, and van Zyl, but their tests are applied to shorter series and are somewhat dated, so more recently developed tests are used here (Kwiatkowski, et al., Schmidt and Phillips).

Orders of integration are more difficult to determine when structural breaks occur in the series. Figures 1 and 2 suggest periods of significant structural changes between the World Wars. Thus, the first step is to check for structural breaks which could affect the properties of the variables. Perron tests, originally used to identify the date of the oil price shock in the 1970s, find a structural break with persistent effects from 1920 as do the unit root tests (Maddala and Kim). This is when the Great Depression first hit agriculture and the agricultural land area stabilized, nine years before the stock market crash (Cochrane, p. 100).

The low power of the most common unit root tests (Dickey-Fuller, Phillips-Perron) make the break likely to influence the test outcomes. As low power increases, the probability of making a type-II error, the KPSS (Kwiatkowski et al.) test is used, in which the null hypothesis of stationarity is the reverse of the usual unit root test, and the break is included following Lee et al. The KPSS results are checked using the Schmidt and Phillips (SP) test for the familiar unit root hypothesis that the coefficient on the lagged term in the random walk should be equal to unity. The SP test is chosen because it is a Lagrange Multiplier test like the KPSS test. The known break is then incorporated in the SP test following Amsler and Lee.

Table 1 begins with the KPSS test results, which find all the variables to be non-stationary in the levels (the test statistic is greater than the critical value at the 95% level of confidence) and stationary in first differences (the test statistic is less than the critical value). The SP test results in the next column corroborate these results, since at the 91% confidence level, most variables are non-stationary in the levels and stationary in first differences. The exceptions are the one case where the test was inconclusive and the results for private research expenditures and wages, which are both stationary in the levels, at the 95% confidence level. These results seem to indicate that the series are long enough, for the annual data from 1910 onward to outweigh the effects of interpolation and extrapolation in the 1880–1910 period.

Apart from establishing that all of the factor and price ratios are integrated of order one and may be cointegrated, these results indicate

5 This tends to make stochastic series look rather more like trended series since a lot of the movement is obviously destroyed. The tests on the time series properties of the series were applied to the whole period and also from 1910 onward and fortunately the results are not different.
that for these annual data, the induced innovation hypothesis should be formulated in the original manner (Hayami and Ruttan 1985), with factor ratios as a function of factor/price ratios, in either the levels or the changes. Olmstead and Rhode’s observation that the two models differ, is taken care of in the ECM which uses the changes to model the shorter run and the levels to test for long-run equilibrium. Tiffin and Dawson’s suggestion that the hypothesis must be altered to make changes in factor ratios a function of factor/price ratios, because the factor ratios may be integrated of order two, can be rejected. With more and better data there is no evidence of this problem.

Cointegrating Relationships and the Error Correction Representation

Having established that the variables are integrated of the same order one, the next stage is to test for cointegrating vectors which imply that non-spurious long-run relationships exist between the variables. The Johansen procedure used is a maximum likelihood test for cointegration (Johansen, Johansen and Juselius) allowing estimation of cointegrating relationships by directly testing the number of cointegrating vectors and the direction of causality.

Table 2 begins with the equation explaining the machinery/labor ratio, where the possible explanatory variables were the own price ratio, the land/labor price ratio, public and private R&D, extension and farm size. The maximum lag length of the VAR was found to be five years, using the Schwartz criterion and with this specification, both the tests in the Johansen procedure indicate that there is a single cointegrating vector for the variables shown. The land/labor price ratio was not significant and extension was collinear with public and private R&D and expenditures. The tests showed that the preferred model retained the private R&D stock, but not public R&D and extension. This result is hardly surprising since the farm machinery industry is responsible for most mechanical innovations (Binswanger 1984) and it is independent of the extension service.

The Schwartz criterion was also used to determine the lag on R&D, which was found to be thirteen years. This is a reasonable time period for the machinery stock to adjust and is close to the eleven-year lag reported by Khatri,

