

Agricultural Supply Response in the Argentinean Economy

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In a paper by A. Herrou-Aragón (2001), it was found that the import-substitution policies of the 70s and the 80s in Argentina resulted in heavy taxation of international trade. In the paper, it is estimated that taxation of the volume of international trade in the Argentinean economy resulting from protectionist policies was about 115 and 80 percent during the 70s and in the 80s, respectively. Short-lived trade liberalization policies were implemented in the late 70s that reduced the overall taxation of trade to about 18 percent but these policies were reversed later on as a result of macroeconomic imbalances. In the 90s, the government introduced in-depth trade reforms that resulted in a reduction of international trade taxation to about 20 percent.

Import-substitution policies resulting in heavy taxation of international trade have certainly harmed the growth potential of the Argentinean exportable sector during those years. The extent of this effect has been subject to considerable debate depending upon the optimism or pessimism about the supply price elasticity of the main export commodity of the Argentinean economy, namely, the agricultural sector. This is, of course, an empirical question that several authors have tried to address.

L. G. Reca (1980) estimated short- and long-run supply price elasticities with data covering the years 1950-1974 using a partial equilibrium Nerlovian model of adjustment. The estimated short- and long- run elasticities are in a range of 0.20-0.35, and 0.4-0.5, respectively, depending on the specification of the model. These estimates are within the range of previous estimates by the author (Reca, 1969) for the years 1934/35-1966/67, and by R. Colomé (1966)

A very comprehensive general equilibrium study about the response of the Argentinean agricultural sector to economic incentives is that of Mundlak, Cavallo and Domenech (1989). They develop a model in which the rate of adoption of new techniques depending upon relative prices of the commodities, sectoral capital-labor ratios, and several macroeconomic variables such as bank crises, inflation etc.. The sectoral allocation of resources in response to economic incentives is also estimated. Their estimates for the period 1913-1984 are consistent with the theoretical model and they are used to simulate the implicit short and long run responses of the Argentinean agricultural sector to prices and other policy interventions.

The results of the simulations of Mundlak, *et. al.*, (1989) indicate that there is a sluggish response of the agricultural sector to economic incentives. According to the simulations, there is 0.40 percent response of agricultural output to a one percent change in its relative price that is achieved in five years. A unitary supply price response to prices is achieved in about twenty years when the output of the agricultural sector reaches its equilibrium state.

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This paper is aimed at presenting new estimates of the long-run supply price elasticity of the agricultural sector and of the speed of adjustment to its stationary state. An aggregate agricultural supply function is derived within the framework of a general equilibrium model and includes relative prices, overall endowments of productive factors and a measure of the government's size as explanatory variables. The agricultural supply function is estimated with the econometric methodology of cointegration analysis of vector autoregression models in which the long-run parameters are estimated without any theoretical specification of the short-run adjustment process.

In Section 1, a general equilibrium agricultural supply function is derived within a model of three goods, one of which is a non traded good, and three factors of production. The results of the estimates of the agricultural supply function are presented in Section 2. The concluding remarks are in Section 3.

1. Some Theoretical Issues

In this section, an aggregate supply function of the agricultural sector is derived within a general equilibrium framework of three goods and three factors of production. The general equilibrium model developed in this section is similar in structure to that of R. Jones (1965), R. Batra and F. Casas (1976), F. Rivera-Batiz (1982), and R. Jones and S. Easton (1983). The main difference with these models is that there are three commodities, one of which is non traded, and three factors of production.

The model includes three goods, namely, exportable agricultural (A), import substitution (M), and non traded goods (H). Factors of production, on the other hand, include land (T), capital (K) and labor (L). Production of the three goods is subject to constant returns to scale. Agriculture is assumed to be intensive in the use of land, and import substitution goods are intensive in the use of capital. Non-traded goods are assumed to be labor intensive. The economy is assumed to be small enough to take the prices of A and M as given in international equilibrium and the price of H is endogenously determined by domestic demand and supply conditions.

The three factors of production are assumed to be mobile among the three sectors of the economy, and prices are flexible in order to achieve full employment of all resources. It is also assumed that the supply of the three factors is perfectly inelastic. In each market, competition and non specialization in production prevail in order to guarantee that average production costs equal the market prices.

Let $a_{i,j}$ denote the amount of input i required to produce one unit of good j . Under full employment,

$$\begin{aligned} a_{TA}A + a_{TM}M + a_{TH}H &= T \\ a_{KA}A + a_{KM}M + a_{KH}H &= K \\ a_{LA}A + a_{LM}M + a_{LH}H &= L \end{aligned} \tag{1}$$

Let w_L , w_K , and w_T be the wage rate, the capital and land rental, respectively. Under zero profit conditions in each market,

$$\begin{aligned}
a_{TA} w_T + a_{KA} w_K + a_{LA} w_L &= p_A \\
a_{TM} w_T + a_{KM} w_K + a_{LM} w_L &= p_M \\
a_{TH} w_T + a_{KH} w_K + a_{LH} w_L &= P_H
\end{aligned} \tag{2}$$

Under the assumption of production functions with constant returns to scale, each input-output coefficient is independent of the scale of output and is a function of factor prices,

$$a_{ij} = a_{ij}(w_T, w_K, w_L) \tag{3}$$

where each function is homogenous of degree zero in all input prices. Let E_{ij}^z be the elasticity of a_{ij} with respect to price of input z , holding constant the prices of the other two factors, that is,

$$E_{ij}^z = \frac{\partial a_{ij}}{\partial w_z} \cdot \frac{w_z}{a_{ij}} \tag{4}$$

