REVISTA DE ECONOMÍA Y ESTADÍSTICA
INSTITUTO DE ECONOMÍA Y FINANZAS
FACULTAD DE CIENCIAS ECONÓMICAS
UNIVERSIDAD NACIONAL DE CÓRDOBA
ISSN 0034-8066

ARTÍCULOS

Factor Productivity in the Argentinean Agriculture

Alberto Herrou Aragón

Revista de Economía y Estadística, Cuarta Época, Vol. 45, No. 1 (2007), pp. 7-29.

http://revistas.unc.edu.ar/index.php/REyE/article/view/3834

Revista de Economía y Estadística - Vol. XLV - N° 1 - (2007) - pp. 7-29
Instituto de Economía y Finanzas - Facultad de Ciencias Económicas
Universidad Nacional de Córdoba - Argentina

Factor Productivity in the Argentinean Agriculture

Alberto Herrou Aragón
albertoherrou@hotmail.com

Resumen:

El objetivo de este trabajo es investigar la endogeneidad de la productividad factorial en la agricultura argentina. En un modelo de endogeneidad de la implementación de nuevas técnicas de producción, se encuentra que ésta dependería de los precios relativos del sector agropecuario y de las dotaciones de capital y de tierra relativas a la de mano de obra. Se encuentra que, con los datos disponibles para los años 1939-1984, las estimaciones no rechazan la hipótesis de endogeneidad del cambio tecnológico como función de las variables mencionadas. Además, se encuentra que el déficit fiscal afecta negativamente el cambio tecnológico.

Palabras claves: Productividad total de los factores, tecnología disponible e implementada, vectores autoregresivos cointegrados, factor de Barttlet, eigenvalues.

Clasificación JEL: C4, F1, O3, Q1

Abstract:

This paper is aimed at investigating the endogeneity of the total factor productivity in the Argentinean agriculture. In a simple model of endogenous technological change, the implementation of new techniques of production would depend upon sectoral relative prices, and upon the overall endowment of capital and land relative to labor. It is found that the estimations using data covering the years 1939-1984 do not reject the hypothesis of endogeneity of technological change as a function of the aforementioned variables. It is also found that the fiscal deficit negatively affects the implementation of new techniques.

Key words: Total factor productivity, available and implemented technologies, cointegrated vector autoregressions, Barttlet correction, eigenvalues.

JEL Classification: C4, F1, O3, Q1

I. Factor Productivity in the Argentinean Agriculture

Changes in total factor productivity (TFP), or changes in the total output-production factors index ratio, constitute one of the most important sources of economic growth. Sources of agricultural growth in the Argentinean economy are shown in Table 1 for the whole sample and for sub-samples. The stock of agricultural capital at 1960 prices includes machinery and equipment, land improvements, and livestock. Land is the total number of hectares devoted to cultivation of cereals, oilseeds, fruits and vegetables, industrial crops, and forage. In order to calculate an aggregate input index, the coefficients of the agricultural production function estimated by L. Reca and J. Verstraeten (1977) are used[1].

Table 1
Sources of Agricultural Output Growth – Years 1940-1984
(in %)

Output growth and sources	1940 - 1984	1940 - 1949	1950 - 1959	1960 - 1969	1970 - 1979	1980 - 1984
Output growth	1.35	0.39	1.37	1.97	2.12	0.43
Input contribution:	0.42	-0.47	0.93	0.97	0.34	0.27
Land	0.32	-0.53	0.69	0.68	0.16	0.86
Labor	-0.10	0.28	-0.20	-0.11	-0.22	-0.42
Capital	0.20	-0.22	0.44	0.40	0.40	-.0.18
Residual (TFP)	0.93	0.86	0.44	1.00	1.78	0.16

Notes: The figures are annually compounded rates of growth. The weights used to aggregate inputs are: (0.437), (0.276), and (0.287) for land, labor and capital, respectively.
Sources: Mundlak, Y., D., Cavallo, and R. Domenech (1989), and L. Reca and J. Verstraeten (1977).

[1] The coefficient of labor in the value added of the agricultural sector (cereals and oilseeds, livestock and fisheries) is 0.28 according to the estimates of the 1973 input-output matrix data, a value that is very close to the estimate used in this paper.

The figures show that for the years 1940-1984, the sectoral TFP has been the most important source of growth of agricultural output, accounting for about 70 percent of the rate of growth[2]. The sub samples of 1950-1959 and 1980-1984 are the exceptions as the contribution of total inputs to the sectoral rate of growth accounted for about 70 and 63 percent of total growth of agricultural output, respectively. V. J. Elías (1989) finds the same pattern of changes in the total factor productivity of the agricultural sector through decades in the Argentinean economy although changes in the TFP have a lesser role in accounting for total agricultural output changes compared to the estimates of this paper. This could be the result of differences in the data used in the studies[3].

If changes in the agricultural TFP were due to policy interventions, then by understanding how they affect the TFP, the agricultural growth process could also be understood and the long-run prospects of sectoral growth changed accordingly. As well documented by C. Diaz-Alejandro (1970) and J. Berlinski (2003), import substitution policies started being implemented in Argentina in the early 30s in response to external shocks. Import permits, increased import tariffs, and foreign exchange controls were the main policy instruments used to this effect. These policies were exacerbated in the mid-40s with the introduction of multiple exchange rates benefiting imports of intermediate goods and import prohibition of almost all imports competing with local production. During the crises of 1952, imports of capital goods were also banned. In 1958, all import prohibitions were eliminated and replaced with import tariffs with a maximum of 300 percent. In the mid-60s, export incentives were introduced to promote exports of manufactured goods thus reversing to some extent the anti-trade bias of import-substitution policies.

