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Testing the induced innovation hypothesis: an error correction model of South African agriculture

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Abstract

This paper applies cointegration techniques to a model of induced innovation based on the two-stage constant elasticity of substitution (CES) production function. This approach results in direct tests of the inducement hypothesis, which are applied to data for South African commercial agriculture for the period 1947–1991. South African data is used because the policy changes have been substantial enough that the factor and price ratios have turning-points, rather than being monotonic. The time series properties of the variables are checked, cointegration is established, and an error correction model (ECM) constructed, allowing factor substitution to be separated from technological change. Finally, the ECM formulation is subjected to causality tests, which show that both the factor price ratios and R&D and extension expenditures are Granger-prior to the factor-saving biases of technological change. Thus, each stage of the analysis corroborates the inducement hypothesis. However, straightforward price-inducement is only part of the explanation of changes in factor ratios. Policy-induced innovation, in response to tax concessions and subsidised credit, is also present.

Keywords: Induced innovation hypothesis; Error correction model; South African

1. Introduction

The notion that at least some inventions may be induced by economic forces has been entertained by historians and economists at least since the 1920s (Mantoux, 1928; Hicks, 1932). There are by now several formulations of the relationship and well over 100 empirical tests (Thirtle and Ruttan, 1987), the great majority of which corroborate some form of the inducement hypothesis. In many cases, the hypothesis is not clearly stated, and the tests amount to no more than establishing a correlation between a measure of factor scarcity and an indicator of the direction, or factor-saving biases, of technical change. In some cases, the results are poor; in nine tests intended to show that changes in the land/labour price ratio cause changes in the land/labour ratio, Ruttan et al. (1978) find that five outcomes are inconsistent with the hypothesis. This leads them to postulate 'an innate labour saving bias' in the technological possibilities. Thus, the hypotheses tested are irrefutable, because they 'fail to forbid any observable state of affairs' (Lakatos, 1970, p. 100).

It is reasonable to infer that the inducement hypothesis implies that there should be a long-run relation-

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ship 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, so although there may be short-run deviations, there should be some equilibrating mechanism that brings 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. This also allows different formulations of the hypothesis to be compared. Then, if a cointegrating vector exists, an error correction model can be constructed to determine the long-run relationships, and the direction of causality can also be established. This paper adds policy variables, such as subsidised credit and tax concessions, which are also important determinants of biases in technical change. The effect of lobbying by interest groups of large and small farmers is also introduced by including farm size.

Section 2 critically examines variants of the hypothesis and discusses the difficulties involved in empirical tests. Section three describes the two-stage constant elasticity of substitution (CES) model and explains why it is still popular, despite the availability of flexible functional forms that are less restrictive. Section four describes the data and applies time series analysis to determine the properties of the variables, and to establish cointegration. Section five formalises the relationships by fitting an ECM, which separates the short-run effects (factor substitution) from the long-run equilibrium path, which depends on technological innovation. Finally, the ECM is tested to establish that the causality runs from the price ratios to the factor ratios.

2. Modelling induced innovation

Hicks (1932) introduced the elasticity of substitution and the idea of 'induced inventions', which endogenised the factor-saving bias of technical change at the level of the firm. However, the two concepts were not clearly separated, as Hicks noted in his Nobel lecture of 1973 (Hicks, 1977, p. 2). This oversight led to critiques of induced innovation such as Blaug (1963) and Salter (1960), which were later shown to depend largely on the definition of the isoquant (Hayami and Ruttan, 1985, p. 86). The induced innovation hypothesis was rehabilitated when Ahmad (1966) introduced the innovation possibility curve (IPC), which is the envelope curve of all the isoquants (representing different technologies) that may be developed, given the state of scientific knowledge.

The IPC (together with its counterpart, the meta-production function) form the basis of Hayami and Ruttan's application of the hypothesis to aggregate agricultural output in a long-run historical development context. They argue that growth in agricultural productivity is generated by technical change that facilitates the substitution of relatively abundant (hence cheap) factors for relatively scarce (hence expensive) factors in the economy (Hayami and Ruttan, 1985, p. 73). Their model is developed by exploiting the identity

$$Q/L \equiv (Q/A)/(A/L) \tag{1}$$

where Q is output, L is labour, and A is land. Land area per worker (A/L) is increased by mechanical technical change, which allows power to be substituted for labour. Similarly, biological advances, such as high-yielding, fertilizer-responsive seed varieties, raise the average product of land (Q/A), and may be referred to as biological/chemical technical change. Thus, technical changes are represented as movements around the IPC, and changes in factor ratios are induced, to a significant extent, 'by the long-term trends in relative factor prices' (Hayami and Ruttan, 1985, p. 181).

The agricultural histories of the United States, where labour was the relatively scarce factor, and Japan, where land was scarce, show that Japanese agricultural technology was relatively yield-increasing (land-saving relative to labour), whereas technical change in the USA was labour-saving relative to land. Biological and chemical technology dominated in Japan, while mechanical technical change was relatively more important in the USA. Hayami and Ruttan's three equation tests of the inducement hypothesis regressed the logarithms of the factor ratios (land/labour, fertilizer/land and machinery/labour) 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.