The zero homogeneity condition of each a_{ij} implies,

$$\sum_z E_{ij}^z = 0 \tag{5}$$

The exogenous variables of the system of equations (1)-(5) are the three factor endowments (L, T , and K), p_A and p_M . In order to derive an aggregate agricultural supply function as a function of the exogenous variables of the system, the full employment conditions are totally differentiated,

$$\begin{aligned}
\lambda_{TA} \hat{A} + \lambda_{TM} \hat{M} + \lambda_{TH} \hat{H} &= \hat{T} - (\lambda_{TA} \hat{a}_{TA} + \lambda_{TM} \hat{a}_{TM} + \lambda_{TH} \hat{a}_{TH}) \\
\lambda_{KA} \hat{A} + \lambda_{KM} \hat{M} + \lambda_{KH} \hat{H} &= \hat{K} - (\lambda_{KA} \hat{a}_{KA} + \lambda_{KM} \hat{a}_{KM} + \lambda_{KH} \hat{a}_{KH}) \\
\lambda_{LA} \hat{A} + \lambda_{LM} \hat{M} + \lambda_{LH} \hat{H} &= \hat{L} - (\lambda_{LA} \hat{a}_{LA} + \lambda_{LM} \hat{a}_{LM} + \lambda_{LH} \hat{a}_{LH})
\end{aligned} \tag{6}$$

where an asterisk denotes a rate of change, and where λ_{ij} is the proportion of the total supply of input i used in the j sector and, of course, $\lambda_{iA} + \lambda_{iM} + \lambda_{iH} = 1$. In the system of equations (6), $\sum_j \hat{a}_{Tj}$, for instance, shows the change in total quantity of land (T) required by the economy, as a result of changes in factor proportions in each sector at unchanged outputs, to maintain full employment of the resource.

The rate of change in sectoral outputs can be defined as functions of the rates of change of input-output coefficients, and of rates of change in factor endowments by inverting the matrix of input requirements (λ),

$$\begin{pmatrix} \hat{A} \\ \hat{M} \\ \hat{H} \end{pmatrix} = \begin{pmatrix} \lambda_{TA}^{-1} & \lambda_{TM}^{-1} & \lambda_{TH}^{-1} \\ \lambda_{KA}^{-1} & \lambda_{KM}^{-1} & \lambda_{KH}^{-1} \\ \lambda_{LA}^{-1} & \lambda_{LM}^{-1} & \lambda_{LH}^{-1} \end{pmatrix} \cdot \begin{pmatrix} \hat{T} \\ \hat{K} \\ \hat{L} \end{pmatrix} - \begin{pmatrix} \lambda_{TA}^{-1} & \lambda_{TM}^{-1} & \lambda_{TH}^{-1} \\ \lambda_{KA}^{-1} & \lambda_{KM}^{-1} & \lambda_{KH}^{-1} \\ \lambda_{LA}^{-1} & \lambda_{LM}^{-1} & \lambda_{LH}^{-1} \end{pmatrix} \cdot \begin{pmatrix} \lambda_{TA} a_{TA} & \lambda_{TM} a_{TM} & \lambda_{TH} a_{TH} \\ \lambda_{KA} a_{KA} & \lambda_{KM} a_{KM} & \lambda_{KH} a_{KH} \\ \lambda_{LA} a_{LA} & \lambda_{LM} a_{LM} & \lambda_{LH} a_{LH} \end{pmatrix} \quad (7)$$

If the inverse of the matrix $\{\lambda\}$ of input requirements is a Minkowski matrix, then, the Rybczynski theorem linking the sectoral outputs to changes in factor endowments can be reproduced in this 3x3 model. This being the case, then the diagonal elements is positive and the off-diagonal elements are negative. Furthermore, as the sum of the row elements of the (λ) matrix is equal to one, then the sum of the row elements of its inverse is also equal to one. This implies that the diagonal elements of the matrix are greater than one. From the system (7),

$$\frac{\hat{A}}{\hat{T}} = \lambda_{TA}^{-1} > 1, \quad \frac{\hat{A}}{\hat{K}} = \lambda_{KA}^{-1} < 0, \quad \frac{\hat{A}}{\hat{L}} = \lambda_{LA}^{-1} < 0, \quad \text{at constant commodity prices.}$$

The impact of the different variables on the agricultural output can be derived from the system of equations (7). In particular, the response of the agricultural output to changes in its own relative price can be assessed as well as those responses to changes in factor endowments. The results would depend upon the interactions among factor intensities, factor substitution and complementarities as well as on the effects of changes in commodity prices on factor prices.