Freer trade policies were implemented in the late 70s but they were short-lived as a result of macroeconomic imbalances. These policies were reverted in the 80s by introducing import quantitative restrictions. According to theses new policies, imports required the approval of a committee formed by public officials and representatives of the import-competing private sector.

[2] As the TFP has been calculated as the residual of the agricultural production function estimated by Reca and Verstraeten (1977), they may also include changes in allocation of resources in addition to technological changes. This might have been an important source of changes in the TFP during the years of the second world war.

[3] For a description of the data used in this paper, see Annex.

These import-substitution policies have certainly negatively affected the economic incentives to exportable agriculture. In order to assess the extent to which import-substitution policies have harmed the agricultural sector, Mundlak, Cavallo and Domenech (1989) estimate a model in which sectoral technological change (the coefficients of a Cobb-Douglas production function and its intercept) is a function of sectoral relative prices, and of state variables such as the overall capital-labor ratio. They also include the lagged dependent variable in the agricultural technical change equation among the explanatory variables to deal perhaps with autocorrelation of residuals (this variable is omitted in the nonagricultural equation). Their estimates support the theoretical model. However, the authors only report the Durbin-Watson statistic to test for autocorrelation of residuals and, as is well known, this test statistic is biased towards rejection of autocorrelation of residuals in presence of lagged endogenous variables. Any unremoved residual autocorrelation will thus yield inconsistent estimates of the parameters.

This paper is thus aimed at finding some of the determinants of the agricultural TFP in the Argentinean agricultural sector during the years 1940-84 for which there is available data utilizing the cointegration analysis of vector autorregressions as all the variables are shown to be non stationary. In Section II, a theoretical framework of endogenous technical change (or changes in the TFP) is presented. The results of the estimations are in Section III and the concluding remarks are in Section IV.

II. An Endogenous Technical Change Framework

According to Mundlak´s technique choice framework (2000), new technologies might be available to firms but the costs of implementing them might be greater than the benefits. Thus, this approach emphasizes the difference between available and implemented technology in which the available technology is exogenously given but the rate of implementation of these techniques depends upon economic incentives and resource constraints.

Within this endogenous technology framework, if new available techniques are capital-intensive, then these techniques are going to be implemented by firms if the relative price of capital compared to other factors of production is low enough to make them profitable to acquire. Otherwise, firms would keep using traditional techniques that are less intensive in the use of capital and the new ones would not be adopted. Coexistence of traditional and new techniques is also feasible at certain threshold relative prices. The rate of adoption of production techniques by firms within the envelope of the available technology set would thus be a matter of economic choice and this would depend upon economic incentives that they face.

Economic incentives affect the relative profitability of available techniques as changes in the relative price of the agricultural sector affect the relative costs of implementing new techniques. C. Rodriguez (1982) shows that in a model with three goods (one of which is non-traded) and three factors of production, an increase in the relative price of the exportable commodity would increase the relative price of land compared to that of capital. On the other hand, the effect of the increase in the relative price of the exportable good over the relative price of labor is ambiguous. If an increase in relative prices causes a reduction in the relative price of capital, then new labor-saving techniques would be adopted by farmers.

The aforementioned analysis has implications for the policy making point of view. This analysis would predict that trade liberalization policies would tend to increase the rate of adoption of modern techniques of production as long as they cause a decline in the relative price of capital if these techniques are capital-intensive. This is illustrated in Fig.1 with an example of a simple model with two goods (import substitution and agricultural exportable goods) and two factors of production (capital and land). The four right angles represent unit value isoquants for the two sectors. They represent the combinations of capital and land that are required to produce a dollar´s worth of output. The fixed coefficient technology is immaterial for the analysis. Two different unit value isoquants are drawn for the agricultural sector representing two different techniques. The traditional technique (Ag0) is more land intensive than the modern technique (Ag1) at the same relative factor prices.

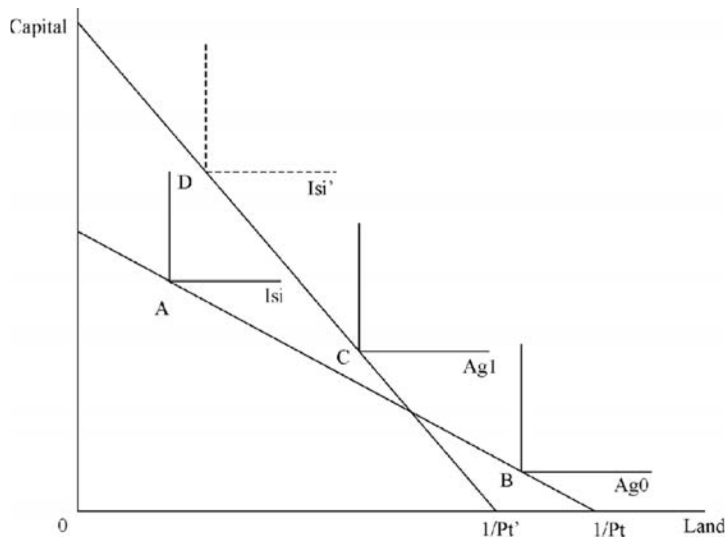

Fig. 1: Isoquants and implementation of capital-intensive techniques

Also in the figure, two unit isocost lines are drawn for two different relative factor prices. The iscost line intersecting the *x*-axis at point 1/Pt goes through the corners of the isocuant of the traditional agriculture (Ag0) and that of the import substitution sector (ISI) unit value isoquants, A and B. The slope of this unit value isocost is the relative price of land compared to that of capital under protection of the import substitution activity. This isocost line is thus compatible with production of the import substitution and traditional agricultural goods. At these factor prices, the modern technique of the agricultural sector is not going to be implemented because the isocost line falls below the unit value isoquant and, consequently, the unit cost exceeds the unit value of the output.