However, the test is entirely ad hoc, and the distinction between factor substitution and induced innovation is not clear. That the causality runs from price ratios to factor ratios is not tested, and nor is the assumption that long-run factor substitution depends upon technical change. These simple tests are not derived from a particular functional form of the production relationship, but the revised edition of Hayami and Ruttan (1985) includes two-stage tests based on the two-stage constant elasticity of substitution (CES) production function. The two-stage CES is estimated, with time-dependent factor augmentation coefficients, to produce estimates of the Allen elasticities of factor substitution and the biases of technical change.

Then, in the 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 et al. (1986), who also apply the model to the USA and Japan, Thirtle (1985), who considers US wheat production, and Karagiannis and Furtan (1990) who use Canadian data. Binswanger (1974, 1978) pioneered a similar path, applying the translog cost function, and using relative factor prices to explain the residual factor shares, net of factor substitution. The 'isotech' analysis of Nordhaus (1973), which has been extended by McCain (1977) and Wyatt (1986) is similar to the Hayami and Ruttan approach, but has not been taken on board by agricultural economists. These developments are more fully outlined by Thirtle and Ruttan (1987).

However, the 'traditional model', developed by Binswanger and many others, uses a time trend in estimating the biases of technical change. Clark and Youngblood (1992) argue that the time series properties of all the variables must be established prior to estimation, since the "parameter estimates for the traditional model are valid only if all independent variables are stationary, and the dependent variables are driven by a deterministic time trend". They correctly suggest that estimates of the biases of technical change should be based on measures such as research expenditures or publications. This is common practice in the duality-based cost and profit function approaches to technical change in agriculture, which are conveniently summarised by Evenson and Pray (1991) (pp. 185–194). 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.

The combination of duality and flexible functional forms employed in studies of this type increases the level of technical sophistication considerably, but is far more demanding in terms of data. The alternative approach, based on the two-stage CES function, is parsimonious in this respect, and the functional form itself gives rise to estimating equations that allow direct testing of the inducement hypothesis.

3. A direct test of the induced innovation hypothesis

The model developed by de Janvry et al. (1989) 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 (1991) uses the model straightforwardly to test the induced innovation hypothesis, and it is this approach that is followed here. The two-stage (CES) production function forms the starting-point

$$Q = [\gamma(Z_1)^{-\rho} + (1-\gamma)(Z_2)^{-\rho}]^{-1/\rho} \quad (2)$$

$$Z_1 = [\beta(L)^{-\rho_1} + (1-\beta)(E_m M)^{-\rho_1}]^{1/\rho_1} \quad (3)$$

$$Z_2 = [\alpha(A)^{-\rho_2} + (1-\alpha)(E_f F)^{-\rho_2}]^{-1/\rho_2} \quad (4)$$

where Q is aggregate agricultural output, Z_1 is the labouresue input group (comprising labour, L, and machinery, M), Z_2 is the landesque inputs (land, A, and fertilizer, F), E_m and E_f are efficiency parameters, and all the Greek letters represent the usual substitution and factor share parameters of the CES function. Rearranging the first-order conditions from the profit maximisation problem, and assuming equilibrium, so that marginal products are equal to factor prices, gives the estimating equations for Eqs. (3) and (4)

$$\ln\left[\frac{M}{L}\right] = \sigma_1 \ln\left[\frac{1-\beta}{\beta}\right] + [\sigma_1 - 1]\ln E_m - \sigma_1 \ln\left[\frac{P_m}{P_1}\right] \quad (5)$$

$$\ln\left[\frac{F}{A}\right] = \sigma_2 \ln\left[\frac{1-\alpha}{\alpha}\right] + [\sigma_2 - 1]\ln E_f - \sigma_2 \ln\left[\frac{P_f}{P_a}\right] \quad (6)$$

where the dependent variables are the natural logs of the machinery/labour and fertilizer/land factor ratios, and the last terms on the right are the equivalent factor price ratios. The direct partial elasticity of substitution of labour for machinery is σ_1, and that for fertilizer and land is σ_2 (Kawagoe et al., 1986). E_{m} and E_{f} are efficiency parameters, and α and B are the share parameters. Thus, the two-stage CES approach leads to a direct test of the inducement hypothesis, since factor ratios are regressed on constants, the price ratios and the efficiency parameters. Frisvold (1991) argues that 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, and if the lagged price ratios explain the factor ratios, then the inducement hypothesis is corroborated.

The treatment of the efficiency parameters remains somewhat arbitrary. de Janvry et al. (1989) and Frisvold (1991) assume that the efficiencies are functions of research activities, so

$$E_{\mathrm{f,m}} = E_{\mathrm{f,m}}[\theta, B, R, t] \qquad (7)$$

where $\boldsymbol{\theta}$ is a vector of shares of past public sector research budgets, B, allocated to land-saving technical change (E_{f}) and labour-saving technical change (E_{m}), $\boldsymbol{R}$ 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, $\boldsymbol{\theta}$, allocating the research budget between activities, itself depends on expected relative factor prices and explicit behavioural assumptions regarding the government research budget allocation, so that

$$\theta^* = (P^e_M/P^e_L,\ P^e_F/P^e_A,\ P^e_A/P^e_L,\ S,\ B,\ R,\ t) \qquad (8)$$

where the price ratios are all expectations. de Janvry et al. (1989) also include transaction costs to explain research allocations. As the transaction costs for labour (supervising, negotiating, information costs) increase with farm size (S), there will be an increasing bias in research towards labour 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 increases the bias towards land saving technology if small farmers' demands prevail. de Janvry et al. (1989) substitute Eqs. (7) and (8) into Eqs. (5) and (6) so the determinants of the optimal technical changes and factor ratios appear in their reduced form equations.