The own relative price elasticity of agricultural supply can be obtained for given factor endowments ($\hat{L} = \hat{K} = \hat{T} = 0$), by introducing the zero homogeneity condition (5) and the definition of price elasticities of input-output coefficients (4) into the system of equations (7),

$$\begin{aligned} \frac{\hat{A}}{(\hat{p}_A - \hat{p}_M)} &= (-\lambda_{TA}^{-1}) \cdot \left[(\delta_T^K) \cdot \left(\frac{\hat{w}_K - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)} \right) + (\delta_T^T) \cdot \left(\frac{\hat{w}_T - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)} \right) \right] + \\ &+ (-\lambda_{TM}^{-1}) \cdot \left[(\delta_K^K) \cdot \left(\frac{\hat{w}_K - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)} \right) + (\delta_K^T) \cdot \left(\frac{\hat{w}_T - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)} \right) \right] + \\ &+ (-\lambda_{TH}^{-1}) \cdot \left[(\delta_L^K) \cdot \left(\frac{\hat{w}_K - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)} \right) + (\delta_L^T) \cdot \left(\frac{\hat{w}_T - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)} \right) \right] \end{aligned} \quad (8)$$

where $\delta_i^z = \sum_j \lambda_{ij} E_{ij}^z$ for $z = K, T$, $i = K, T, L$, and $j = A, M, H$

In equation (8), the terms $\{\delta_i^z\}$ show the response of the use of capital and land by the three sectors to changes in relative prices of capital and land compared to that of labor. The terms $\{\lambda_{Tj}^{-1}\}$ show the response of the output of the A sector that is needed

(along with those of the other two sectors) to achieve full employment of capital, labor and land. The signs and magnitudes of the terms in brackets depend upon the effects of the change in the relative price of agriculture on the prices of factors of production as well as on the interactions among factor intensities and substitution and complementarity relationships among factors¹. Furthermore, the impact of a change in p_A is going to be reflected in p_H as shown in L. Sjaastad (1980). If substitution in consumption as well as in production is assumed, then an increase in the price of A will increase the price of H but less than proportionately.

In order to obtain a relationship between relative factor prices and prices of goods, the system of equations (2) is totally differentiated to get,

$$\begin{pmatrix} \theta_{TA} & \theta_{KA} & \theta_{LA} \\ \theta_{TM} & \theta_{KM} & \theta_{LM} \\ \theta_{TH} & \theta_{KH} & \theta_{LH} \end{pmatrix} \cdot \begin{pmatrix} \hat{w}_T \\ \hat{w}_K \\ \hat{w}_L \end{pmatrix} = \begin{pmatrix} \hat{p}_A \\ \hat{p}_M \\ \hat{p}_H \end{pmatrix} \quad (9)$$

where θ_{ij} is the share of input i in production costs of sector j .

To arrive at equation (9), the envelope property of cost minimization has been used, that is,

$$\sum_i \theta_{ij} \hat{a}_{ij} = 0 \quad (10)$$

As indicated earlier, the price of the non traded good is endogenously determined by its domestic demand and supply and is affected only indirectly through substitution and complementarity effects by the prices of the two traded goods A and M . In particular, if substitution effects in both consumption and production prevail, it can be shown that,

$$1 > \frac{\hat{p}_h}{\hat{p}_A} > 0 \quad \text{and} \quad 1 > \frac{\hat{p}_H}{\hat{p}_M} > 0, \quad \text{or that}$$

$$\hat{p}_H = (1 - \omega) \cdot \hat{p}_A + \omega \cdot \hat{p}_M \quad (11)$$

where $1 > \omega > 0$. By the zero homogeneity condition in prices of both demand and supply functions of the non traded good, a proportional increase in both prices of traded goods will leave the relative price of the non traded good constant vis-à-vis prices of the two traded goods. Thus, factor prices can be obtained from (9) and (11) as functions of p_A and p_M ,

$$\begin{pmatrix} \hat{w}_T \\ \hat{w}_K \\ \hat{w}_L \end{pmatrix} = \begin{pmatrix} \theta_{TA}^{-1} & \theta_{KA}^{-1} & \theta_{LA}^{-1} \\ \theta_{TM}^{-1} & \theta_{KM}^{-1} & \theta_{LM}^{-1} \\ \theta_{TH}^{-1} & \theta_{KH}^{-1} & \theta_{LH}^{-1} \end{pmatrix} \cdot \begin{pmatrix} \hat{p}_A \\ \hat{p}_M \\ (1 - \omega) \cdot \hat{p}_A + \omega \cdot \hat{p}_M \end{pmatrix} \quad (12)$$

¹ The percentage response of A 's output to a percentage change in p_A is a general equilibrium supply price elasticity calculated along the transformation schedule of the economy. See, R. Jones (1965).

where the matrix $\{\theta^{-1}\}$ is a Minkowski matrix. Thus,

$$(\hat{w}_T - \hat{w}_L) = [(\theta_{KH}^{-1} - \theta_{KA}^{-1}) + (\theta_{LH}^{-1} - \theta_{LA}^{-1}) \cdot \omega] \cdot (\hat{p}_A - \hat{p}_M) \quad (13)$$

$$(\hat{w}_K - \hat{w}_L) = [(\theta_{KH}^{-1} - \theta_{KM}^{-1}) + (\theta_{LH}^{-1} - \theta_{LM}^{-1}) \cdot \omega] \cdot (\hat{p}_A - \hat{p}_M) \quad (14)$$

The impact of changes in the price of agriculture on the relative prices of land and capital is ambiguous. The sign of the first term in parenthesis of equation (13) depends upon the intensity of use of capital in the agricultural and non traded sectors and is negative under the factor intensity assumptions made earlier. The second term is positive and greater than one as the non traded sector is intensive in the use of labor but it is multiplied by the factor ω that is less than one. The first term in parenthesis in equation (14) is negative and the second term is positive and greater than one but it is multiplied by ω . Thus, one plausible result is that an increase in the price of agriculture would increase the relative price of land and reduce the relative price of capital compared to that of labor as agriculture and the non traded good are intensive in the use of land and labor, respectively.