If the import tariff is removed or reduced, the relative price of land increases and the new isocost line intercepts the *x*-axis at 1/Pt'. With the reduction in the relative price of the importable good, its unit value isoquant shift upwardly to the right (ISI') as more capital and land are required to produce a dollar's worth of output. This unit cost line goes through the corners of the ISI' and Ag1 isoquants, C and D, and falls below the unit value isoquant Ag0. As a result, the modern technique of production of the agricultural sector is going to be implemented and the traditional technique is going to be discarded. There is of course a threshold relative factor price at which the two techniques in the agricultural sector are going to coexist.

Resource constraints also affect the implementation of new techniques. The main constraint is the level of the overall capital stock when new available techniques are more intensive in the use of capital as they appear to be[4]. Herrou-Aragón (2006) shows the conditions under which an increase in the capital-labor and capital-land ratios would result in a reduction in the relative price of capital in a general equilibrium model with three goods (one of which is non traded) and three factors of production (capital, labor, and land). These conditions are: (a) production of non traded goods is intensive in the use of labor; and (b) production of goods competing with imports is more labor intensive than that of production of exportable goods. If production of non traded goods is labor-intensive, then increases in the stock of capital and land would result in an excess demand for these goods at constant relative prices that, in turn, would require of an increase in the relative price of non-traded goods to clear the market. By the zero-profit condition, this increase in the relative price of non traded goods would increase the price of capital compared to that of labor. It is also shown in the aforementioned paper that under condition (b), an increase in the capital-labor and land-labor ratios would result in a reduction in the price of capital compared to that of land.

The available technology set is hard to measure as it is embodied in knowledge and, thus, in human capital. Schooling and expenditure in research and development can be measures of the available technology as they represent investment in human capital. Quality of schooling and profitability of research and development are issues that are hard to deal with actual data. Alternatively, as the human capital factor is a complement of the other factors of production, these factors are going to be positively related with knowledge.

The above discussion regarding the variables that could help to explain endogenous technological change in the agricultural sector (*A*) (TFP_A) can be summarized in a function such as the following,

4 For some empirical evidence about factor intensity of new techniques, see Mundlak (2000), chapter 6.

$$TFP_A = f\left(\frac{p_a}{p_m}, \frac{K}{L}, \frac{T}{L}\right)$$

where (TFP_A) is the total factor productivity in the agricultural sector, (P_a/P_m) is the price of the agricultural activity p_a compared to that of the import-competing activity, (p_m), (K/L) is the overall capital-labor ratio, and (T/L) is the land-labor ratio. It is expected that $f_1 > 0$ if new techniques are capital intensive, and an increase in the relative price of agriculture reduces the relative cost of capital, and $f_2 > 0$ if capital accumulation leads to a reduction the relative price of capital. If an increase in the land-labor ratio reduces de relative price of capital, then $f_2 > 0$.

III. The Results of the Estimation

There are at least two methodological issues with the estimation of the TFP function. First, variables can respond to changes in other variables with lags and this introduces a short-run dynamics into the system of equations.

If this is the case, then economic theory could tell us very little about the identification of short-run relationships. This suggests using a vector autoregressive representation of the system of equations through which long-run relationships can be identified. Consider first the following autoregressive model:

(1) $$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \ldots + \Pi_k X_{t-k} + \varepsilon_t \qquad (t=1,\ldots,T)$$

where ε_t's are independent Gaussian variables with 0 mean and variance Ω, and X_t is a $p \times 1$ vector of stochastic variables.

Secondly, many economic variables are non stationary and estimating a functional relationship in the levels of the variables could lead to find spurious correlations between them as they have common trends. Proper differencing of the variables can remove the common trends and they are thus going to be uncorrelated. If the variables are non stationary, the vector autoregressive model can then be rewritten as:

(2) $$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{t-k+1} \Delta X_{k-1} + \Pi X_{t-1} + \varepsilon_t$$

where $\Delta = 1 - L$, and L is the lag operator

$$\Gamma_i = -\sum_{j=i+1}^{k} \Pi_j \quad \text{, and}$$

$$\Pi = -\left(1 - \sum_{i=1}^{k} \Pi_i\right)$$

Since $\Delta X_t, \ldots, \Delta X_{t-k+1}$ are stationary, that is, I(0) butbut X_{t-1} is $I(1)$, in order that this equation be consistent, Π should not be of full rank, say, of rank r. The hypothesis that the rank of Π is r can be formulated as the restriction that $\Pi = \alpha\beta'$ where α and β are $p \times r$ vectors and the vector β is the cointegrating vector with the property that $\beta' X$ is stationary. If the hypothesis that $r = 0$ is rejected, then the matrix Π contains information about long-run relationships between the variables in the data. The vector α is usually interpreted as the average rate of adjustment of the variables towards their long run equilibrium values. Campbell and Shiller (1988), however, demonstrate that error correction models do not necessarily reflect partial adjustment which, in turn, is the result of adjustment costs. They show that error correction models may also arise because one variable helps to forecast another.

Johansen (1990, 1991) has developed two test statistics to test the cointegration rank of the Π matrix, namely, the eigenvalue and the trace statistics. Asymptotic critical values for these test statistics are provided by Doornik, J. A. (1998). The asymptotic distribution of the test statistics depends upon the assumptions about the deterministic terms included in (2).