The same approach is followed here, but the model is adapted to the South African case. The factor ratios are assumed to be functions of the past R&D expenditures, that generated the technologies, extension expenditures that transmitted the results to the farmers, so diffusing the technology, and the education level of the farmers, which affects both their own creative and managerial abilities, and their skill in appraising and adapting exogenous technologies. Chemical and mechanical patent data is used to account for private sector research, and farm size is included as a cause of factor-saving biases. The policy distortions of the apartheid period are also allowed for by including dummy variables for tax concessions, negative real interest rates for farm credit, and the operation of the Pass Laws, which made farm labour artificially scarce.

4. Time series analysis of the data

The data are for the commercial agricultural sector of the Republic of South Africa, for the period 1947–1991. The main source is the *Abstract of Agricultural Statistics* (Republic of South Africa, 1993) supplemented by historical material and unpublished information from the Department of Agriculture.

The land measure is the total hectares of agricultural land and the land price is the rental value. Labour is all hired labour and the price is the wage for hired labour. Fertilizer is a constant price series for an aggregate of all fertilizer and the price is for the average fertilizer mix. The machinery series is the service flow (interest, depreciation and running costs) from the capital stock of machinery, implements, motor vehicles and tractors, taken from Thirtle et al. (1993). This avoids the problem of mixing stock and flow variables. The series for R&D, extension, farmer education and patents are from Thirtle and van Zyl (1994). The patents are counts of chemical and mechanical patents, pertaining to agriculture, registered by all countries in the US. They are used both to account for private R&D activity and to try and capture some of the international R&D spillovers. The model was customised to suit the special problems of South Africa by including dummy variables to capture the effect of taxation policy, interest rates and the control of labour by means of the Pass Laws. The dummy variable for tax policy covers the period up to 1982, when concessions such as agricultural buildings subsidises and rapid tax write-offs for machinery were withdrawn. The dummy for the negative real interest rate covers 1973–1982 and 1985–1987; the Pass Law dummy covers 1968–1985. Farm size is used in the model to represent transaction costs of land purchases and sales and of labour management, as in de Janvry et al. (1989).

Cointegration techniques are used to establish valid relationships between the variables. If the variables are cointegrated, then deviations from the long-run equilibrium path should be bounded. 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 and second, a linear combination must exist that is integrated of an order one less than the original variables; that is, if the variables are integrated of order one (denoted $I(1)$), the error term from the cointegrating regression should be stationary (or $I(0)$).

The time series properties of the variables, all in logarithms, are reported in Table 1. The standard Dickey–Fuller (Dickey and Fuller, 1981) (DF) test and the Sargan and Bhargava (1983) CRDW test are used to determine the order of integration. Column two shows that all the variables are first-order autoregressive (AR) processes, which indicates that the error terms in the DF tests are white noise, as required. This was confirmed using the Ljung and Box (1978) Q^* statistic for serial correlation, which is appropriate for small samples.

The next column shows that the DF test statistics indicate that all the variables are non-stationary in the levels, except one (all except the −3.81 for mechanical patents are greater than the critical value of −2.90). However, the DF test values for the first differences are all less than the critical value, indicating that all the variables are $I(1)$, except for mechanical patents, which appears to be $I(0)$. The CRDW tests in column four confirm the $I(1)$ results and suggest that mechanical patents is also borderline $I(1)$. The Dickey–Fuller ϕ_2 test (Dickey and Fuller, 1981) in the last column jointly tests for a unit root and a deterministic trend. The test statistics for all the variables except one are less than the critical value, indicating that there are no deterministic trends. The exception is again the mechanical patents series, which later proved to have no explanatory power. The weather variable is rainfall,

Table 1
Testing the variables for order of integration

Variable name and abbreviation	AR order	DF tests	CRDW tests	D-F ϕ_2 tests
Log ratio machinery/labour (M/L)	1	−2.40	0.07	3.07
Δ Log ratio machinery/labour		−5.66	1.75	
Log ratio price machinery/labour (P_M/P_L)	1	−1.36	0.18	1.37
Δ Log ratio price labour/machinery		−7.48	2.29	
Log ratio fertilizer/land (F/A)	1	−1.79	0.03	2.24
Δ Log ratio fertilizer/land		−5.94	1.85	
Log ratio price fertilizer/land (P_F/P_A)	1	−2.25	0.04	3.76
Δ Log ratio price fertilizer/land		−5.47	1.67	
Log ratio price land/labour (P_A/P_L)	1	−1.41	0.12	2.48
Δ Log ratio price land/labour		−6.76	2.11	
Log of real R&D expenditures (RD)	1	−1.07	0.04	1.63
Δ Log of real R&D expenditures		−5.32	1.42	
Log of real extension expenditures (EXT)	1	−2.07	0.12	4.28
Δ Log of real extension expenditures		−7.57	1.97	
Log of farm size (SIZE)	1	−0.37	0.02	1.68
Δ Log of farm size		−5.15	1.65	
Log of chemical patents (CHP)	1	−2.68	0.11	6.45
Δ Log of chemical patents		−8.46	2.54	
Log of mechanical patents (MEP)	1	−3.81	0.66	13.50
Δ Log of mechanical patents		−6.26	1.95	
Critical values		−2.90	0.67	6.73

which is not included in the table, as it was $I(0)$, as expected. This means that it can explain short-run deviations, but will play no role in the long-run relationships. Thus, the induced innovation hypothesis clears the first hurdle, in that all but one of the variables have the same time series properties and may cointegrate.