In the case that there were positive and negative responses of the prices of land and capital (relative to the price of labor) to an increase in the relative price of the agricultural sector (compared to the price of import substitutes), sufficient but not necessary conditions for a positive response of the agricultural output to changes in its own relative price can be obtained as follows: (i) the non traded good uses land less intensively than the other two sectors, and (ii) either the three factors of production are substitutes ($E_{ij}^k > 0$) and the degree of substitution between capital and labor is poorer than that between land and capital in the three sectors, capital and labor are complements ($E_{Lj}^K < 0$), or capital and labor are used in fixed proportions ($E_{Lj}^K = 0$).

The impact of changes in factor endowments on agricultural output can be analyzed with equations (7) and (12). An increase in the endowment of one factor would increase production of the good that is intensive in the use of the factor y reduce that of the other two goods at constant sectoral relative prices. But prices of traded goods compared to those of non traded goods would not be constant because equilibrium in the markets for non traded goods would require of a price adjustment. This change in relative prices would result, in turn, in a change in relative factor prices and, hence, in a change in the optimal factor intensities in the three activities.

If, as assumed for expositional purposes in this paper, the non traded good sector is intensive in the use of labor, then, an increase in the endowment of labor will increase its sectoral output and will decrease those of the other two sectors under constant factor prices. The proportional increase in output of the non traded activity will be greater than the proportional increase in labor endowment in order to keep all factors fully employed. At the same time, the increase in the endowment of labor will raise national income and, as a result, the demand for non-traded goods will increase at constant commodity prices. The final impact of these two forces on relative prices of non-traded goods is ambiguous depending upon their relative magnitudes.

If, for simplicity, homothetic preferences are assumed and, consequently, unitary income demand elasticities for the three goods, then the percentage increase in national income (\hat{Y}) and in the demand for the non-traded good (\hat{H}^d) will be equal to $\hat{H}^d = \hat{Y} = \theta_l \hat{L}$ at constant commodity prices, where θ_l is the share of labor income in national income and \hat{L} is the percentage increase in the endowment of labor. On the other hand, the supply of the non-traded commodity will increase proportionally more than the increase in labor endowment at constant commodity prices according to the Rybczynski theorem. The proportional increase in output of the non-traded activity (Y_h^s) is given by $\hat{H} = \lambda_{LM}^{-1} \cdot \hat{L}$, and, according to the assumptions about sectoral labor intensities made earlier, $\lambda_{LM}^{-1} > 1$. The other two outputs will decline in response to an increase in labor supply at constant commodity prices as $\lambda_{LM}^{-1} < 0$ and $\lambda_{LA}^{-1} < 0$.

As a result, the price of the non-traded commodity will decline compared to those of exportable and importable goods to eliminate the excess supply in the market. The decline in the relative price of non-traded goods will, in turn, reduce the relative price of the factor that the activity uses intensively, namely, labor². That is, prices of capital and land will rise in relation to the price of labor and these increases will reduce the optimal capital and land-labor ratios in the three sectors. In particular, as the capital-labor ratio falls, then the output of the capital intensive good will have to increase in order to keep full employment of the existing stock of capital absorbing at the same time resources from the other two sectors. A reduction in the land-labor price ratio, on the other hand, will need of an increase in the output of the agricultural sector in order to keep land fully employed. Thus, the net effect of an increase in the amount of labor on the output of agriculture is ambiguous. The same ambiguity in the response of agriculture to changes in the stock of capital and land can be found from the systems (7) and (12).

So far, the coefficients λ_{ij} and θ_{ij} have been assumed constant under a given technology environment. If an exportable sector supply function is going to be estimated with long-run time series, then a general equilibrium framework within which this response is analyzed should to incorporate changes in these coefficients because of changes in technology.

There are two approaches in the literature to deal with the issue of technological change on resources allocation among the different activities in the economy³. One approach is to assume that changes in techniques of production are exogenously given and that the new more productive techniques totally replace the old ones for a given endowment of resources as it saves resources needed to produce the same volume of output. An alternative approach is that the adoption of new techniques is a matter of cost-benefit analysis undertaken by firms.

According to the technique choice framework, new technology might be available to firms but the cost of implementing them might be greater than the benefits. In particular, if

² It can be easily shown from the system (12) that $\hat{w}_L = \theta_{LH}^{-1} \cdot \hat{p}_H$, $\hat{w}_K = \theta_{LM}^{-1} \cdot \hat{p}_H$, and $\hat{w}_T = \theta_{LA}^{-1} \cdot \hat{p}_H$, under constant prices of agricultural and imports substitution activities.

³ For a summary of the approaches, see Y. Mundlak (2000).

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. Thus, this approach emphasizes the differences between available technology and implemented production techniques. 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 to implement new techniques include relative prices and the capital-labor and land-labor ratios. If there were an increase in the relative price of exportable agriculture resulting, for instance, from a reduction in import tariffs, and if this would result in increases in the prices of labor and land compared to that of capital. As a result, labor- and land-saving techniques were going to be adopted. If, on the other hand, there were an increase in the overall capital-labor ratio and, as a result, prices of labor compared to that of capital and land would rise, then firms would find more profitable to adopt techniques that labor-saving.

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. This analysis suggests specifying a technique choice function (T_i) for firms in sector i as,

$$T_i = f^i \left(\frac{P_x}{P_m}, K, T, L \right) \quad (15)$$

The results of this section can be summarized in an aggregate agricultural supply function like the following:

$$\ln A = \beta_0 + \beta_1 \ln \left(\frac{P_x}{P_m} \right) + \beta_2 \ln K + \beta_3 \ln T + \beta_4 \ln L \quad (16)$$

where the β 's include not only the direct effect of the variables on agricultural output but also their impact on the rate of adoption of technology given the available state of the arts.