Podivinski (1990) finds that the tabulated critical values of Johansen´s tests based on the asymptotic distribution may be inappropriate when applied to sample sizes of 100 or smaller. S. Johansen (2002) derives a Bartlett correction factor of the trace test statistic to improve its finite sample properties[5]. The Bartlett procedure amounts to find the

[5] A small sample Bartlett correction of the maximum eigenvalue test statistic has not been developed in the literature. Cheung and Lai (1993) estimated a response surface function to correct both the trace and the maximum eigenvalue statistics for the small sample bias as a function of the sample size and of the degrees of freedom. However, their estimates used the table A2 of Johansen and Juselius (1990) that includes only five variables and it cannot be used for larger dimensions of the vector autoregression. If the Cheung-Lai correction is applied to the maximum eigenvalue statistic of the five variable vector autoregression, the results indicate that the null hypothesis of one cointegration vector cannot be rejected by the data.

expectation of the likelihood ratio test and correcting it to have the same mean as the limit distribution. The correction factor is a function of the estimated values of the parameters $(\hat{\alpha}, \hat{\beta}, \hat{\Gamma}_i, \text{and } \hat{\Omega})$ under the null hypothesis about the number of cointegration vectors and of the deterministic terms, and under the assumption of Gaussian errors. If, for instance, it is assumed that $r = 0$, then the correction factor will only be a function of $(\hat{\Gamma}_i \text{ and } \hat{\Omega})$. If, on the other hand, $r = n$, the correction factor is calculated using the estimates of $(\hat{\alpha}, \hat{\beta}, \hat{\Gamma}_1, ..., \hat{\Gamma}_n, \text{and } \hat{\Omega})$.

The TFP function is estimated with annual data covering the period 1941-1984 for which the needed data is available with data of the years 1939 and 1940 used as initial conditions[6]. The unrestricted parameters of the vector autoregression (1) are estimated with two lags in the levels of the variables based on the likelihood ratio test and the Hannan and Quinn criterion. In small samples, however, the use of the likelihood ratio test would lead to spurious rejection of the null hypothesis because the small sample distribution of the test statistic differs from its asymptotic distribution. Thus, the likelihood ratio test is adjusted for degrees of freedom to correct the small sample bias of the unadjusted likelihood ratio.

The underlying assumptions of the statistical model that is, that the residuals are normally distributed, uncorrelated and homoskedastic, are tested in order to ensure that the statistical properties of the estimates are met. The test of the null hypothesis of Gaussian residuals is based on the multivariate Jarque-Bera test statistics as proposed by Doornik and Hansen (1994). The Doornik and Hansen´s procedure transforms skewness and kurtosis to approximately χ^2 in small samples. The residuals are orthogonalized according to the procedure of Doornik and Hansen (1994) that makes the test statistic invariant with respect to the ordering of the variables (the alternative Choleski orthogonalization depends upon the ordering of the variables) and to the scaling of the variables (as it uses the correlation rather than the covariance matrix of residuals). For the system as a whole, the null hypothesis of normality of residuals cannot be rejected by the data as the $\chi^2(8)$ test statistic is calculated for the system as a whole in 8.32 with a marginal significance level (the *p*-value) of 0.40.

[6] The results were obtained using CATS in RATS, version 2.

The null hypothesis of serially uncorrelated residuals is also tested as residual correlation yields inconsistent estimates of the parameters. The multivariate Lagrange multiplier test statistics at one and two lags of the residuals are calculated in 5.4 and 5.8 with p-values of 0.99 in both cases, respectively, and these values indicate that the null hypothesis cannot be rejected by the data. In addition, the null hypothesis of no autoregressive conditional heteroskedastic disturbances cannot be rejected by the data as the multivariate Lagrange multiplier test statistics at one and two lags of the residuals that are approximately distributed as χ^2 with 100 and 200 degrees of freedom are calculated in 90.2 and 179.5 with marginal probabilities of 0.75 and 0.85, respectively.

In order to test the rank of the Π matrix, the model (2) is fitted with one lag of the variables in first differences and a constant in the cointegration space and a linear trend in the data as most of the variables seem to have a trend in their levels. The results of the tests of the rank of Π are presented in Table 2. The Barttlet corrected trace statistic[7] is calculated in 59.41 and this amounts to reject the hypothesis of no cointegration vector (r = 0) with a marginal probability of 0.002. The hypothesis of one cointegrating vector cannot be rejected with a marginal probability of 0.22 and thus the data supports the existence of one cointegrating vector.

Table 2
Trace Test Statistics for Testing Cointegrating Vectors

r	Trace Statistic	Trace Statistic*	p-value*
0	68.58	59.41	0.002
1	25.97	23.76	0.218
2	11.45	10.03	0.284
3	0.11	0.10	0.749

Note: The model includes a constant in the cointegration space and a linear trend in the data. The corrected trace statistic (*) is the trace statistic divided by the Barttlet correction factor. The p-values (*) are approximated using the Γ- distribution, see Doornik (1998).

[7] The parameters to calculate the correction factor have not been tabulated in Johansen, Nielsen, and Fachin (2005) for an unrestricted constant. This problem is avoided by using the parameters of a slightly larger model with a linear trend restricted to the cointegration space that is the same as under the null hypothesis of no trend but the tails of the distribution of the trace statistic are larger than those of the model with an unrestricted constant.

Usually, univariate tests for unit roots precede tests for cointegration. These tests may have low power because they are based on univariate time series and do not take into account information in related series[8]. Thus, stationarity of individual series can be formulated in terms of the parameters in the multivariate system as a null hypothesis given the cointegration space. If economically meaningful variables included in the system are found to be stationary, then an extra cointegrating vector is added to the cointegration space. The test statistic is distributed asymptotically as $\chi^2(p\text{-}r)$ where p is the number of variables in the system and r is the number of cointegrating vectors. The results of the tests are presented in Table 3 and they show that the null hypothesis of stationary variables is strongly rejected under the hypothesis of one cointegrating vector.