These results also indicate that for South Africa, the induced innovation hypothesis should be formulated in the original manner (Hayami and Ruttan, 1985), with factor ratios being a function of factor price ratios. This is contrary to Tiffin and Dawson (1995), who suggest modifying the hypothesis to make *changes* in factor ratios a function of factor price ratios. This result comes from their finding that for the US data from Hayami and Ruttan (1985), the factor ratios are $I(1)$, while the price ratios are $I(2)$.

The key relationship between factor ratios and factor prices are considered next, before incorporating the range of other variables. Fig. 1 shows that after 1950, there is a remarkably close relationship between the ratio of machinery to labour, and the price of labour relative to the price of machinery. In the first few years, after the end of the war, the arable area was being extended, which led to increased use of labour (especially since harvesting was not mechanised) as well as machinery. Also, supplies of agricultural machinery and equipment were limited, and animal power was still important. Then, for three decades, the price of machinery falls relative to the price of labour and the machinery labour ratio rises, as the induced innovation hypothesis predicts. Indeed, regressing the factor ratio on the price ratio from 1954 onwards gives an R^2 of 0.94, indicating that the farmers respond strongly to prices. Then, after the 1980 collapse of the gold price, the Rand was drastically devalued (Thirtle et al., 1993) and the favourable credit and tax policies were largely withdrawn. These events had a combined effect of making a domestic input like labour far cheaper relative to capital, and led to a dramatic reversal of the historical trend, with labour increasing as it was substituted for expensive capital. It is because the policy changes are so dramatic, giving price reversals of this magnitude, combined with the quality of the data, which is as good as any for the western countries, that makes South Africa ideal for testing the inducement hypothesis. The machinery/labour ratio is also shown net of long-run factor substitution, following Eq. (5) and using the elasticity estimate from the ECM developed in Section 5.

The induced innovation hypothesis appears to be well supported by these data, in that there is a lot of change in the factor ratio not explained by factor substitution, which must be true for the hypothesis

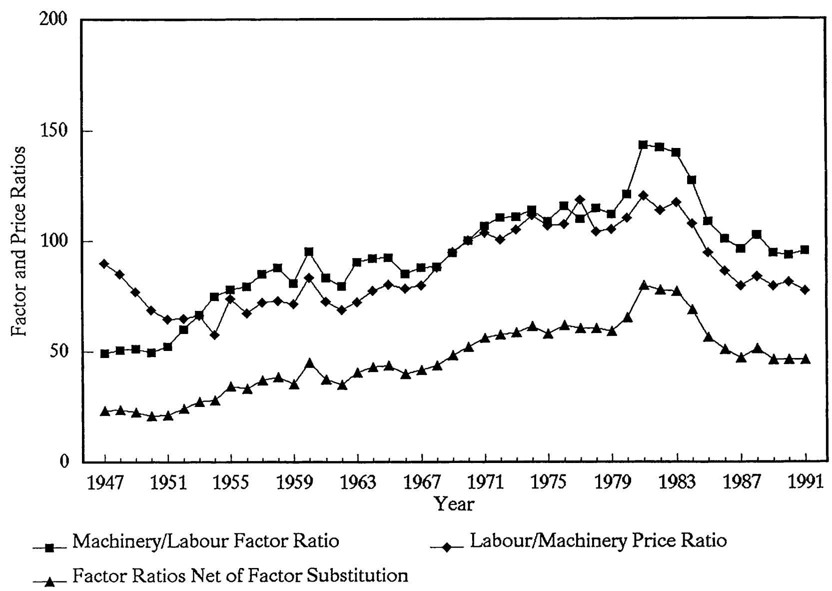

Fig. 1. Price and factor ratios, Gross and net of factor substitution.

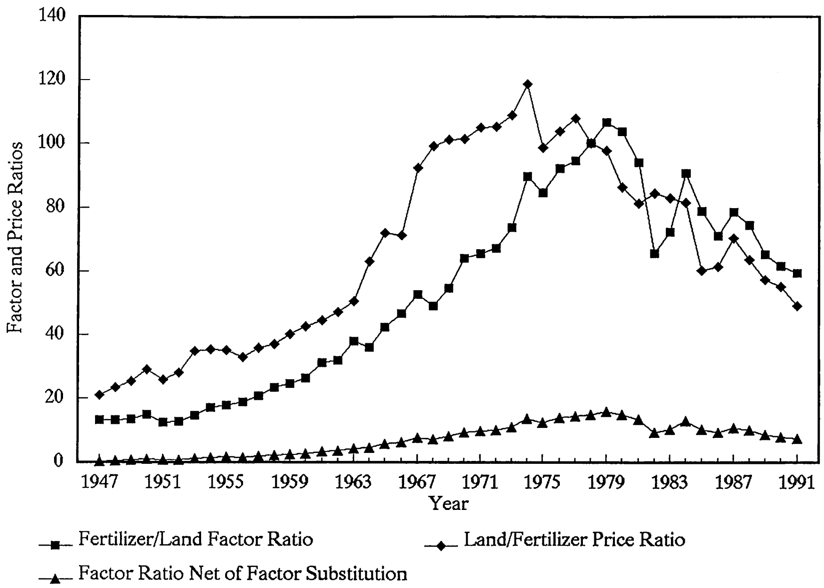

Fig. 2. Price and factor ratios, Gross and net of factor substitution.

