Factor Productivity in the Argentinean Agriculture

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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 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.

Table 1: Sources of Agricultural Output Growth – Years 1940-1984

<table>
<thead>
<tr>
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<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Output growth</td>
<td>1.35</td>
<td>0.39</td>
<td>1.37</td>
<td>1.97</td>
<td>2.12</td>
<td>0.43</td>
</tr>
<tr>
<td>Input contribution of which:</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Land</td>
<td>0.32</td>
<td>-0.53</td>
<td>0.69</td>
<td>0.68</td>
<td>0.16</td>
<td>0.86</td>
</tr>
<tr>
<td>Labor</td>
<td>-0.10</td>
<td>0.28</td>
<td>-0.20</td>
<td>-0.11</td>
<td>-0.22</td>
<td>-0.42</td>
</tr>
<tr>
<td>Capital</td>
<td>0.20</td>
<td>-0.22</td>
<td>0.44</td>
<td>0.40</td>
<td>0.40</td>
<td>-0.18</td>
</tr>
<tr>
<td>Residual (TFP)</td>
<td>0.93</td>
<td>0.86</td>
<td>0.44</td>
<td>1.00</td>
<td>1.78</td>
<td>0.16</td>
</tr>
</tbody>
</table>

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).

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. 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. Elias (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.

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

1 For a description of the data used in this paper, see Annex.
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 these new policies, imports
required the approval of a committee formed by public officials and representatives of the
import-competing private sector.

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 1, a theoretical framework of endogenous technical
change (or changes in the TFP) is presented. The results of the estimations are in Section
2 and the concluding remarks are in Section 3.

I. 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 rate of implementation of new techniques. If there were an increase in the relative price of exportable agriculture resulting, for instance, from a reduction in import tariffs, this would in turn increase the marginal product of investing in new techniques in the exportable sector. On the other hand, changes in the relative price of the agricultural sector affect the relative costs of implementing new techniques. If an increase in relative prices causes an increase in the relative price of labor compared to that of capital, then new labor-saving techniques would be adopted.

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. An increase in the capital labor ratio would result in the adoption of new capital intensive techniques as the wage-rental ratio is expected to increase.

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 effects of the determinants of endogenous technological change in the agricultural sector \( (A) \) \( (TFP_A) \) can be summarized in a function such as the following,

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

where \( (TFP_A) \) is the total factor productivity in the agricultural sector, \( \left(\frac{p_a}{p_m}\right) \) is the price of the agricultural activity \( (p_a) \) compared to that of the import-competing activity \( (p_m) \), \( \left(\frac{K}{L}\right) \) is the overall capital-labor ratio, and \( \left(\frac{T}{L}\right) \) is the land-labor ratio. It is expected that \( f_1 > 0 \) and, if new techniques are capital intensive, then, \( f_2 > 0 \). If \( f_2 > 0 \) and if an increase in the land-labor ratio reduces de relative price of capital, then \( f_3 > 0 \).

II. **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

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

where \( \varepsilon_t \) are independent Gaussian variables with 0 mean and variance \( \Omega \), 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

\[ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{t-k} \Delta X_{t-k-1} + \Pi L X_{t-1} + \varepsilon_t \]

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

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

\[ \Pi = \begin{pmatrix} 1 \sum_{i=1}^k \Pi_i \end{pmatrix} \]

Since \( \Delta X_t, \ldots, \Delta X_{t-k+1} \) are stationary, that is, \( I(0) \) but \( X_{t-1} \) is \( I(1) \), in order that this equation be consistent, \( \Pi \) should not be of full rank, say, of rank \( r \). The hypothesis that the rank of \( \Pi \) is \( r \) can be formulated as the restriction that \( \Pi = \alpha \beta' \) where \( \alpha \) and \( \beta \) are \( p \times r \) vectors and the vector \( \beta \) is the cointegrating vector with the property that \( \beta' X \) is stationary. If the hypothesis that \( r=0 \) is rejected, then the matrix \( \Pi \) contains information about long-run relationships between the variables in the data. The vector \( \alpha \) 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 \( \Pi \) 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. The Bartlett procedure amounts to find the 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.
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 \( \hat{\gamma}_1 \) and \( \hat{\Omega} \). If, on the other hand, \( r = n \), the correction factor is calculated using the estimates of \( (\hat{\alpha}, \hat{\beta}, \hat{\gamma}_1, ..., \hat{\gamma}_n, \hat{\Omega}) \).