2. The Results of the Estimation

In this section, the agricultural supply function (16) is estimated with annual data covering the period 1939-1984 for which data is available⁴. On the basis of the Phillips-Perron test, the null hypothesis that all the variables of the agricultural supply function are non stationary cannot be rejected by the data (see Table 1 below). The p -values are obtained from the surface response function estimated by MacKinnon (1993) that permits

⁴ For a description of the data, see Annex B.

the calculation of the Phillips and Perron critical values for any sample size and for any right-hand variables.

Table 1: Phillips-Perron (PP) Test for Unit Roots

Variables	PP statistic	p-values
ln(A)	-1.31	0.62
ln(ρ_x/ρ_m)	-0.68	0.42
ln(K)	1.55	0.99
ln(L)	-1.25	0.64
ln(T)	0.70	0.99

Notes: With the exception of the logarithm of the relative price of agriculture, all the variables show a trend in their levels and a deterministic constant was thus added to the first differences of the variables in the PP unit root tests.

Non stationary variables mean that a linear combination of them may be stationary and this is what has been referred in the literature as cointegration vectors that are interpreted as long run equilibrium relationships. One method of estimating the long run parameters of the agricultural supply function is that of Johansen and Juselius (1990). In order to estimate the number of stationary long run equilibrium equations, a cointegration test has to be performed. Consider first the following autoregressive model

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \dots + \Pi_k X_{t-k} + \varepsilon_t \quad (t=1, \dots, T) \quad (17)$$

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.

As most economic time series are non stationary, it is convenient to rewrite the model (17) as:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \dots + \Gamma_{t-k+1} + \Pi X_{t-k} + \varepsilon_t \quad (18)$$

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

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

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

As demonstrated by Johansen and Juselius (1990), testing the number of cointegration vectors amounts to determine the rank of the matrix π . The hypothesis that the rank of π is r can be formulated as the restriction that $\pi = \alpha\beta'$ where the vector β is the cointegrating vector with the property that $\beta'X$ is stationary, and α can be interpreted as the average rate of adjustment of the variables towards their long run equilibrium values, and

α and β are $p \times r$ vectors. If the hypothesis that $r=0$ is rejected, then the matrix π contains information about long-run relationships between the variables in the data.

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 (17).

As demonstrated by Cheung and Lai (1993), the critical asymptotic values of the two tests statistics tend to overestimate the number of statistically significant cointegration vectors in small samples. They also find that the trace test statistic shows little bias in the presence of either skewness or excess kurtosis and that the maximal eigenvalue test shows substantial bias in the presence of large skewness although it is quite robust to excess kurtosis.

Cheung and Lai (1993), and Ahn and Reinsel (1990) proposed small sample corrections based on the degrees of freedom. While Cheung and Lai (1993) correct the asymptotic critical values of the test statistics, Ahn and Reinsel (1990) correct the test statistics. Cheung and Lai show that the Ahn-Reinsel method does not yield unbiased estimates of the finite sample critical values for Johansen's tests. The distribution of the test statistics in small samples depends not only on the degrees of freedom but also on the parameters of the vector autoregression under the null hypothesis about the number of cointegrated vectors. Both of the mentioned degrees of freedom corrections capture part of this dependence on the lag length but not the dependence upon the parameters.

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 $(\alpha, \beta, \hat{\Gamma}_i, \text{ and } \hat{\Omega})$ under the null hypothesis about the number of cointegration vectors (r) and 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 $(\alpha, \beta, \hat{\Gamma}_1, \dots, \hat{\Gamma}_n, \text{ and } \hat{\Omega})$.

The unrestricted parameters of the vector autoregression (17) 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. In addition, the White test statistic indicates that the hypothesis of homoskedastic disturbances cannot be rejected with a marginal probability of about 20 percent.

The test of the null hypothesis of Gaussian residuals is based on the Jarque-Bera test statistic with a correction for small samples as proposed by Doornik and Hansen (1994) for both the individual equations and the system of equations as a whole. 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 (as with the Choleski orthogonalization. Furthermore, the Doornik and Hansen's procedure transforms skewness and kurtosis to approximately χ^2 in small samples.

For the system as a whole, the null hypothesis of normality cannot be rejected by the data as the χ^2 (10) test statistic is calculated for the system as a whole in 4.09 with a marginal significance level (the p -value) of 94 percent. The null hypothesis of serially uncorrelated residuals is also tested and the Lagrange multiplier test statistic indicates that it cannot be rejected by the data at one and two lags of the residuals with p -values of 35 and 44 percent, respectively. There is no evidence of heteroskedastic disturbances according to the ARCH test⁵.

In order to calculate the small sample correction factor of the trace statistic under the null hypothesis of $r=0$, the model (18) is fitted with one lag of the variables in first differences and a constant in the cointegration space and a linear trend in the data⁶. Since the power of the test is likely to be low against near cointegration alternatives, it seems reasonable to follow a test procedure that rejects the null hypothesis for higher p -values than the usual 5 percent significance level, say, at 10 percent.