to hold. However, the claim that this residual is explained by induced innovation still rests on the argument of Hayami and Ruttan (1985) that long-run factor substitution is only made possible by innovation. Indeed, there are other factors, such as policy changes, which also contribute to the explanation of the residual. Also, the lack of lags between price changes and quantity adjustments is surprising, although it is partly caused by the machinery series being the service flow rather than the capital stock.

Fig. 2 shows that the fertilizer/land factor ratios and price ratios are also highly correlated. The fit is not as close as for labour and machinery, but for the full period the R^2 was 0.74. There was some growth in land during the period 1947 to 1959, when the cultivated area was being expanded. The fertilizer/land factor ratio shows two distinct trends, growing rapidly at 7.38% per annum from 1947–1979, and then falling at a rate of −3.98% per year. The decrease in fertilizer use in 1975 can be attributed to the rise in the relative price of fertilizer that resulted from the OPEC oil crisis. Then, the fertilizer/land ratio rose again as land rents increased rapidly in the late 1970s (explained in Van Schalkwyk et al., 1994). Finally, following the collapse of the gold price and devaluation of the Rand at the beginning of the 1980s, the land/fertilizer price ratio rose while the fertilizer/land ratio fell in 1981 and 1982. This was the results of two years of severe drought, which made heavy fertilizer use pointless. At the same time, agricultural subsidies and tax concessions ended, and land rents fell relative to the price of fertilizer from 1984, with the factor ratio following the price ratio closely, with no lag. Groenewald (1986) and Van Schalkwyk and Groenewald (1992) suggest that many farmers in South Africa, particularly grain farmers, over-fertilized, so the reduction contributed to the increase in productivity growth reported by Thirtle et al. (1993). The fertilizer/land ratio net of long-run factor substitution shows that the residual left to be explained by induced innovation and other factors is initially smaller than in the case of machinery, but is still substantial from the late 1960s onwards.

5. Cointegration and the error correction model

Having established that the variables are integrated of the same order, the next stage is to test for cointegrating vectors which imply that non-spurious long-run relationships exist between the variables. Three tests were used to test for cointegration between combinations of variables, namely the DF test, the CRDW test and the Johansen procedure. The maximum likelihood approach of Johansen (1988) and Johansen and Juselius (1990) allows estimation of all cointegrating relationships and tests for the number of cointegrating vectors and the direction of causality. The procedure begins by defining a VAR of a set of variables X,

$$X_t = \pi_1 X_{t-1} + \ldots + \pi_k X_{t-k} + e_t \qquad t = 1, \ldots, T \tag{9}$$

where, if there are n variables, this becomes an n-dimensional kth order vector autoregression model with Gaussian errors. X_t is a vector of the n variables and k is large enough to make the error term white noise. The VAR model can be reparameterized in error correction form (Cuthbertson et al., 1992) as

$$\Delta x_t = \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \prod x_{t-k} + e_t, \qquad t = 1, \ldots, T$$

$$\text{where } \Gamma = [(\mathbf{I} + \pi_1), (\mathbf{I} + \pi_1 + \pi_2), \ldots, (\mathbf{I} + \pi_1 + \ldots + \pi_k)] \tag{10}$$

$$\prod = \mathbf{I} - \pi_1 - \pi_2 - \ldots - \pi_k$$

where $\mathbf{I}$ is the identity matrix, so this leads to the ECM that is developed next.

The Johansen testing procedure is a multivariate likelihood ratio test for an autoregressive process with independent Gaussian errors. The procedure involves the identification of rank of the matrix $\mathbf{\Pi}$. Three possible cases can arise concerning the rank of $\mathbf{\Pi}$. First, if $\mathbf{\Pi}$ has full rank, x_t is stationary. Secondly, if $\mathbf{\Pi}$ has zero rank, implying that $\mathbf{\Pi}$ is the zero matrix, x_t is non-stationary and not cointegrated and Eq. (9) is simply a VAR in first differences. Thirdly, $\mathbf{\Pi}$ can have reduced rank $0<r<p$ implying cointegration. If $\mathbf{\Pi}$ is of rank ($0<r<p$), then $\mathbf{\Pi}$ can be expressed as $\mathbf{\Pi}=\alpha\beta'$ where α and β are mxr full column rank matrices. β may be interpreted as the mxr matrix of cointegrating vectors representing long-run relationships, and α as the mxr matrix of loading weights showing how the cointegrating vectors are loaded into each equation of the system. Causality within the system is determined below, by testing zero restrictions on the α-matrix using Wald tests (Hall and Milne, 1994; Caporale and Pitts, 1995).