### Table 1. Testing the Variables for Order of Integration with a Structural Break in 1920

<table>
<thead>
<tr>
<th>Variable Name and Abbreviation</th>
<th>KPSS Tests</th>
<th>SP Tests</th>
</tr>
</thead>
<tbody>
<tr>
<td>Log of ratio machinery/labor (M/L)</td>
<td>0.1775</td>
<td>-3.504</td>
</tr>
<tr>
<td>Δ Log of ratio machinery/labor</td>
<td>0.1221*</td>
<td>-3.750**</td>
</tr>
<tr>
<td>Log of ratio price machinery/labor (P_M / P_L)</td>
<td>0.2145</td>
<td>inconclusive</td>
</tr>
<tr>
<td>Δ Log of ratio price machinery/labor</td>
<td>0.1356*</td>
<td>inconclusive</td>
</tr>
<tr>
<td>Log of ratio fertilizer/land (F/A)</td>
<td>0.1484</td>
<td>-2.472</td>
</tr>
<tr>
<td>Δ Log of ratio fertilizer/land</td>
<td>0.1163*</td>
<td>-7.092**</td>
</tr>
<tr>
<td>Log of ratio price fertilizer/land (P_F / P_A)</td>
<td>0.2193</td>
<td>-3.568</td>
</tr>
<tr>
<td>Δ Log of ratio price fertilizer/land</td>
<td>0.1356*</td>
<td>inconclusive</td>
</tr>
<tr>
<td>Log of ratio price land/labor (P_A / P_L)</td>
<td>0.1729</td>
<td>-3.623</td>
</tr>
<tr>
<td>Δ Log of ratio price land/labor</td>
<td>0.1222*</td>
<td>-3.782**</td>
</tr>
<tr>
<td>Log of real public R&amp;D expenditures (RDPUB)</td>
<td>0.2287</td>
<td>-3.562</td>
</tr>
<tr>
<td>Δ Log of real public R&amp;D expenditures</td>
<td>0.1136*</td>
<td>-4.047**</td>
</tr>
<tr>
<td>Log of real private R&amp;D expenditures (RDPRI)</td>
<td>0.1692</td>
<td>-3.562</td>
</tr>
<tr>
<td>Δ Log of real private R&amp;D expenditures</td>
<td>0.1259*</td>
<td>-3.517</td>
</tr>
<tr>
<td>Log of real extension expenditures (EXT)</td>
<td>0.1839</td>
<td>1.51</td>
</tr>
<tr>
<td>Δ Log of real extension expenditures</td>
<td>0.1283*</td>
<td>-4.257**</td>
</tr>
<tr>
<td>Log of farm size (size)</td>
<td>0.1705</td>
<td>-3.249</td>
</tr>
<tr>
<td>Δ Log of farm size</td>
<td>0.1144*</td>
<td>-3.674**</td>
</tr>
<tr>
<td>Log of electrical power (ELEC)</td>
<td>0.1757</td>
<td>-3.584</td>
</tr>
<tr>
<td>Δ Log of electrical power</td>
<td>0.1188*</td>
<td>-3.772**</td>
</tr>
<tr>
<td>Log manufacturing wage/farm wage (W_M / W_F)</td>
<td>0.1976</td>
<td>-5.573*</td>
</tr>
<tr>
<td>Δ Log manufacturing wage/farm wage</td>
<td>0.1047*</td>
<td>-3.393</td>
</tr>
<tr>
<td>Critical values</td>
<td>0.146 (95%)</td>
<td>-3.63 (99%)</td>
</tr>
</tbody>
</table>

* Indicates stationarity at 95% confidence level.
** Indicates stationarity at 99% confidence level.
Thirle, and van Zyl for South Africa. Granger’s Representation Theorem shows that in asymptotic results such lags should not be necessary, but with limited data imposing more structure in this way made the cointegrating relationship considerably stronger.

For the fertilizer/land equation, the list of possible independent variables is similar to the first equation, with the addition of rural electrification which was included as a proxy for infrastructure development, which was expected to affect the price of land.6 In this case, the own price ratio, the land/labor price ratio, the public R&D stock and farm size all proved to be significant in the cointegrating equation. Although the trace test version of the Johansen model identifies two or possibly three cointegrating vectors, the eigenvalue test indicates only one and no other combination of variables cointegrated according to both tests. The maximum lag in the VAR was again five years and the R&D lag was found to be twenty-three years. This is long in comparison with most studies, but in the only previous application of similar tests to time series this long, Pardey and Craig found that lags of up to thirty years were necessary to capture all the effects of public R&D on agricultural output.

### Estimation and Results of the Error Correction Model (ECM)

Granger’s Representation Theorem proves that a cointegrated system of variables can be adequately represented as an ECM (Engle and Granger) and Campos, Ericsson, and Hendry show that in the presence of a structural break, error correction models are generally more powerful than Engle and Granger’s two-step procedure. Using the variables for which cointegration was established, as shown in table 2, the ECMs for the two equations are

$$\Delta \ln(M/L)_t = \varphi_0 + \sum_{i=1}^{2} \varphi_{1i} \Delta \ln(M/L)_{t-i} + \sum_{i=0}^{2} \varphi_{2i} \Delta \ln(P_M/P_L)_{t-i} + \sum_{i=1}^{15} \varphi_{3i} \Delta \ln RDPRI_{t-i} + \lambda \left[ \ln(M/L)_{t-1} - \alpha_1 \ln(P_M/P_L)_{t-1} - \sum_{i=1}^{15} \alpha_{2i} \Delta \ln RDPRI_{t-i} \right]$$

