Supply Response of Agriculture in the Argentinean Economy

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In the paper by A. Herrou-Aragón (2001), it was found that the tax on the volume of international trade in the Argentinean economy was about 110 percent in the second half of the 80s and about 55 percent during the 90s compared to that of the second half of the 30s. The commercial policies leading to the heavy taxation of international trade during the second half of the 80s have certainly harmed, to some extent, the growth potential of the Argentinean exportable sector during those years. This paper is aimed at estimating the extent to which taxation of international trade has had an impact on the Argentinean agricultural sector during the period from 1939 until 1984 for which the needed data is available.

So far, the most comprehensive 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 adoption of new techniques is endogenous 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 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. The results of the simulations indicate a unitary price elasticity of aggregate agricultural supply but it takes about twenty years to achieve the output equilibrium state.

In support of Mundlak´s hypothesis of endogenous rate of adoption of new techniques, Herrou-Aragón (2004) finds a positive relationship between the relative price of labor compared to the return to capital, and total factor productivity of the agricultural sector using Johansen´s cointegration methodology because the test about non stationary variables fails to reject the null hypothesis.

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,

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

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,

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

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,

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

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,

$$E_{ij}^z = \frac{\partial a_{ij}}{\partial w_z} \cdot \frac{w_z}{a_{ij}} \qquad (4)$$

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

$$\sum_z E_{ij}^z = 0 \qquad (5)$$

The exogenous variables of the system of equations (1)-(5) are the three factor endowments $(L, T, \text{and } K)$, 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,

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

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), $\sum_j \hat{\alpha}_{Tj}$, 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),

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

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 are 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),

$\frac{\hat{A}}{\hat{T}} = \lambda_{TA}^{-1} > 1$, $\frac{\hat{A}}{\hat{K}} = \lambda_{KA}^{-1} < 0$, $\frac{\hat{A}}{\hat{L}} = \lambda_{LA}^{-1} < 0$, 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),

$$\frac{\hat{A}}{(\hat{p}_A - \hat{p}_M)} = \left(-\lambda_{TA}^{-1}\right)\cdot\left[\left(\lambda_{TA}E_{TA}^{K} + \lambda_{TM}E_{TM}^{K} + \lambda_{TH}E_{TH}^{K}\right)\cdot\left(\frac{\hat{w}_K - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)}\right) + \left(\lambda_{TA}E_{TA}^{T} + \lambda_{TM}E_{TM}^{T} + \lambda_{TH}E_{TH}^{T}\right)\cdot\left(\frac{\hat{w}_T -}{(\hat{p}_A -}\right.\right.$$

$$+\left(-\lambda_{TM}^{-1}\right)\cdot\left[\left(\lambda_{KA}E_{KA}^{K} + \lambda_{KM}E_{KM}^{K} + \lambda_{KH}E_{KH}^{K}\right)\cdot\left(\frac{\hat{w}_K - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)}\right) + \left(\lambda_{KA}E_{KA}^{T} + \lambda_{KM}E_{KM}^{T} + \lambda_{KH}E_{KH}^{T}\right)\cdot\left(\frac{\hat{w}_T - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)}\right)\right] +$$

$$\left(-\lambda_{TH}^{-1}\right)\cdot\left[\left(\lambda_{LA}E_{LA}^{K} + \lambda_{LM}E_{LM}^{K} + \lambda_{LH}E_{LH}^{K}\right)\cdot\left(\frac{\hat{w}_K - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)}\right) + \left(\lambda_{LA}E_{LA}^{T} + \lambda_{LM}E_{LM}^{T} + \lambda_{LH}E_{LH}^{T}\right)\cdot\left(\frac{\hat{w}_T - \hat{w}_L}{(\hat{p}_A - \hat{p}_M)}\right)\right] \quad (8)$$

In equation (8), the terms in brackets 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 $\left\{\lambda_{Tj}^{-1}\right\}$ show the needed response of the output of the *A* sector 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 factors[1]. 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,

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

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

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,

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

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

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

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

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 less than one in absolute value 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 *ù* 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 *ù*. 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.

If there is a positive response of the price of land and a negative one of the price of capital compared to that of labor, then, a positive response of the agricultural sector to changes in its own relative price can be obtained under the following conditions: (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 $\left(E_{Lj}^K < 0\right)$, or capital and labor are used in fixed proportions $\left(E_{Lj}^K = 0\right)$.

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 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 commodity 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 $\left(\hat{H}^d\right)$ 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 $\hat{H} = \lambda_{LM}^{-1} . \hat{L}$, 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, labor[2]. 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, if 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 economy[3]. 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 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 and land compared to that of capital would rise, then firms would find more profitable to adopt techniques that are more intensive in the use of capital.