The results of the three tests are reported in Table 2, which does not include the parameter values, because the ECM that follows is a preferable representation, since it models the dynamics. Comparing the test statistics with the critical values shows that the machinery labour equation cointegrates according to all four tests. Mechanical patents, which appeared to be $I(0)$ was not significant and was dropped at this stage, as were extension expenditures, and the dummy variable representing the Pass laws. The fertilizer/land equation narrowly fails to cointegrate according to the DF test, but passes the CRDW test and the two Johansen tests find three cointegrating vectors, which indicates that there may well be feedback effects. R&D and extension were collinear, and in this case extension gave more significant results, which is reasonable, since the extension service recommended (excessive) fertilizer application rates over most of the period. Farm size was non-significant in the cointegrating fertilizer–land equation, suggesting that there is no difference between large and small farmer demands for fertilizer technology. The poor DW statistic for the second equation (CRDW in the table) is a reflection of the omitted dynamics rather than a long-run misspecification. Thus, there are non-spurious long-run relationships between the variables, and the ECM is a valid representation (Engle and Granger,

Table 2
Cointegration tests

Preferred equations factor and price ratios in logs	DF test	CRDW test	Johansen model	
			eigenvalue test	trace test
M/L C P_M/P_L P_A/P_L RD SIZE TAXD INTD	−5.5 (−5.1)	1.5 (0.49)	(1) 40.15 (34.4)	93.25 (76.07)
F/A C P_F/P_A P_A/P_L EXT CHP Rainfall	−4.5 (−5.1)	1.2 (0.49)	(1) 68.57 (34.4)	137.42 (76.1)
			(2) 31.05 (28.1)	68.89 (53.1)
			(3) 24.02 (22.0)	37.85 (34.9)

The critical values are in parentheses.

1987), so the inducement hypothesis clears this second hurdle.

Using the variables reported in Table 2, the ECMs for explaining the machinery to labour and the fertilizer to land factor ratios are

$$\begin{aligned}\Delta\ln(M/L)_t &= \phi_0 + \sum_{t=t-1}^{t-i} \phi_1\Delta\ln(M/L)_t + \sum_{t=t}^{t-j} \phi_2\Delta\ln(P_M/P_L)_t \\ &+ \sum_{t=1}^{t-k} \phi_3\Delta\ln(P_A/P_L)_t + \sum_{t=1}^{t-l} \phi_4\Delta\ln RD_t \\ &+ \sum_{t=1}^{t-o} \phi_5\Delta\ln\mathrm{SIZE}_t + \phi_6\mathrm{TAXD} + \phi_7\mathrm{INTD} \\ &+ \lambda[\ln(M/L) - \alpha_1\ln(P_M/P_L) - \alpha_2\ln(P_A/P_L) \\ &- \alpha_3\ln\mathrm{RD} - \alpha_4\ln\mathrm{SIZE}]_{t-1} \qquad (11)\end{aligned}$$

$$\begin{aligned}\ln(F/A)_t &= \delta_0 + \sum_{t=t-1}^{t-p} \delta_1\Delta\ln(F/A)_t + \sum_{t=t}^{t-q} \delta_2\Delta\ln(P_F/P_A)_t \\ &+ \sum_{t=1}^{t-r} \delta_3\Delta\ln(P_A/P_L) + \sum_{t=1}^{t-u} \delta_4\Delta\ln\mathrm{CHP}_t \\ &+ \sum_{t=1}^{t-v} \delta_5\ln\mathrm{EXT}_t + \delta_6\mathrm{RAIN} + \lambda[\ln(F/A) \\ &- \beta_1\ln(P_F/P_A) - \beta_2\ln(P_A/P_L) - \beta_3\ln\mathrm{CHP} \\ &- \beta_4\ln\mathrm{EXT}]_{t-1} \qquad (12)\end{aligned}$$

M/L=machinery/labour factor ratio, P_M/P_L=machinery/labour price ratio, P_A/P_L=land to labour price ratio, P_F/P_A=fertilizer to land price ratio, RD=public research expenditure, EXT=extension expenditure, SIZE=average farm size, TAXD=tax dummy, INTD=interest rate dummy, CHP=chemical patents.

The difference terms and the dummy variables are $I(0)$ and cover the short-run situation, whereas the long-run relationship is captured by the $I(1)$ level terms, within the square brackets. The seemingly unrelated estimation (SURE) procedure was used in order to gain efficiency in estimation, since the errors of the two equations were correlated. The results of the ECM reported in Table 3 were chosen on the criteria of goodness-of-fit (variance dominance), data coherence, parameter parsimony and consistency with theory (Hendry and Richard, 1982). For the machinery/labour equation, 66% of the variance is explained, and there is no evidence of serial correlation. The short-run own-price coefficient may be interpreted as the direct elasticity of substitution and the value of -0.37 is significantly different from unity, suggesting that the Cobb Douglas would not be suitable for these data. The other price ratios were not significant in the short run. Since the variables for tax concessions and negative interest rates are dummies, they are included in the levels and prove to be significant and positive. Thus, there is evidence that policy-induced innovation matters.