$$\Delta \ln(F/A)_t = \delta_0 + \sum_{i=1}^{2} \delta_{1i} \Delta \ln(F/A)_{t-i} + \sum_{i=0}^{2} \delta_{2i} \Delta \ln(P_F/P_A)_{t-i} + \sum_{i=0}^{2} \delta_{3i} \Delta \ln(P_A/P_L)_{t-i} + \sum_{i=1}^{25} \delta_{4i} \Delta \ln RDPUB_{t-i} + \sum_{i=1}^{3} \delta_{5i} \Delta \ln SIZE_{t-i} + \lambda \left[ \ln(F/A)_{t-1} - \beta_1 \ln(P_F/P_A)_{t-1} - \beta_2 \ln(P_A/P_L)_{t-1} - \sum_{i=1}^{25} \beta_{3i} \Delta \ln RDPUB_{t-i} - \sum_{i=1}^{3} \beta_{4i} \Delta \ln SIZE_{t-i} \right]$$

### Table 2. Johansen Cointegration Results

<table>
<thead>
<tr>
<th>Variables Tested</th>
<th>Test Statistic (95% Critical Value)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Eigenvalue Test</td>
</tr>
<tr>
<td>M/L $P_a/P_L$ Private R&amp;D stock (−13)</td>
<td>23.54 (22.00)</td>
</tr>
<tr>
<td></td>
<td>13.23 (15.67)</td>
</tr>
<tr>
<td>F/A $P_f/P_a$ $P_a/P_L$ Public R&amp;D stock (−23) SIZE</td>
<td>34.69 (34.4)</td>
</tr>
<tr>
<td></td>
<td>24.18 (28.1)</td>
</tr>
<tr>
<td></td>
<td>20.21 (22.0)</td>
</tr>
</tbody>
</table>

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6 Olmstead and Rhode argue that improvements in transport and communications led to increases in land values. Since the railroad system was largely in place by 1880, electrification is a better proxy for infrastructure.
price coefficients are the direct elasticities of either equation and were dropped. and electrical power were not significant in average farm size. Extension expenditures RDPUB, is public research stock, and SIZE, is price ratio, RDPRI, is private research stock, variables are in natural logarithms.7 The own parts. The coefficients, which were used in conjunction with the relative price changes, to calculate the changes in the factor ratios that resulted from factor substitution. The total changes, net of this factor substitution, were plotted in figures 1 and 2. The coefficient λ, in the second part of the ECM, is the adjustment elasticity, which for stability, must be negative. This indicates that when the system is not at long-run equilibrium, it will be moving toward it.

The results of the ECMs reported in table 3 were chosen on the criteria of goodness of fit (variance dominance), data coherence (white noise error process), parameter parsimony and consistency with theory (Hendry and Richard). Seemingly unrelated estimation was used to gain efficiency, since the errors of the two equations are correlated, as they should be, since the prices of machinery and land are correlated.

For both equations the results are reported for the full period (the starting dates allow for the lags on R&D) and from 1920 onward, both because of the structural break in 1920 and to avoid using the early data that involved the lag. The shorter run may have been less erratic, rather than the production economics view in which at least one factor is fixed.

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7 The term “shorter-run” in the ECM context means that the system is not in long-run equilibrium, which in a model that includes R&D expenditures implies that technology generation and adoption must occur for equilibrium to be attained. The shorter run may be an lesser period, rather than the production economics view in which at least one factor is fixed.