Table 3
Test of Stationarity of Variables

r	5% critical values	$\ln(TFP)$	$\ln(p_a/p_m)$	$\ln(K/L)$	$\ln(T/L)$
1	7.82	40.11 (0.00)	31.58 (0.00)	39.56 (0.00)	27.76 (0.00)
2	5.99	12.97 (0.00)	9.29 (0.01)	11.57 (0.00)	2.85 (0.24)
3	3.84	9.90 (0.00)	7.51 (0.01)	11.23 (0.00)	2.60 (0.11)

Notes: The numbers in parenthesis are the p-values

The estimated cointegrating vector is as follows (the numbers in parenthesis are the asymptotic t-statistics corrected for degrees of freedom):

β coefficients:

$\ln(TFP_A)$	$\ln(P_a/P_m)$	$\ln(K/L)$	$\ln(T/L)$
1.000	-2.127	-3.219	-9.314
--	(-6.294)	(-6.592)	(-6.528)

[8] See Maddala, G. S., and In-Moo Kim (1999), pp. 231

All the estimated parameters are positive and significantly different from zero at the usual levels of significance. In particular, favorable terms of trade for the agricultural sector tend to increase the sectoral rate of adoption of new techniques as expected by the theoretical considerations of Section I. In addition, the results show that increases in the capital- and land-labor ratios have positive effects on the rate of implementation of new techniques.

The estimated α coefficients are presented below and they show that the coefficient of the cointegrating vector in the TFP equation is negative and statistically significant. This provides additional support to the existence of one cointegrating vector.

α coefficients:

$\Delta\ln(TFP_A)$	$\Delta\ln(P_a/P_m)$	$\Delta\ln(K/L)$	$\Delta\ln(T/L)$
-0.075	0.157	0.014	0.043
(-3.410)	(3.605)	(1.506)	(4.298)

The results of the estimation of the α coefficients also indicate that they are statistically significant in the case of the relative price and land-labor ratio variables. The calculated value of the likelihood test statistic (corrected for degrees of freedom) to test the null hypothesis of weak exogeneity of the relative price variable, that is, the hypothesis that the α coefficient in the relative price variable is zero, is calculated in 8.56 with a marginal probability of 0.003 that amounts to reject the null. Interpretation of causal orderings is not always straightforward.

As indicated earlier, Campbell and Shiller show that a causal ordering in a cointegrated vector autoregression can arise because one variable helps to predict another if economic agents have superior information than that of the econometrician. If, for instance, the rate of adoption of new techniques were a function of the present value of expected future relative prices and agents had superior information, then the estimated cointegrating vector would incorporate this superior information and would cause relative prices because it contains agents' forecasts about prices in the next period.

On the other hand, the value of the test statistic calculated under the null hypothesis that the land-labor ratio is weakly exogenous is 11.36 with a marginal significance of about 0.001 that amounts to reject the null hypothesis. The hypothesis that the overall capital-labor ratio is weakly exogenous cannot be rejected as the likelihood ratio test statistic is calculated in 1.90 with a marginal probability of about 0.17.

One interpretation of the endogeneity of the overall land-labor ratio could be that the cointegrating vector contains, besides the aforementioned relative price forecast component, information about stationary supply shocks with zero mean. As the measure of the endowment of land only includes cultivated area with agricultural crops and excludes the pasture area devoted to livestock raising, it could be that supply shocks affecting the production function of agricultural crops would have an impact on the cultivated area devoted to these crops vis-à-vis that of livestock production. As a result, a causal ordering would follow between the cointegrating vector and the land-labor ratio.

So far, the theoretical framework does not include any macroeconomic variable affecting farmers´ decisions about the rate of adoption of new production techniques. Mundlak, Cavallo and Domenech (1989) find that high inflation rates have a negative impact on the rate of adoption of new techniques in the agricultural sector. A measure of macroeconomic disequilibria, namely, the fiscal deficit, is included in the system of variables. The fiscal deficit could negatively affect incentives to implement available capital-intensive techniques if it would be associated by economic agents with current and future taxation that negatively affect current expectations of future agricultural relative prices.

To this effect, a five-variable vector autoregression is estimated including a measure of the fiscal deficit (*d*) defined as the change in foreign and domestic indebtedness of the overall consolidated public sector as a percentage of the gross domestic product at current prices. The vector autoregression (2) is estimated with one lag in the first differences of the variables and an unrestricted constant as most of the variables seem to have a linear trend in their levels[9].

As done before, the assumptions about the behavior of residuals are tested. The null hypothesis of normality of residuals cannot be rejected as the calculated $\chi^2(10)$ test statistic is calculated in 8.82 with a marginal

[9] If the statistical model is estimated with a trend restricted to the cointegration space, the trend variable is not statistically different from zero and this amounts not to reject the model with an unrestricted constant. These results are available from the author upon request.

probability of 0.55. On the other hand, the Lagrange multiplier test statistics calculated under the null hypothesis of no autoregressive conditional heteroskedasticity of the residuals at one and two lags that are distributed as $\chi^2(225)$ and $\chi^2(450)$ are 202.52 and 451.14, respectively, and these values amount not to reject the null with marginal probabilities of 0.86 and 0.48. The null hypothesis of uncorrelated residuals cannot be rejected as the Lagrange multiplier test statistics calculated under the null hypothesis of no residual autocorrelation at one and two lags are 14.22 and 16.62 with marginal probabilities of 0.96 and 0.90, respectively.