The adjusted R^2 for the fertilizer/land equation of 0.51 is still reasonably high for an ECM, and the serial correlation in the levels regression, reported in Table 2, has been corrected by the dynamic specification. The short-run coefficient on the fertilizer/land price is the direct partial elasticity of substitution, which has a value of -0.39. By comparison, the US results from Frisvold (1991) are -0.27 for fertilizer/land and -0.45 for machinery/labour, which is more counter-intuitive. The land/labour price ratio coefficient is positive and significant in the short run, suggesting that labour and fertilizer are complements. Rainfall was not reported in Table 1 as it conforms to the expectation of being $I(0)$. It is included in the short run and is only significant at low confidence levels.

The long-run results, derived from the adjustment terms, are entirely consistent with the induced innovation hypothesis. The negative coefficient on the long-run (lagged) own-price ratio is taken by Frisvold (1991) to indicate that a decrease in the machinery/labour price ratio generates labour-saving technological change. However, the lagged price ratio terms could be taken to indicate long-run factor substitution, rather than induced innovation. This issue is taken up below. The error correction term for the machinery/labour equation is -0.70, which indicates that when the system is not in equilibrium, there is 70% correction towards the long-run equilibrium level in the current period. The positive sign on the lagged land/labour price coefficient is in agreement with the prediction by Frisvold (1991). Public R&D expenditures are positive and significant, and since induced innovation requires R&D, the significance of R&D in the model helps corroborate the inducement hypothesis. Lastly, the positive coefficient for farm size is in

Table 3
Unrestricted ECM results: significant variables only

Variable	Coefficient	(M/L) ratio	Coefficient	(F/A) ratio
Short run				
Constant	Φ_0	3.21 (3.5)	δ_0	−5.11 (−4.5)
$\Delta(P_M/P_L)_t$	Φ_2	−0.37 (−3.4)		
$\Delta(P_F/P_A)_t$			δ_2	−0.39 (−2.7)
$\Delta(P_A/P_L)_t$	Φ_3		δ_3	0.17 (1.9)
Tax dummy	Φ_6	0.16 (2.6)		
Interest dummy	Φ_7	0.09 (2.5)		
Rainfall			δ^6	0.12 (1.3)
Long run				
$(M/L)_{t-1}$	λ	−0.70 (−5.6)		
$(F/A)_{t-1}$			λ	−0.41 (−4.7)
$(P_M/P_L)_{t-1}$	$\lambda\alpha_1$	−0.33 (−2.0)		
$(P_F/P_A)_{t-1}$			$\lambda\beta_1$	−0.36 (−3.3)
$(P_A/P_L)_{t-1}$	$\lambda\alpha_2$	0.12 (2.8)	$\lambda\beta_2$	0.14 (1.9)
$(\mathrm{RD})_{t-1}$	$\lambda\alpha_3$	0.09 (1.4)		
$(\mathrm{CHP})_{t-1}$			$\lambda\beta_3$	0.06 (1.6)
$(\mathrm{EXT})_{t-1}$			$\lambda\beta_4$	0.20 (1.7)
$(\mathrm{SIZE})_{t-1}$	$\lambda\alpha_4$	0.23 (1.5)		
Adjusted R^2		0.66		0.51
DW		2.3		1.9

The critical t values are 1.3 for 90% confidence, 1.68 for 95% and 2.02 for 97.5%, where a one-tailed test is appropriate.

agreement with the suggestion of de Janvry et al. (1989) that large farmer demands predominate, increasing the machinery-using bias.

In the long run, the error correction term for the fertilizer to land equation is −0.41, where the negative sign shows that the direction of correction is towards equilibrium. This indicates slower adjustment towards the long-run equilibrium level in the current period, with full adjustment taking about 2 1/2 years. The adjustment for machinery covers a multitude of items (including public R&D on irrigation equipment and specialised harvesting machinery), but capital stocks should take time to adjust, whereas the results show machinery adjusts more quickly than fertilizer. The negative coefficient on the long-run own-price variable can be taken to mean that a decrease in the fertilizer/land price ratio generates land-saving technological change, in the manner predicted by the induced innovation hypothesis. Chemical patents, representing international technological spillovers, are positive and significant, and so is the effect of extension efforts. The public R&D expenditures fail to show any effect due to collinearity with these two variables, but this weakness suggests that the South African research system has tended to adapt foreign technology rather than developing its own basic genetic material. Thus, technology-related variables are again significant in explaining the factor ratio, which corroborates the inducement hypothesis.

Long-run elasticities can be calculated from the results in Table 3 simply by dividing the long-run coefficients by the appropriate adjustment elasticity, λ. Whereas the short-run elasticities of substitution in Table 3 may be viewed as movements around isoquants, the long-run equivalents calculated here would correspond to movements around innovation possibility surfaces, which encompass all the techniques that can be developed, given the state of scientific knowledge. If this meaning is attributed to the disequilibria in the system, it implies that long-run equilibrium is attained only after the innovations that have been induced by the changes in prices are adopted. These long-run elasticities are reported in Table 4.