The corrected trace statistic of 72.80 shows that the hypothesis of no cointegration vector ($r=0$) is rejected by the data (see table 2 below) with a marginal probability of 2.7 percent. The hypothesis of one cointegration vector ($r=1$) cannot be rejected as the corrected trace statistic is calculated in 42.93 with a marginal probability of 13.4 percent.

Table 2: Trace Test Statistics for Testing Cointegrating Vectors

r	Bartlett Correction Factor	Trace Statistic	Corrected Trace Statistic	p - values
0	1.388	101.051	72.802	0.027
1	1.237	53.117	42.931	0.134

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 Γ -distribution, see Doornik (1998).

The estimated cointegration vector (the β_1 's) is as follows (the t -statistics are the numbers in parenthesis):

$$\ln(A) = 46.055 + 1.256 \ln(p_x/p_m) + 2.226 \ln(K) + 4.679 \ln(T) - 6.701 \ln(L)$$

(7.970) (6.466) (4.014) (-7.541)

All the estimated coefficients are significantly different from zero at the usual significance levels. In particular, the long-run supply price elasticity is estimated in 1.26 and is about three times higher than those estimated by Reca (1980). The estimated

⁵ See Annex A

⁶ The coefficients to calculate the correction factor have not been tabulated in Johansen, Nielsen, and Fachin (2005). This problem is avoided by using the coefficients of a slightly larger model with a linear trend restricted to the cointegration space.

supply price elasticity is higher than the simulated by Mundlak, Cavallo, and Domenech (1989) with a time horizon of twenty years. The economic interpretation of the coefficients of the factor endowments is not straightforward because they include not only the effect of their changes on agricultural output but also the effects of these changes on relative prices of non-traded goods. In addition, they may include the impact of these variables on the rate of adoption of new technologies.

The average speeds of adjustment of the variables towards their long-run equilibrium states (the α vector) are estimated as follows (the number in parenthesis are the t-statistics):

Table 4: Estimates of the α vector

$\ln A$	$\ln(p_x/p_m)$	$\ln K$	$\ln T$	$\ln L$
-0.114 (-3.662)	0.300 (4.587)	0.007 (1.452)	0.050 (3.079)	-0.011 (-1.035)

The average speed of adjustment of the agricultural output is -0.11 and this indicates that convergence to its long-run equilibrium state is achieved in about nine-ten years. The likelihood ratio test statistic (adjusted by degrees of freedom) under the null hypothesis that this coefficient is zero is calculated in 4.44 and this hypothesis is rejected with a marginal probability of 3.5 percent. In their simulations, Mundlak, Cavallo, and Domenech (1989) implicitly estimate a unit elastic response of agricultural output to changes in relative prices but the long run equilibrium output is achieved in twenty years⁷.

Weakly exogeneity tests of the variables can be conducted under the null hypothesis of $\alpha_i=0$ using the likelihood ratio test adjusted by degrees of freedom that is distributed as $\chi^2(1)$. The value of the test statistic under the null hypothesis that the relative price of agriculture (P_x/P_m) is weakly exogenous is 5.14 with a p -value of 2.3 percent and this amounts to reject it at the usual significance levels. On the other hand, the test statistic calculated under the null hypothesis that land is weakly exogenous is 4.85 and this value amounts to reject it with a p -value of about 2.8 percent. Furthermore, the hypothesis that capital and labor are weakly exogenous cannot be rejected by the data.

The interpretation of causal relationships among variables is not straightforward. For instance, a government policy reaction function can induce causality from output to prices although they are structurally exogenous and simulations based on the estimated parameters can be subject to the Lucas' criticism (1981). In the case of the relative price variable, a plausible explanation for the lack of weakly exogeneity could be a feedback policy reaction that relates taxation of exports to a booming agricultural economy. An alternative hypothesis is that the international prices of agricultural commodities are affected by Argentinean agricultural output variations in which case the country would not be a small economy in international markets. This is certainly an issue that deserves additional research because if the country is a price maker in world markets, then an

⁷ See Mundlak, Y., D. Cavallo, and R. Domenech (1989), Table 19, page 96.

export tax would maximize welfare from the country's standpoint in absence of import taxes.

As mentioned earlier, the hypothesis that land is weakly exogenous is rejected by the data. One explanation could be that if the error correction vector contains, in addition to its own equilibrium correction process, information about stationary technological shocks that are land saving and if these shocks affect the use of land by the various agricultural crops differently, then this could be reflected in the output weighted cultivated land and a causal ordering would follow.

Mundlak, Cavallo, and Domenech (1989) include several macroeconomic variables such as inflation, deflation, and bank failures in their estimates of agricultural production functions. A model consisting of the six variables including fiscal deficit (d as a percentage of GDP⁸) is fitted with two lags, and an unrestricted constant term. Fiscal deficits can have negative effects on sectoral outputs through potential distortions caused by the inflation tax and/or crowding out effects of public sector borrowing over the private sector.

The results of the estimation indicate that the hypothesis of normally distributed residuals cannot be rejected as the multivariate Jarque-Bera test statistic ($\chi^2(12)$) is calculated in 8.46 with a marginal probability of 75 percent. On the other hand, the LM statistics to test the hypothesis of uncorrelated residuals are calculated at one and two lags in 34.92 and 33.56, respectively, and these values indicate that the null hypothesis cannot be rejected. On the other hand, the hypothesis of homoskedastic disturbances is not rejected by the data as the multivariate ARCH test statistics are calculated in 454.0 and 924.0 for one and two lags, respectively with marginal probabilities of 32 and 16 percent.