The rank of the Π matrix is tested with the trace test statistic using the Barttlet correction mentioned earlier. The results (see table 4) indicate that the hypothesis of one cointegating vector cannot be rejected by the data. Under the null hypothesis of $r = 0$, the Barttlet corrected trace statistic is calculated in 80.96 with a marginal significance of 0.004 that amounts to reject the null. The null hypothesis of $r = 1$ cannot be rejected as the corrected trace test statistic is calculated in 40.60 with a marginal probability of about 0.20.

Table 4
Trace Test Statistics for Testing Cointegrating Vectors

r	Trace Statistic	Trace Statistic*	p-value*
0	95.63	80.96	0.004
1	46.13	40.60	0.203
2	23.97	21.77	0.322
3	6.51	5.79	0.722
4	0.00	0.00	0.981

Note: See Table 2

The $\chi^2(4)$ tests statistics calculated under the null hypothesis of stationarity of the variables are shown below and they indicate that for $r = 1$, the null hypothesis is strongly rejected.

The results of the test of stationary variables (the marginal significance levels are shown in parentheses) are the following:

$\ln(TFP_A)$	$\ln(P_a/P_m)$	$\ln(K/L)$	$\ln(T/L)$	d
46.77	35.08	47.01	34.48	39.57
(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

The estimated parameters of the cointegrating vector are presented below and they show that all of them are statistically different from zero at the usual significance levels. In particular, the coefficient of the fiscal deficit variable is negative and statistically different from zero at the usual significance levels. A plausible interpretation given in this paper is that of the associated expected increase in current and future distortionary taxes to finance current deficits that could negatively affect the economic incentives of the agricultural sector. An alternative explanation could be that the expansion of credit to the public sector needed to finance the fiscal deficit could have a crowding out effect over the private sector by increasing the real interest rate and reducing thus the incentives to adopt more capital intensive techniques.

The β coefficients (the numbers in parenthesis are asymptotic t-statistics corrected for degrees of freedom) are the following :

$\ln(TFP_A)$	$\ln(P_a/P_m)$	$\ln(K/L)$	$\ln(T/L)$	d
1.000	-0.790	-1.679	-4.018	0.018
--	(-5.864)	(-7.967)	(-6.912)	(2.448)

The estimated α coefficients are shown below:

$\Delta\ln(TFP_A)$	$\Delta\ln(P_a/P_m)$	$\Delta\ln(K/L)$	$\Delta\ln(T/L)$	Δd
-0.217	0.347	0.017	0.099	-3.874
(-4.804)	(3.412)	(0.788)	(4.312)	(-1.435)

The null hypotheses of weak exogeneity of the capital/labor ratio and the fiscal deficit variables cannot be rejected as the likelihood ratio tests statistics are calculated in 0.52 and 1.57, respectively, with marginal probabilities of 0.47 and 0.21. On the other hand, the null hypotheses of weakly exogenous relative prices and land-labor ratio variables are rejected with marginal significance levels of 0.008 and 0.001, respectively.

The estimates presented above could have been subject to some bias if the estimated total factor productivity does not reflect the actual weights of the production factors in the agricultural production function. In order to test the hypothesis that these weights do not differ from the actual ones, a seven variable vector autoregression is estimated including the agricultural capital-labor K_{ag}/L_{ag} and the land-labor ratios (T/L_{ag}). If there were measurement errors in the calculation of the total factor productivity, then, under constant returns to scale in the agricultural production function, the ratio of agricultural value added to the calculated index of primary factors would be also a function of the agricultural capital and land-labor ratios. It is thus expected that the long-run parameters of these variables do not differ from zero if the coefficients used to calculate the TFP are accurately measured.

The results of the estimation of the seven variable vector autoregression indicate that the hypothesis of two cointegration vectors cannot be rejected by the data (see Table 5 below). When the number of cointegrating vectors is higher than one there is an identification problem because linear combinations of the cointegrating vectors are also cointegrating relationships and thus the parameters of the vectors are not identified.

Table 5
Trace Test Statistics for Testing Cointegrating Vectors

r	Trace Statistic	Trace Statistic*	p-value*
0	147.39	109.94	0.003
1	95.60	73.67	0.022
2	48.74	36.84	0.359

Note: See Table 2

One method of identifying the long-run parameters of the cointegrating relationships is the triangular representation of Phillips (1991). Under this representation, the parameters of the cointegrating variables are expressed as functions of the non-cointegrating variables. Let these two cointegration vectors be:

$$\begin{bmatrix} 1 & \beta_{12} & \beta_{13} & \beta_{14} & \beta_{15} & \beta_{16} & \beta_{17} \\ \beta_{21} & \beta_{22} & \beta_{23} & \beta_{24} & \beta_{25} & \beta_{26} & \beta_{27} \end{bmatrix} . \begin{bmatrix} \ln TFP_{sg} \\ \ln(K_{ag} / L_{ag}) \\ \ln(T / L_{ag}) \\ \ln(P_x / P_m) \\ \ln(K / L \\ \ln(T / L) \\ d \end{bmatrix}$$

It can be shown that, in the triangular representation of Phillips, the coefficient of the $\ln(T / L_{ag})$ variable is equal to $(\beta_{13} - \beta_{12}\beta_{23})/(1 - \beta_{21}\,\beta_{12})$ if the second vector is normalized on the agricultural capital-labor ratio, $\ln(K_{ag}/L_{ag})$. On the other hand, if the second cointegrating relationship is normalized on the agricultural land-labor ratio, $\ln(T / L_{ag})$, then the coefficient of is equal to $(\beta_{12} - \beta_{13}\beta_{22})/(1 - \beta_{21}\,\beta_{23})$.