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. 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 \( \chi^2 \) 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 \( \chi^2 \) test statistic is calculated for the system as a whole in 8.32 with a marginal significance level (the \( p \)-value) of 40 percent.

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 99 percent 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 \( \chi^2 \) with 100 and 200 degrees of freedom are calculated in 90.2 and 179.5 with marginal probabilities of 75 and 85 percent, respectively.

In order to test the rank of the \( \Pi \) 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 \( \Pi \) are presented in Table 2. The Bartlett corrected trace statistic is \( \chi^2 \) with 100 and 200 degrees of freedom are calculated in 90.2 and 179.5 with marginal probabilities of 75 and 85 percent, respectively.

\(^2\) 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
calculated in 59.41 and this amounts to reject the hypothesis of no cointegration vector \((r=0)\) with a marginal probability of 0.2 percent. The hypothesis of one cointegrating vector cannot be rejected with a marginal probability of 22 percent and thus the data supports the existence of one cointegrating vector.

Table 2: Trace Test Statistics for Testing Cointegrating Vectors

<table>
<thead>
<tr>
<th>(r)</th>
<th>Trace Statistic</th>
<th>Trace Statistic*</th>
<th>(p)-value*</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>68.58</td>
<td>59.41</td>
<td>0.002</td>
</tr>
<tr>
<td>1</td>
<td>25.97</td>
<td>23.76</td>
<td>0.218</td>
</tr>
<tr>
<td>2</td>
<td>11.45</td>
<td>10.03</td>
<td>0.284</td>
</tr>
<tr>
<td>3</td>
<td>0.11</td>
<td>0.10</td>
<td>0.749</td>
</tr>
</tbody>
</table>

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 Bartlett correction factor. The \(p\)-values (*) are approximated using the \(\Gamma\)-distribution, see Doornik (1998).

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\(^3\). 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 \(\chi^2(p-r)\) 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

<table>
<thead>
<tr>
<th>(r)</th>
<th>5% critical values</th>
<th>(\ln(TFP))</th>
<th>(\ln(p/p_m))</th>
<th>(\ln(K/L))</th>
<th>(\ln(T/L))</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>7.82</td>
<td>40.11</td>
<td>31.58</td>
<td>39.56</td>
<td>27.76</td>
</tr>
<tr>
<td></td>
<td>(0.00)</td>
<td>(0.00)</td>
<td>(0.00)</td>
<td>(0.00)</td>
<td>(0.00)</td>
</tr>
<tr>
<td>2</td>
<td>5.99</td>
<td>12.97</td>
<td>9.29</td>
<td>11.57</td>
<td>2.85</td>
</tr>
<tr>
<td></td>
<td>(0.00)</td>
<td>(0.01)</td>
<td>(0.00)</td>
<td>(0.01)</td>
<td>(0.24)</td>
</tr>
<tr>
<td>3</td>
<td>3.84</td>
<td>9.90</td>
<td>7.51</td>
<td>11.23</td>
<td>2.60</td>
</tr>
<tr>
<td></td>
<td>(0.00)</td>
<td>(0.01)</td>
<td>(0.00)</td>
<td>(0.00)</td>
<td>(0.11)</td>
</tr>
</tbody>
</table>

Notes: The numbers in parenthesis are the \(p\)-values

\(^3\) See Maddala, G. S., and In-Moo Kim (1999), pp. 231
The estimated cointegrating vector is as follows (the numbers in parenthesis are the t-statistics):

**β coefficients:**

<table>
<thead>
<tr>
<th></th>
<th>$\ln(TFP_a)$</th>
<th>$\ln(P_a/P_m)$</th>
<th>$\ln(K/L)$</th>
<th>$\ln(T/L)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1.00</td>
<td>-2.13</td>
<td>-3.22</td>
<td>-9.32</td>
<td></td>
</tr>
<tr>
<td>--</td>
<td>(-7.06)</td>
<td>(-7.39)</td>
<td>(-7.32)</td>
<td></td>
</tr>
</tbody>
</table>

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 coefficients of the resource constraints are all positive suggesting that new techniques are capital intensive or labor saving.