The cointegration rank of the cointegrating model is determined by comparing the trace statistic adjusted by the Bartlett small sample correction factor with the asymptotic critical values at 10 percent significance level. The results indicate (see table 3) that the hypothesis of $r=1$ cannot be rejected at this significance level.

Table 3: Trace Test Statistics for Testing Cointegrating Vectors

r	Bartlett Correction Factor	Trace Statistic	Corrected Trace Statistic	p -values
0	1.4137	129.889	91.173	0.098
1	1.2790	77.805	60.832	0.211

Note: See Table 2.

The estimated cointegrating vector β is as follows:

$$\ln(A) = 33.224 + 1.055 \ln(p_x/p_m) + 2.578 \ln(K) + 3.137 \ln(T) - 6.547 \ln(L) - 0.023 \ln(d)$$

(7.810)
(6.433)
(2.925)
(-7.873)
(-2.478)

⁸ The null hypothesis that the fiscal deficit follows a unit root autoregressive process with a linear trend in the level of the variable cannot be rejected by the data as the PP test statistic is calculated in -2.63 and this amounts not to reject the null with a marginal probability of 9.5 percent.

The supply price elasticity is estimated in about 1.0 and this estimate is close to that of the supply price function that excludes the fiscal deficit. This estimated supply price elasticity is similar to that estimated in Mundlak, Cavallo, and Domenech (1989) for a time horizon of twenty years. As expected, the coefficient of the fiscal deficit is negative and statistically different from zero. The remaining coefficients are all statistically different from zero and they are quite robust to the inclusion of the fiscal deficit variable.

The speed of convergence of agricultural output towards the long-run equilibrium is estimated in -0.134 and is statistically different from zero. This is another piece of evidence supporting the existence of one cointegration vector. The hypothesis that relative prices and land are weakly exogenous are rejected by the data as the likelihood ratio tests statistics are calculated in 6.6 and 4.1, respectively with marginal probabilities of 1.0 and 4.2 percent.

In his aforementioned study, L. G. Reca (1980) includes a dummy variable that takes a value of zero for the years 1950-1964 and one afterwards to define two different technological levels. According to Reca's estimates, technological change was responsible for an increase in agricultural production of about 8-10 percent between 1965 and 1974. As the coefficients of the supply function estimated in this paper includes the effect of prices changes on the adoption of new production techniques, the existence of two technological levels before and after 1964 could affect the stability of the parameters over time.

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 and non-constancy of these parameters will be reflected in the time path of the estimated eigenvalues. The time paths of the estimated eigenvalues for the sub sample 1967-1984 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 model⁹. The results are presented in Figure 1 and, although it is not a formal test of stability of parameters, they do not seem to indicate non-constancy of the parameters.

A formal test of stability of parameters over time developed by Hansen and Johansen (1999) is presented in Figure 2, where the fluctuation test of the transformed eigenvalues is presented. The eigenvalues, λ_i , are transformed into $\varepsilon_i = \log\left(\frac{\lambda_i}{1-\lambda_i}\right)$ to obtain a better approximation of their limiting distribution. 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 α and β and concentrating out the short term coefficients (the R1-form). It can be seen that the values of the test statistics are below the 5% critical level and, consequently, the hypothesis of constancy of parameters cannot be rejected.

⁹ In the six variable system including the fiscal deficit, the sub sample to calculate the recursive eigenvalues includes the period of time from 1975 until 1984 and, thus, the estimates of the coefficients with the base data could be affected by the technological innovations of the mid-sixties and, consequently, the stability tests might not be reliable. For this reason, the eigenvalues of the six variable system have not been recursively estimated.

Eigenvalues

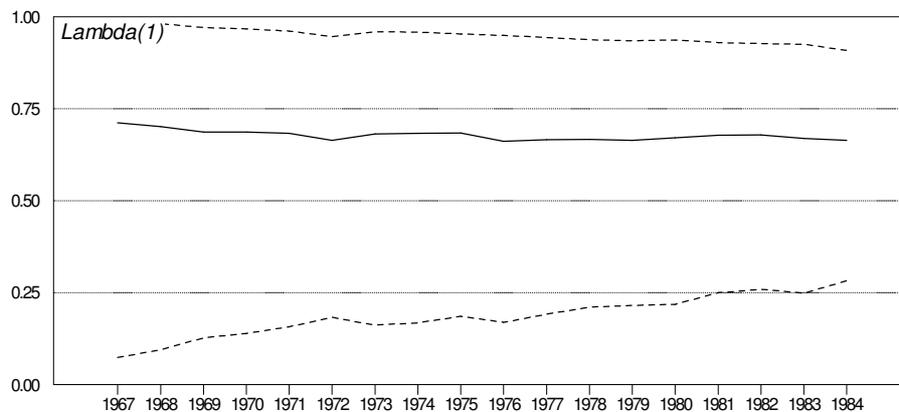
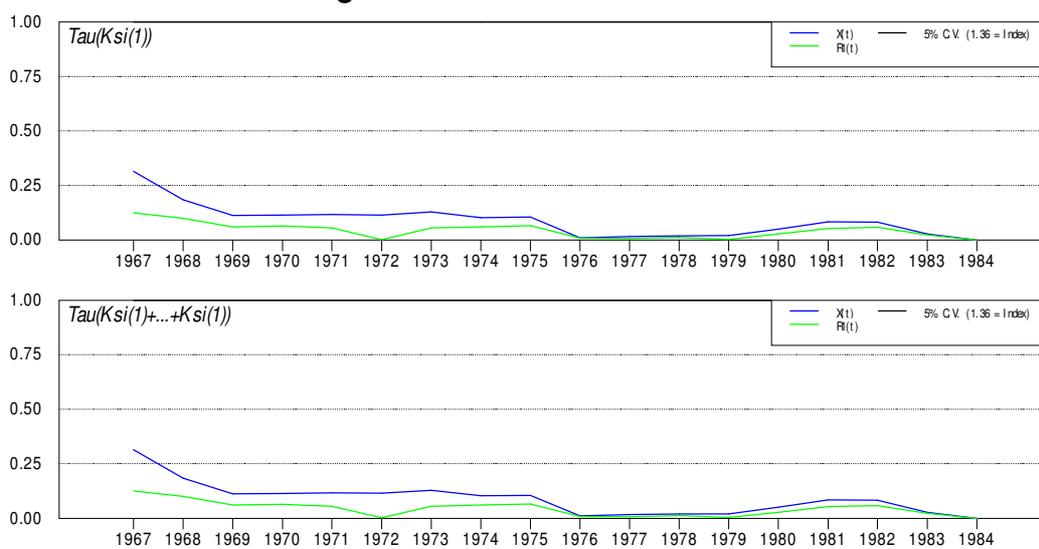