### Table 3. Unrestricted ECM Results

<table>
<thead>
<tr>
<th></th>
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<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>CONSTANT</td>
<td>φ₀</td>
<td>−0.042 (−0.9)</td>
<td>−0.259 (−2.1)</td>
<td>6.373 (−3.6)</td>
<td>−3.050 (−2.6)</td>
<td></td>
</tr>
<tr>
<td>Δ(M/L),₋₁</td>
<td>φ₁</td>
<td>0.299 (3.0)</td>
<td>0.251 (2.2)</td>
<td>0.585 (−3.9)</td>
<td>0.586 (−3.8)</td>
<td></td>
</tr>
<tr>
<td>Δ(M/L),₋₂</td>
<td>φ₂</td>
<td>0.307 (2.9)</td>
<td>0.245 (2.0)</td>
<td>0.598 (−5.1)</td>
<td>−0.603 (−5.4)</td>
<td></td>
</tr>
<tr>
<td>Δ(P_M/P_L)₋₁</td>
<td>φ₃</td>
<td>−0.068 (−1.6)</td>
<td>−0.080 (−1.5)</td>
<td>0.111 (2.1)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δ(P_A/P_L)₋₁</td>
<td>φ₄</td>
<td>−0.068 (−1.6)</td>
<td>−0.080 (−1.5)</td>
<td>0.111 (2.1)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Stock Public</td>
<td>λα₁</td>
<td>0.009 (1.2)</td>
<td>0.047 (2.2)</td>
<td>0.552 (3.4)</td>
<td>0.394 (1.8)</td>
<td></td>
</tr>
<tr>
<td>Stock Private</td>
<td>λα₂₋₁₂</td>
<td>0.009 (1.2)</td>
<td>0.047 (2.2)</td>
<td>0.552 (3.4)</td>
<td>0.394 (1.8)</td>
<td></td>
</tr>
<tr>
<td>Farm Size₋₁</td>
<td>λβ₄₋₁</td>
<td>0.35</td>
<td>0.39</td>
<td>0.34</td>
<td>0.38</td>
<td></td>
</tr>
<tr>
<td>Adj. R²</td>
<td>1.82</td>
<td>1.73</td>
<td>1.97</td>
<td>1.91</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: The critical t-values for one-tailed tests are 1.29 for 90% confidence, 1.66 for 95% and 1.98 for 97.5%.
break. The coefficient of the constant term is insignificant and the two significant lagged changes in the dependent variable are required to remove the autocorrelation from the system. The shorter-run elasticity of interest is the own price coefficient, which may be interpreted as the direct elasticity of substitution (\(c_0\) here and \(c_1\) in equation (1)). This is significant and the value of \(-0.08\) in the shorter period suggests that the shorter-run substitution possibilities were extremely limited, relative to South Africa, where Thirtle, Townsend, and van Zyl estimated the equivalent elasticity as \(-0.37\). However, that figure must be inflated by the perverse labor policies such as the Pass Laws, which caused artificial labor scarcity and forced farmers to substitute machinery. The price of capital was also artificially cheapened by tax incentives and cheap credit.

The fertilizer/land model performed equally well in terms of explaining the deviations in the dependent variable for both time periods. The shorter-run coefficient on the fertilizer-to-land price ratio, of \(-0.58\), is the direct partial elasticity of substitution (\(\delta_2\) here and \(\delta_2\) in equation 2), which is low, but higher than for South Africa, where the Thirtle, Townsend, and van Zyl estimate was \(-0.39\). This suggests that the shorter-run substitution possibilities between land and fertilizer are not as limited as for machinery and labor. The land/labor price ratio coefficient is also negative and highly significant in the shorter run, suggesting that labor and fertilizer are substitutes.\(^8\)

In the long run, the error correction term, \(\lambda\), for the machinery/labor equation is \(-0.02\). This indicates little adjustment toward the long-run equilibrium level in the current period, with full adjustment taking as much as fifty years. This is perhaps excessive, even for durable capital items, but for the shorter period the value of \(-0.06\) suggests full adjustment in just under seventeen years, which is only slightly longer than the fifteen years suggested by Antle. The negative sign shows that the direction of correction is toward equilibrium, which is essential for ECM stability.

The negative coefficient on the long-run own price variable indicates that a decrease in the machinery/labor price ratio generates labor-saving technological change, in the manner predicted by the induced innovation hypothesis.\(^9\) However, the poor quality of the early data is perhaps responsible for the insignificance of the key variables in the full period. For the shorter period, the own price ratio is significant and so is private R&D (whereas public R&D was not).

The fertilizer/land model is more robust and gives significant and similar results for both periods. The error correction term for the fertilizer-to-land equation is \(-0.44\), which indicates quite rapid adjustment toward the long-run equilibrium level in the current period, with full adjustment taking a little more than two years. The negative sign shows that the direction of correction is toward equilibrium. The negative coefficient on the long-run own price variable indicates that a decrease in the fertilizer/land price ratio generates land-saving technological change, in the manner predicted by the induced innovation hypothesis. The negative sign on the land/labor price coefficient indicates substitutability and is in agreement with Frisvold’s prediction. Public R&D expenditures are positive and significant, leading to increased fertilizer use. The farm size parameter is also positive and significant, indicating that larger farms are more fertilizer-intensive. While the R&D result is consistent with de Janvry et al., the farm size outcome is not. However, this is not a matter of small farmers using less fertilizer than large farmers. The farm size measure in this time series is just the average farm size for each year, and since farms both grew larger and used more fertilizer in later years, this result is generated. Extension and electrification, the rural infrastructure variables, are not significant in the cointegration tests or the ECM.