The estimated $\alpha$ 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 evidence to support the existence of one cointegrating vector.

**α coefficients:**

<table>
<thead>
<tr>
<th></th>
<th>$\Delta \ln(TFP_a)$</th>
<th>$\Delta \ln(P_a/P_m)$</th>
<th>$\Delta \ln(K/L)$</th>
<th>$\Delta \ln(T/L)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>-0.075</td>
<td>0.157</td>
<td>0.014</td>
<td>0.043</td>
<td></td>
</tr>
<tr>
<td>(-3.823)</td>
<td>(4.042)</td>
<td>(1.689)</td>
<td>(4.189)</td>
<td></td>
</tr>
</tbody>
</table>

The results of the estimation of the $\alpha$ 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 by degrees of freedom) to test the null hypothesis of weak exogeneity of the relative price variable, that is, the hypothesis that the $\alpha$ coefficient in the relative price variable is zero, is calculated in 8.56 with a marginal probability of 0.3 percent 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 0.1 percent 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 17 percent.
An 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 agricultural 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 Domenec (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 adopt available 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.

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 probability of 55 percent. 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 86 and 48 percent. 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 96 and 90 percent, respectively.

The rank of the \(\Pi\) 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 cointegrating 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 0.4 percent 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 20 percent.

Table 4: Trace Test Statistics for Testing Cointegrating Vectors

<table>
<thead>
<tr>
<th>(r)</th>
<th>Trace Statistic</th>
<th>Trace Statistic*</th>
<th>(p)-value*</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\[\text{Table 4: Trace Test Statistics for Testing Cointegrating Vectors}\]
<table>
<thead>
<tr>
<th></th>
<th>0</th>
<th>95.63</th>
<th>80.96</th>
<th>0.004</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>46.13</td>
<td>40.60</td>
<td>0.203</td>
<td></td>
</tr>
<tr>
<td>2</td>
<td>23.97</td>
<td>21.77</td>
<td>0.322</td>
<td></td>
</tr>
<tr>
<td>3</td>
<td>6.51</td>
<td>5.79</td>
<td>0.722</td>
<td></td>
</tr>
<tr>
<td>4</td>
<td>0.00</td>
<td>0.00</td>
<td>0.981</td>
<td></td>
</tr>
</tbody>
</table>

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:

Test of stationary variables (the marginal significance levels are shown in parentheses):

\[
\begin{array}{cccccc}
\text{ln}(TFP_A) & \text{ln}(P_a/P_m) & \text{ln}(K/L) & \text{ln}(T/L) & d \\
46.77 & 35.08 & 47.01 & 34.48 & 39.57 \\
(0.00) & (0.00) & (0.00) & (0.00) & (0.00)
\end{array}
\]

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 and with the expected signs. 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.

\( \beta \) coefficients (the numbers in parenthesis are the \( t \)-statistics):

\[
\begin{array}{cccccc}
\text{ln}(TFP_A) & \text{ln}(P_a/P_m) & \text{ln}(K/L) & \text{ln}(T/L) & d \\
1.00 & -0.79 & -1.78 & -4.02 & 0.02 \\
-- & (-6.77) & (-9.20) & (-7.98) & (2.83)
\end{array}
\]

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 estimated \( \alpha \) coefficients are shown below:

\( \alpha \) coefficients:

\[
\begin{array}{cccccc}
\Delta \ln(TFP_A) & \Delta \ln(P_a/P_m) & \Delta \ln(K/L) & \Delta \ln(T/L) & \Delta d \\
-0.22 & 0.35 & 0.20 & 0.10 & -3.87 \\
(-5.55) & (3.94) & (0.91) & (4.98) & (-1.65)
\end{array}
\]

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 47 and 21 percent. On the other
hand, the null hypotheses of weakly exogenous relative prices and land-labor ratio
variables are rejected with marginal significance levels of 0.8 and 0.1 percent,
respectively.

III. 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. 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 policies have certainly contributed to the poor
performance of the sector during the period 1941-1984 in which the 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.
References


Annex: Data Description


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.