Figure 1: Recursive estimates of one non-zero eigenvalue (black solid line) with asymptotic 95% confidence bands (dotted lines), 1967-1984.

Eigenvalue Fluctuation Test



$$\text{Tau}(Ksi) = C(T) \|Ksi(t) - Ksi(T)\|$$

Figure 2: The blue and green lines correspond to the X - and $R1$ forms of the test statistics. The graphs are scaled by the 5% critical value (1.36). The graph at the bottom corresponds to the sum of the transformed eigenvalues and it is equal to the graph of the top because there is only one cointegrating vector.

3. Concluding Remarks

This paper estimates a reduced-form agricultural supply function within the framework of a general equilibrium model. The supply function includes not only sectoral relative prices but also the stock of factors of production. The response of agricultural output to changes in factor endowments reflect not only changes in the transformation surface of the economy caused by these changes but also the movement along it resulting from the induced changes in relative factor prices. The model is also estimated with the inclusion of a fiscal deficit variable to incorporate the effects of macroeconomic policies on agricultural output.

Technological change is regarded as a variable subject to economic choice by firms given the available technology and, hence, it is endogenous. Variables that affect this choice of techniques include relative sectoral prices and factor endowments. In addition, the endowment of physical capital (and, perhaps, land) is regarded as a carrier of the available technology as it is correlated with human capital and, hence, with knowledge.

The results of the estimations show that there is a statistically significant supply price elasticity of the agricultural output in the range of 1.0-1.25 and that the convergence to the equilibrium state is achieved in a range of about eight to ten years, depending upon the model estimated. The estimated supply price response is much higher than previous estimates of partial equilibrium models by Reca (1969, 1980), and Colome (1966). The speed of convergence is much shorter than that estimated by Mundlak, Cavallo, and Domenech (1989) of about twenty years for a unitary response of agricultural output to changes in prices. It is also found that fiscal deficits have a statistically significant negative effect on agricultural output.

The results also show that, according to the time path of one non-zero eigenvalue, there is no indication of non-constancy of parameters for the period covering 1967-1984, years during which new techniques of production were adopted. A formal fluctuation test of the transformed eigenvalues also indicates that the null hypothesis of constancy of parameters cannot be rejected by the data.

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Annex A: Univariate Tests of Residuals

Normality test – Doornik-Hansen procedure:

Equation	<i>J-B</i> Test Statistic	<i>p</i> -value
ln <i>A</i>	0.331	0.848
ln(<i>P_x/P_m</i>)	3.193	0.203
ln <i>K</i>	0.203	0.903
ln <i>T</i>	2.609	0.271
ln <i>L</i>	0.745	0.689

ARCH test of heteroskedasticity:

Equation	ARCH(2)	<i>p</i> -value
ln <i>A</i>	0.038	0.981
ln(<i>P_x/P_m</i>)	0.149	0.928
ln <i>K</i>	1.068	0.586
ln <i>T</i>	2.963	0.227
ln <i>L</i>	0.113	0.945

Normality test – Doornik-Hansen procedure:

Equation	<i>J-B</i> Test Statistic	<i>p</i> -value
ln <i>A</i>	0.602	0.740
ln(<i>P_x/P_m</i>)	0.616	0.735
ln <i>K</i>	0.079	0.961
ln <i>T</i>	1.350	0.509
ln <i>L</i>	1.320	0.517
<i>d</i>	4.034	0.133

ARCH test of heteroskedasticity:

Equation	ARCH(2)	p -value
$\ln A$	1.403	0.496
$\ln(P_x/P_m)$	0.366	0.833
$\ln K$	1.966	0.374
$\ln T$	1.457	0.483
$\ln L$	1.460	0.482
d	0.159	0.924

Annex B: Data Description

A: Agricultural output at 1960 prices. Source: IERRAL, op. cit.

(P_x/P_m) : Price of agricultural goods divided by the 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: Stock of capital employed in production of goods and services in australes at 1960 prices. Source: IEERAL, op. cit.

T: Total planted area 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.

D: Consolidated fiscal deficit as percentage of the gross domestic product at current prices. Source: IERAL, op. cit.