The long-run elasticities can be calculated from the results in table 3 simply by dividing the long-run coefficients by \(-\lambda\), to preserve the signs on the estimated coefficients. Whereas the shorter-run elasticities of substitution in table 3 may be viewed as movements around the isoquants, the long-run equivalents can be viewed as movements around innovation possibility surfaces, which encompass all the techniques which can be developed, given the state of scientific knowledge (following Ahmad’s definition of the innovation possibility curve). If this meaning is attributed to disequilibria in the system, it implies that long-run equilibrium

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8 Perhaps the most obvious omitted variable is pesticide, which does not fit naturally into the model and is most likely to have been a substitute for labor. Pesticide can be included (it is positively correlated with the machinery/labor ratio) but only at the cost of dropping one of the other variables which are more directly relevant to the current argument.

9 In this aggregate time series, the farm size measure only varies with time, so all this establishes is that over time farms get bigger and use more machinery. The causality issue is pursued later.
Table 4. Long-Run Elasticities for the ECM

<table>
<thead>
<tr>
<th>Variable</th>
<th>$P_M / P_L$</th>
<th>$P_F / P_A$</th>
<th>$P_A / P_L$</th>
<th>R&amp;D Stock</th>
<th>Size</th>
</tr>
</thead>
<tbody>
<tr>
<td>$M/L$ Equation 1894–1990</td>
<td>−0.84</td>
<td>−</td>
<td>NS</td>
<td>0.41</td>
<td>NS</td>
</tr>
<tr>
<td>$M/L$ Equation 1920–1990</td>
<td>−0.64</td>
<td>−</td>
<td>NS</td>
<td>0.72</td>
<td>NS</td>
</tr>
<tr>
<td>$F/A$ Equation 1904–1990</td>
<td>−0.61</td>
<td>−0.61</td>
<td>0.16</td>
<td>1.26</td>
<td></td>
</tr>
<tr>
<td>$F/A$ Equation 1920–1990</td>
<td>−0.58</td>
<td>−0.58</td>
<td>0.28</td>
<td>0.98</td>
<td></td>
</tr>
</tbody>
</table>

Table 5. Decomposition of ECM Results into Factor Substitution and Induced Factor Bias

<table>
<thead>
<tr>
<th>Period</th>
<th>Machinery/labor</th>
<th>Fertilizer/land</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Factor Substitution</td>
<td>Induced Innovation</td>
</tr>
<tr>
<td>1880–1920</td>
<td>31</td>
<td>69</td>
</tr>
<tr>
<td>1920–1990</td>
<td>19</td>
<td>81</td>
</tr>
</tbody>
</table>

is attained only after the adoption of innovations that are induced by changes in relative prices.\(^{10}\)

Thus, figure 1 showed the machinery/labor factor ratio net of own price effects. The figure suggests that when factor substitution has been accounted for, the major proportion of the change in factor ratios is still left unexplained. This remaining portion can be explained by the lagged effect of relative prices, causing technology-induced substitution around the IPC, private R&D expenditures that produced the mechanical technologies, and farm size which also drives the process. Table 4 shows that the low speed of adjustment for the machinery/labor equation means that the long-run own price effect, which includes innovations, is almost fifteen times as large as for the shorter run, with a value of 0.84 for the full period and 0.64 for post-1920.\(^{11}\) Similarly, a 1% increase in private R&D expenditures will, in the long run, increase the machinery/labor ratio by 0.41%, or 0.72% for the post-1920 period. Applying a similar model to South African data Thirtle, Townsend, and van Zyl found a long-run elasticity of substitution of 0.47 and an R&D impact of only 0.13%.

Figure 2 showed the fertilizer/land factor ratio net of substitution effects and again there remained a large portion of the change in factor ratios to be explained by the long-run effect of the price ratios, public R&D expenditure, and farm size. The long-run elasticities for the fertilizer/land equation are less different from the shorter run, due to the higher adjustment coefficient of −0.44. This gives an elasticity of long-run substitution, around the innovation possibility curve, of −0.61 for the full period. Coincidentally, the elasticity with respect to the land/labor price ratio is also −0.61. In the long run, a 1% increase in public R&D changes the factor ratio by 0.16% and a 1% increase in farm size increases the fertilizer/land ratio by 1.26%.

The relative shares of factor substitution and technical change, in explaining changes in the two factor ratios, are summarized in table 5. Over time, the share of technological change becomes greater, so that in the later period, factor substitution explains less than 20% of the change in the machinery/labor ratio and only one-third of the change in the fertilizer/land ratio. This demonstrates Hayami and Ruttan’s proposition that innovations are required to allow continuing, long-run substitution, over periods of this length.