If the β_{12} and β_{13} coefficients were zero, then, under the triangular representation the coefficients of the agricultural capital-labor ratio or the land-labor ratio in the first cointegration relationship would be equal to zero, leaving aside compensation of parameters, depending upon the normalization of the second vector. In addition, the coefficient of adjustment of the second cointegration vector in the $\Delta\ln(TFP_{ag})$ equation should not be statistically different from zero. Thus, a likelihood ratio test statistic can be calculated under these null joint hypotheses and it is distributed as χ^2 with two degrees of freedom.

The first cointegrating vector is normalized on the total factor productivity variable and the second one on the variable measuring the agricultural capital-labor ratio. Although the trace statistic indicates that the hypothesis of two cointegrating vectors cannot be rejected by the data, the $\chi^2(1)$ likelihood ratio test statistic calculated under the hypothesis that the coefficient of adjustment of the second cointegrating vector in the $\Delta\log(K_{ag}/L_{ag})$ equation is equal to zero is 0.2318 and this amounts not to reject the null with a marginal probability of 0.63.

Under the normalization on the variable measuring the (logarithm of) total factor productivity, the $\chi^2(2)$ likelihood ratio test statistic (adjusted for degrees of freedom) calculated under the null hypothesis that the agricultural capital- and land-labor ratios are equal to zero is equal to 1.986 that amounts not to reject the null hypothesis with a p-value of 0.37.

It does seem that the parameters used to calculate the total factor productivity index does not contain any significant bias leading to wrong statistical inferences.

On the other hand, if the second cointegrating vector is normalized on the agricultural land-labor ratio, then the likelihood ratio test under the null hypothesis that the coefficient of adjustment of this vector in the $\Delta \ln(T / L_{ag})$ equation is zero is rejected by the data and this result supports the existence of two cointegrating vectors. The $\chi^2(2)$ likelihood ratio test statistic (adjusted for degrees of freedom) under the aforementioned joint null hypotheses is calculated in 1.7992 that amounts not to reject the null with a marginal probability of 0.40.

The stability of the long-run parameters over time is analyzed. Hansen and Johansen (1999) suggest a graphical procedure to evaluate the constancy of the long-run parameters over time in cointegrated vector autoregressive models. The procedure is based on recursively estimated non-zero eigenvalues as these provide information about the adjustment coefficients and the cointegrated vectors. Non-constancy of these parameters will thus be reflected in the time path of the estimated eigenvalues.

The eigenvalues, λ_I, are transformed into $\xi_i = \ln(\lambda_i / 1 - \lambda_i)$ to obtain a better approximation of their limiting distribution and to ensure that the confidence bounds for λ_I stay in the interval [0, 1]. The time paths of the transformed estimated eigenvalue for the sub sample 1967-1984 with an autoregression vector containing five variables and one cointegrating vector are used as a diagnosis tool in the model evaluation. The size of the sub sample has been chosen as a function of the parameters of the model. The results are presented in Figure 2 and, although it is not a formal test of stability of parameters, they do not seem to indicate non-constancy of the parameters.

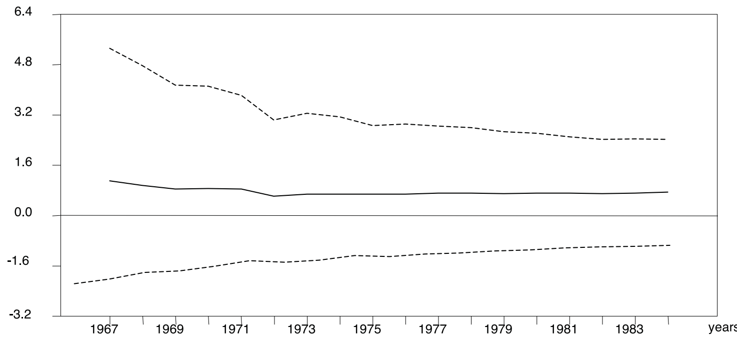

Fig. 2: Recursive estimates of the transformed eigenvalues, $\hat{\xi}_i = \log(\hat{\lambda}_i) - \log(1-\hat{\lambda}_i)$ (black solid line), with the 95% confidence bands (dotted lines), 1967-1984.

A formal test of stability of parameters over time developed by Hansen and Johansen (1999) is presented in Figure 3, in which there are the plots of the sample paths of:

$$\tau_T^{(t)}(\xi_i) = \frac{t}{T}\left|\left(T^{-1}\Sigma_{ii}^{0}\right)^{-\frac{1}{2}}\hat{\xi}_i^{(t)}\right|, \quad \text{i=1, 2}$$

and

$$\tau_T^{(t)}\left(\sum_{i=1}^{2}\xi_i\right) = \frac{1}{T}\left|\left(T^{-1}\sum_{j,k=1}^{2}\Sigma_{jk}^{0}\right)^{-\frac{1}{2}}\sum_{i=1}^{2}\hat{\xi}_i^{(t)}\right|$$

where $\Sigma_{j,k}^{0}$ is the variance of the transformed eigenvalues.

In the recursive analysis, the test statistics are calculated either by reestimating recursively all the parameters (the so-called *X*-form), or by reestimating only the long-run parameters *a* and *β* and concentrating out the short term coefficients (the *R1*-form). The fluctuations tests are sup tests and are generally regarded as conservative, meaning that if the null hypothesis of stability of parameters is rejected, it is a signal of rather large deviations from the null. The quantiles of their distribution have been tabulated by Ploberger, Krämer, and Kontrus (1989). It can be seen in Fig. 3 that the values of the test statistics are below the 0.05 critical level of 1.36 and, consequently, the hypothesis of constancy of parameters over time cannot be rejected.