The large adjustment coefficient for the machinery/labour equation means that the elasticity of substitution for the long run is −0.47, which is only 30% greater than for the short run. The long-run elasticity

Table 4
Long-run elasticities

Variable	P_M/P_L	P_F/P_A	P_A/P_L	RD	EXT	CHP	SIZE
M/L equation	−0.47	–	0.17	0.13	–	–	0.33
F/A equation	–	−0.85	0.34	–	0.49	0.15	–

of the land/labour price ratio is 0.17, and the elasticity of R&D suggests that in the long-run, a 1% increase in R&D will increase the machinery/labour ratio by 0.13%. Similarly, a 1% increase in farm size will eventually increase the machinery/labour ratio by 0.33%.

The fertilizer/land results are more satisfactory, due to the lower adjustment coefficient of 0.41. This gives a long-run direct partial elasticity of substitution, around the innovation possibility curve, of −0.85. The corresponding elasticity for the land/labour price ratio is 0.34 and in the long run, a 1% increase in extension and chemical patents increases the factor ratio by 0.49 and 0.15%, respectively.

Thus, if the scheme of Frisvold (1991) for directly testing the inducement hypothesis is followed, the ECM extends the tests substantially, and the results entirely corroborate the hypothesis. The snag is that factor substitution can be costly, and take time to accomplish. Thus, all the own-price induced changes in the factor ratios could be attributed to long-run factor substitution. It is the long-run elasticities from Table 4 that are used to entirely remove factor substitution from the factor ratios in Figs. 1 and 2. The figures show the factor ratios net of all price-driven factor substitution effects. When substitution has been accounted for, there is still a large proportion of the change in factor ratios to be explained, and in both cases the technology variables play an important explanatory role. This is a stricter test of the inducement hypothesis, but it still passes, although it is clear that policy variables matter also.

Thus, the inducement hypothesis is corroborated by the significance of the variables in the ECM. The final stage is to apply causality tests to the ECM to ensure that the price ratios and technology-related variables are Granger-prior to the factor ratios. Causality tests within the ECM framework can be conducted by testing the loading matrix of the Johansen model. If the α matrix has a complete column of zeroes, then no cointegrating vector will appear in a particular block of the model, indicating that there is no causal relationship. The restrictions are tested by direct Wald tests on the loading parameters (Hall and Milne, 1994).

The loading matrix $\varkappa^2$ tests shown in the first row of Table 5 confirm the causality from the price ratios and other variables to the factor ratios. This is shown by the significant value of 30.64 for the VAR in which the machinery/labour ratios is the dependent variable. The loading matrix $\varkappa^2$ tests shown in the rest of the first row indicate that the disequilibrium in the machinery/labour factor ratio does not feedback to the price ratios, R&D or farm size. This suggests that the price ratios, R&D and land size can be treated as weakly exogenous, which is consistent with the single-equation model used here.

The directions of causality in the fertilizer to land system is more complex, as was already suggested by the existence of three cointegrating vectors. The results from the loading matrix indicate that there is causality from the price ratios and other variables to the factor ratio, but there is also feedback from the disequilibrium in the fertilizer/land ratio, to the ferti-

Table 5
Causality tests: Wald test results of zero restrictions on the loading matrix, α

Machinery/labour factor ratio	Machinery/labour price ratio	Land/labour price ratio	R&D expenditure	Farm size
30.64**	0.11	1.44	1.83	0.37
Fertilizer/land factor ratio	Fertilizer/land price ratio	Land/labour price ratio	Extension expenditure	Chemical patents
12.19**	13.82**	2.69	15.23**	0.38

lizer/land price ratio and extension expenditures. This suggests that the fertilizer/land price ratio and extension are not weakly exogenous, hence the single equation model may produce biased results. However, in both cases, there clearly is causality from the price ratios, and technology variables to the factor ratios, so the inducement hypothesis passes the last test.

6. Conclusion

This paper tests the induced innovation hypothesis of Hayami and Ruttan (1985) using data for South African agriculture. The two-stage CES production function leads to estimating equations that directly test the inducement hypothesis by making factor ratios a function of factor price ratios. The model also incorporates the variables that generate the new technologies (public R&D and extension and private patents), and other variables that affect the factor saving biases, such as farm size, and policy variables like negative interest rates and tax concessions.

The use of cointegration methods allows the inducement hypothesis to be subjected to a series of tests. First, since the variables are non-stationary, the time series properties of the variables are determined to ensure that cointegration is possible. Then, cointegration is established, and an error correct model constructed. Removing factor substitution shows that there are still considerable residual changes in factor ratios to be explained. The ECM shows that lagged price ratios and technology variables (public R&D and extension and private patents) are significant in explaining changes in factor ratios, and so are policy variables (subsidised credit and tax concessions). Finally, causality tests show that the price ratios and technology variables are causally prior to the factor ratios.

All of these results corroborate price-induced technical change, but also suggest that the model needs to be expanded to include farm size, research and extension expenditures and policy variables, which are also important determinants of the observed technological bias. This suggests that although relative input prices affect the biases of technical change, the relevant technology and policy variables also matter.

In the new South Africa, more attention should be focused on the technological needs of small-scale farmers, as the combination of the lobbying power of the large-scale commercial farmers and the policies followed by the apartheid regime have influenced the allocation of R&D expenditures between labour and land-saving technical change. This distorted the technological bias towards labour-saving technical change, which is hardly appropriate for a labour-surplus economy, in which the small farmers in the ex-homelands face chronic land scarcity.

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