The share of technology can be further decomposed into the effects of relative price changes, of R&D, and of increases in farm size. Although this is perhaps useful, it does not imply that these are separate effects. The

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\(^{10}\) The long run here allows for technological change since for the system to be in “long run” equilibrium the effects of R&D should have had time to take effect: this is often called the very long run.

\(^{11}\) This is well below the estimate of the long-run elasticity of substitution reported by Kislev and Peterson (1982), which was −1.7, but it is close to the previous studies they quote, in which Griliches and Binswanger’s results were around unity.
Table 6. Decomposition of Long-Run Technical Change in ECM Results

<table>
<thead>
<tr>
<th>Variable</th>
<th>Share of Change (%)</th>
<th>1880–1990</th>
</tr>
</thead>
<tbody>
<tr>
<td>Own price ratio</td>
<td>38</td>
<td>36</td>
</tr>
<tr>
<td>Land-labor price</td>
<td>–</td>
<td>14</td>
</tr>
<tr>
<td>R&amp;D stock</td>
<td>62</td>
<td>22</td>
</tr>
<tr>
<td>Farm size</td>
<td>–</td>
<td>28</td>
</tr>
</tbody>
</table>

R&D generated the price-induced effects and both of these changed farm size, which itself fed back into changing the factor ratios. Thus, all the effects are from R&D, which has a direct impact and indirect effects through prices and farm size. With this proviso, table 6 shows that private R&D accounts for almost two-thirds of the change in the machinery/labor ratio. The own price ratio accounts for the other third and has a very similar share in explaining the fertilizer/land ratio, but here the direct public R&D share falls to 22%, as the land/labor price ratio and farm size also make substantial contributions.

In summary, to corroborate the induced innovation hypothesis, the crucial requirement is that the coefficients on the lagged own-price terms should be negative and significant, which is true in both models. Also, since it is R&D that produces new technologies, its significance in affecting both factor ratios directly, further corroborates the hypothesis. Finally, public R&D was concentrated on biological technology, so it produced land-saving technical change, whereas much private R&D was in the farm machinery industry, so it resulted in labor-saving innovations. These roles may not be as clear as they once were, since the private sector has been moving into the biological area over the last ten years, to exploit the promise of biotechnology. Recent results by Huffman and Evenson (2001) show that between 1950 and 1982, private R&D and market forces had a larger impact on farm size and specialization in crops than they had in the livestock sector.

Causality Tests

The final stage in the procedure is to apply causality tests to the ECMs to ensure that it is also true that the price ratios are causally prior to the factor ratios. Causality tests within the ECM framework can be conducted by testing the loading matrix in Johansen’s model. If the α matrix has a complete column of zeros, no cointegrating vector will appear in a particular block of the model, indicating no causal relationship. Direct Wald tests on the loading parameters (Hall and Milne; Hall and Wickens) are used to determine these restrictions. The rank tests on the loading matrices, reported in table 7 confirm causality from own price ratios to factor ratios and indicate that disequilibrium in the machinery/labor factor ratio does not feed back to the machinery/labor price. There is, however, feedback to private R&D because both are significant.

This two-way causality is unfortunate since private R&D is causing mechanization and the machinery/labor ratio is also causing private R&D to increase. So, labor could have been pushed off by mechanization, or pulled off by rising urban wages, as suggested by Kislev and Peterson. This issue remains unresolved and is further investigated in the next section.

The dynamics of the causal relations in the fertilizer/land system are similar. The results indicate causality from own price ratios to factor ratios and that disequilibrium in the fertilizer/land factor ratio does not feed back to the fertilizer/land price. There is feedback from disequilibrium in the fertilizer/land factor ratio to the land/labor price ratio and farm size. This suggests that the land/labor price ratio and farm size are not weakly exogenous and it may have improved the results to create

Table 7. Causality Tests. Wald Test Results of Zero Restrictions on the Loading Matrix, α

<table>
<thead>
<tr>
<th>Machinery/Labor Factor Ratio</th>
<th>Machinery/Labor Price Ratio</th>
<th>Stock of Private R&amp;D</th>
</tr>
</thead>
<tbody>
<tr>
<td>3.38*</td>
<td>0.1</td>
<td>139.23**</td>
</tr>
<tr>
<td>12.55**</td>
<td>0.37</td>
<td>3.65*</td>
</tr>
<tr>
<td>51.06**</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Key: **Significant at 99%; *significant at 90%.
instruments for these variables. This was not done because it was possible to establish cointegration in both the machinery/labor and fertilizer/land equations without instrumented variables.