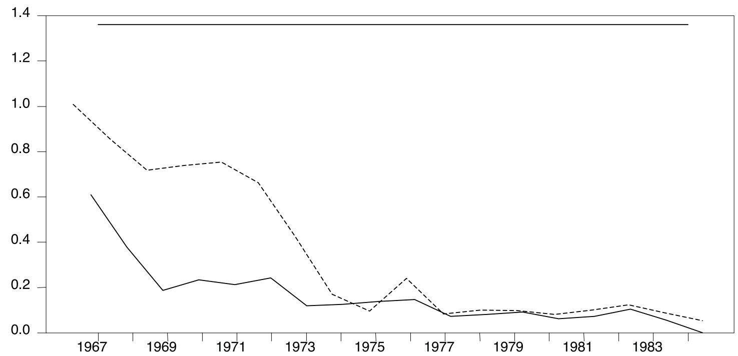

Figure 3: Fluctuation tests of the eigenvalues, 1967-1984. The black solid and black dotted lines correspond to the R1- and the X-forms of the test statistics, respectively. The horizontal black solid line corresponds to the critical value of the test statistics at 0.05 (1.36).

IV. Concluding Remarks

It is shown in this paper that the null hypothesis of endogenous total factor productivity in the Argentinean agricultural sector that is associated with technological change cannot be rejected by the data. In particular, it is found that economic incentives to the agricultural activity, namely, agricultural relative prices, have significant positive effects on the adoption of new techniques. The findings of this paper also indicates that the overall resource constraints of the economy, namely, the land- and capital-labor ratios have positive effects over the rate of implementation of newly available techniques of production in the agricultural sector. It is also found that the fiscal deficit has a negative and statistically significant effect on the total factor productivity.

The main lesson that can be learned from this paper is that policymakers who support import-substitution policies in Argentina have severely underestimated the response of the Argentinean exportable agricultural activity to the adoption of new techniques of production. These anti-trade policies have certainly contributed to the poor performance of the sector during the period 1941-1984 in which the average annual rate of growth of agricultural GDP (1.4 percent) was below the rate of growth of total population (1.7 percent) by depressing agricultural relative prices to foster import-substitution activities.

V. REFERENCES

Berlinski, J. (2003). "International Trade and Commercial Policy", in Della Paolera, G. and A. Taylor (eds.), *A New Economic History of Argentina*, Cambridge University Press.

Campbell, J. and R. Shiller, (1988). "Interpreting Cointegrated Models", Working Paper No 2568, *NBER,* April.

CATS in Rats, (2005) version 2, by J. G. Dennis, H. Hansen, S. Johansen, and K. Juselius, Estima.

Diaz Alejandro, C., (1970). *Essays on the Economic History of the Argentine Republic*, Yale University Press.

Doornik, J. and H. Hansen, (1994). "An Omnibus Test for Univariate and Multivariate Normality", Working Paper, Nuffield College, Oxford.

Doornik, J. A. (1998). "Approximation to Asymptotic Distribution of Cointegrating Tests", *Journal of Econometric Surveys*, 12, 573-593.

Elías, V. J. (1999). *Sources of Economic Growth*, International Center for Economic Growth.

Hansen, Henrik and Søren Johansen, (1999). "Some tests for parameter constancy in cointegrated VAR-models", *Econometrics Journal*, Royal Economic Society, vol. 2(2), pages 306-333.

Herrou-Aragon, A. (2006). "Factor Prices and Commercial Policies in the Argentinean Economy", Working paper.

IERAL, (1986). "Estadísticas de la Evolución Económica de Argentina 1913-1984", *Estudios*, No 39, July-September.

Johansen, S. (1991). "Estimation and Hypothesis Testing of Cointegation Vectors in Gaussian Vector Autoregressive Models", *Econometrica*, 59.

Johansen, S. and K. Juselius, (1990). "Maximum Likelihood Estimation and Inference in Cointegration – With Applications to the Demand for Money", *Oxford Bulletin of Economics and Statistics*, 52, 2.

Johansen, S. (2002). "A small sample correction for tests of hypotheses on the cointegrating vectors", *Journal of Econometrics*, Elsevier, vol. 111(2), pages 195-221, December.

Maddala, G. S. (2001). *Introduction to Econometrics*, 3rd edition, John Wiley and Sons, Ltd.

Maddala, G.S. (1999). and In-Moo Kim, *Unit Roots, Cointegration, and Structural Change*, Cambridge University Press.

Mundlak, Y., D. Cavallo, and R. Domenech, (1989). *Agriculture and Economic Growth in Argentina, 1913-1984*. Research Report 76, IFPRI, November.

Mundlak, Y. (2000) *Agriculture and Economic Growth*: Theory and Measurement, Harvard University Press.

Phillips, P. C. B. (1991). "Optimal inference in cointegrated systems", *Econometrica*, 59.

Reca, L., and J. Verstraeten, (1977). "La Formación del Producto Agropecuario Argentino: Antecedentes y Posibilidades", *Desarrollo Económico*, Vol. 17, Nº 67, Oct.-Dec. pp. 371-389.

ANNEX: DATA DESCRIPTION

A: Agricultural gross domestic product at 1960 prices. Source: IEERAL, op. cit.

(pa/pm): Price of wholesale agricultural goods divided by the wholesale price of imported goods. Source: from 1939 until 1965, Diaz-Alejandro, C. F., Ensayos sobre la Historia Económica Argentina, Amorrortu editores. From 1966 until 1984: INDEC.

K: Total stock of capital employed in production of goods and services in australes at 1960 prices. Source: IEERAL, op. cit.

T: Total planted area with agricultural crops in thousand of hectares weighted by the value of production of each crop. Source: IEERAL, op. cit.

L: Total labor force in million people. Source IEERAL, op. cit