Wage-Pull or Technology-Push?

The choice between the competing explanations of machinery-labor substitution were not resolved above because the causality was two-way. Kislev and Peterson (1981) and especially Peterson and Kislev suggest that labor may leave agriculture because of better opportunities in urban areas, as opposed to being forced off the farm by mechanization. Table 8 reports the results of cointegration tests intended to distinguish between these two causes.

The first two rows show that the machinery (including animals) and labor series do cointegrate, according to both tests, for both periods. This implies causality in at least one direction. If the decline in labor is causally prior to the increase in machinery, it would suggest that there is no sense in which labor was forced off the land by mechanical technology. Conversely, if the rise in the use of machinery is Granger-prior (1969) to the decline in labor, then it would seem probable that the labor was displaced by mechanical innovations. The coefficients are not reported to save space, but single-equation Granger causality tests support the proposition that the decline in labor is causally prior to mechanization far more strongly than the labor displacement argument. Five lags in the VAR are sufficient to give white noise errors and labor is found to be causally prior to machinery at lags of two, three and five years, at the 95% confidence level. Testing for reverse causality, machinery is Granger-prior to labor only with an eight-year lag. This is reasonable evidence that mechanization is required because labor leaves, but a second test below is more powerful.

The second two rows show that according to both tests, the machinery/labor factor ratio ($\frac{M}{L}$) and the ratio of the change in the manufacturing/farm wage ($\frac{W_U}{W_R}$) are cointegrated. The coefficients are not interesting, but single-equation Granger causality tests show unidirectional causality, with changes in the ratio of urban-to-rural wages causally prior to the machinery-labor ratio at lags 1 through 5, and no reverse causality from the $\frac{M}{L}$ ratio to changes in the ratio of wages at the same lag lengths with 90% confidence or better. This supports the hypothesis that labor was pulled off the land by the higher income opportunities in urban areas. Thus the technology-push hypothesis is rejected and the wage-pull hypothesis is supported by these simple tests, but other unavailable variables, such as increasing levels of rural education, that enable the farm population to move to urban jobs, must also have a role.

Table 8. Wage-Pull versus Technology-Push Explanations of Labor Displacement

<table>
<thead>
<tr>
<th>Variable</th>
<th>Johansen Test (95% Critical Value)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Eigenvalue</td>
</tr>
<tr>
<td>$M$ and $L$ (1882–1990)</td>
<td>20.8 (15.7)</td>
</tr>
<tr>
<td>$M$ and $L$ (1920–1990)</td>
<td>19.8 (15.7)</td>
</tr>
<tr>
<td>$M/L$ and change in $W_U/W_R$</td>
<td>34.3 (15.7)</td>
</tr>
<tr>
<td>(1886–1990)</td>
<td>3.1 (9.2)</td>
</tr>
<tr>
<td>$M/L$ and change in $W_U/W_R$</td>
<td>26.2 (15.7)</td>
</tr>
<tr>
<td>(1920–1990)</td>
<td>2.1 (9.2)</td>
</tr>
</tbody>
</table>

*The VAR lengths are 2 over both periods for the $M$ and $L$ equation and 5 over both periods for the change in the wage ratio equation as selected by the Schwartz Criterion.

Conclusion

This article tests Hayami and Ruttan’s (1985) induced innovation hypothesis using annual data for U.S. agriculture from 1880 to 1990. The two-stage CES production function leads to estimating equations that directly test the inducement hypothesis, by making factor ratios (net of factor substitution) functions of factor/price ratios. The model also incorporates the variables that generate new technologies (public and private R&D and extension) and farm size, which also affect factor-saving biases.

The results corroborate the inducement hypothesis with respect to the role of factor/price ratios in generating long-run changes in factor ratios net of factor substitution. The relationships are shown to cointegrate, the correlation is negative, as predicted and the causality runs from the price ratios to the factor ratios. The results also show public and private research expenditures and farm size are further important determinants of the observed rates and biases of technological change. Hence, the biases
of technical change depend not only on input price ratios, as in Hayami and Ruttan’s formulation, but are also a function of R&D expenditures and farm structure, as de Janvry et al. suggested. Consistent with de Janvry et al., public R&D expenditures generate land-saving technical change.

Tests to determine whether labor was pulled off the land by urban wages, as suggested by Kislev and Peterson (1982), or pushed off by new technology, showed that the first hypothesis was supported, whilst the second was rejected. However, the forces driving structural change must be more complex and should include the role of rural education in making non-farm jobs more attainable and the less measurable pull factor of urban amenities.

References


Appendix: The Data

Land


Fertilizer


Labor


Animal and Mechanical Power and Machinery

This item is the service flow (depreciation and running costs) from capital stocks of machinery and animals, which required aggregation.

Quantity series. 1880–1910, Mechanical Power and Machinery, USDA Input Index Catalogue, item 3000, Mechanical Power and Machinery, decennial observations, with interpolation. 1910–26, Mechanical power and machinery, Table 55, from Changes in Farm Production and Efficiency, USDA, ESCS, Statistical Bulletin No.612, 1977 (base 1967). Updated to 1990 from Table 36, Production and Efficiency Statistics, 1991, USDA, ERS, ECIFS 10-3. Then, the series for 1926–90 was re-flated using the USDA Price Index (see below) and re-deflated.
with the BLS Index for Agricultural Machinery and Equipment from BLS Labstat Series Report, series WF0111 (also available in U.S. Department of Labor, Bureau of Labor Statistics, Handbook of Labor Statistics). The series constructed can be viewed as depreciation and running costs for farm machinery and equipment, evaluated using the USDA’s rules, but scaled up to allow for the BLS index treatment of quality improvement.

**Quantity series.** 1880–1950, Horses and Mules, from Historical Statistics of the United States: Colonial Times to 1970 (annual data from 1867). 1950–60, Agricultural Statistics, various issues. From 1960-90, the figures are only available for the agricultural census years, as Hayami and Ruttan (1985) explain. These figures are for capital stocks, so to derive a service flow a constant price value series was generated and interest was fixed at four percent for the Tornguist-Theil weights, which is the figure used in capital stock calculations by the USDA. The average working life of horses and mules was assumed to be eight years, so they are depreciated at 12.5% per annum. The running cost element of the service flow is taken to be mainly the cost of feed, and a stall-fed animal is assumed to require three tons of hay and a ton of oats per annum. However, farm animals may be put out to graze when not working intensively, so only an arbitrary 50% of the feed requirement was incorporated. This intake is valued at a constant price and added to the depreciation, to complete the service flow of horses and mules.

**Price series.** Machinery and Equipment. These figures follow the approach of Hayami and Ruttan (1985), Table C-2. 1880–1890, Warren-Pearson Wholesale Price Index for Metal and Metal Products, From Historical Statistics. 1890–1910, BLS Wholesale Price Index for Metal and Metal Products, From Historical Statistics. 1910–60, Farm Machinery, Indexes of Average Prices, Table 85, in, USDA, ERS, Statistical Bulletin No.368, Costs and Returns on Commercial Farms: Long Term Study, updated to 1990 from Table 16, USDA Database at the Albert R. Mann Library at Cornell University (USDA.MANNLIB.CORNELL.EDU). However, the USDA series lack the quality adjustment of the BLS series that is available from 1926 to 1990. So, from 1926 onward, the BLS index is used; Agricultural Machinery and Equipment, Table 127, in U.S. Department of Labor, BLS, Handbook of Labor Statistics, Bulletin 2000, 1979. This is updated to 1990 with the BLS Index for Agricultural Machinery and Equipment from BLS Labstat Series Report, series WF0111.

**Price series.** Horses and Mules, from Historical Statistics of the United States: Colonial Times to 1970 (annual data from 1867). Both the numbers and values are given in Series K 195–212, so prices can be retrieved with no difficulty, and updated to 1960, using Agricultural Statistics, various issues. From 1960 to 1990, the figures are only available for the agricultural census years, as Hayami and Ruttan (1985) explain, so the series used is Feeder Livestock, Indexes of Prices Paid, from USDA, Economics and Statistical Service, Crop Reporting Board, Agricultural Prices: Annual Summary, 1980, updated to 1990 from USDA, Agricultural Statistics, 1992. Note that by 1960, the share of animals in the total is only just over 2%, so the errors imparted cannot be important.

**Aggregation.** Input and Price series for Animals, Machinery and Equipment. The approach taken is to form a value-share weighted aggregate of animal and mechanical power service flows. The shares begin at two thirds to one third in 1880, but machinery overtakes in 1915 and by 1990 the animals account for only a little over 1% of the total value.

**R&D and Extension**

Public R&D from Huffman and Evenson (1993). The figures for 1880–87 were extrapolated. Private R&D from Huffman and Evenson (1993). The figures are decennial until 1956, then annual. Public Extension From Huffman and Evenson (1993). The series for total extension begins at 1915. From 1888 to 1915, the extension series is assumed to be collinear with the expenditures of the Office of Experiment Stations (Huffman and Evenson 1993, Table 4.2). The figures for 1880–87 were extrapolated.