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 physical capital and, perhaps, of land, these other production 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,

$$T_i = f^i\left(\frac{p_x}{p_m},\ K, T, L\right) \qquad (15)$$

So far, the role of government (*G*) in the economy has been omitted. If government is thought as an agency that purchase inputs from the private sector and produces goods and services that are mostly non traded goods, then, the impact of government activities on agricultural output can be assessed. Government activities may have positive effects on the production possibilities of the economy as they can generate externalities to other sectors (security, defense, and enforcement of private contracts and law). On the other hand, if the government is labor intensive, then an increase in the size of the government would tend to reduce the output of non-traded goods that are privately produced. This, in

turn, would increase prices of these goods and that of labor compared to the prices of the other factors. As was mentioned earlier, the net impact of these changes on resources allocation is ambiguous. It has also to be mentioned that if the government does not follow the profit maximizing rules of the private sector, then the marginal rates of substitution among factors of production are not going to be equalized across public and private sectors of the economy. As a result, the economy would be forced to operate inside its production possibilities frontier. Thus, the impact of government activities on resources allocation is not predictable on a *priori* basis.

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

$$\ln A = \beta_0 + \beta_1 \ln\left(\frac{p_x}{p_m}\right) + \beta_2 \ln K + \beta_3 \ln T + \beta_4 \ln L + \beta_5 \ln G \qquad (16)$$

where the $\hat{a}$'s include not only the direct effect of the variables on agricultural output but also their impact on the rate of adoption of technology.

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 available[4]. On the basis of the Phillips-Perron test[5], the null hypothesis that all the variables of the agricultural supply function are non stationary cannot be rejected by the data (see table1 below). The *p*-values are obtained from the surface response function estimated by MacKinnon (1993) that permits 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(p_x/p_m)	-0.68	0.42
ln(*K*)	1.55	0.99
ln(*L*)	-1.03	0.73
ln(*T*)	0.70	0.99
ln(*G*)	-2.37	0.15

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 can be 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:

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

where ε_t´s are independent Gaussian variables with 0 mean and variance , and X_t is a px1 vector of stochastic variables.

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

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

where Ä=1-L, and L is the lag operator

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

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

As demonstrated by Johansen and Juselius (1990), testing the number of cointegration vectors amounts to determine the rank of the matrix ð. The hypothesis that the rank of ð is r can be formulated as the restriction that ð=áâ' where the vector â is the cointegrating vector with the property that â'X is stationary, and á can be interpreted as the average rate of adjustment of the variables towards their long run equilibrium values, and á and â are p x r vectors. If the hypothesis that r=0 is rejected, then the matrix ð 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 ð matrix, namely, the eigenvalue and the trace statistics. Asymptotic critical values for these test statistics are provided by Johansen and Juselius (1990) and also by Osterwald-Lenum (1992) and Johansen and Nielsen (1993). 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) proposes a small sample correction of the trace test statistic by finding 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 $(\hat{a}, \hat{b}, \hat{\Gamma}_i, \text{and } \hat{\Omega})$ 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 $(\hat{a}, \hat{b}, \hat{\Gamma}_1, ..., \hat{\Gamma}_n, \text{and } \hat{\Omega})$.

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-Vera test statistic for both the individual equations and the system of equations as a whole. The results show (see Table 2) that the null hypothesis of normally distributed residuals cannot be rejected. 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 four lags of the residuals with *p*-values of 0.27 and 0.37 percent, respectively.

Table 2: Jarque-Vera Tests Statistics

Component	Jarque-Vera Statistic	*p*-values
ln(*A*)	0.07	0.97
ln(p_x/p_m)	3.25	0.20
ln(*K*)	0.35	0.84
ln(*L*)	3.84	0.15
ln(*T*)	1.92	0.38
ln(*G*)	1.53	0.46
System	17.42	0.13

Note: the Jarque-Vera statistic is defined as $\left(\frac{T-m}{6}\right)\left(SK^2 + \frac{EK^2}{4}\right) \sim \chi^2(2)$, where *SK* is skewness, *EK* is excess kurtosis, *T* is the sample size, and *m* is the number of regressors. The Jarque-Vera statistic for the system as a whole is distributed as $\chi^2(12)$.

In order to calculate the small sample correction factor, the model (18) is fitted with one lag of the variables in first differences and deterministic constant inside the cointegration vectors and a linear trend in the data. The corrected trace statistic shows that the hypothesis of no cointegration vector (r=0) is rejected by the data (see table 3 below) at a p- value of 10 percent. On the other hand, the hypothesis of one cointegration vector (r=1) cannot be rejected at 10 percent significance level. 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[6].

Table 3: Trace Test Statistics for Testing Cointegrating Vectors

r	Johansen´s Correction Factor	Trace Statistic	Corrected Trace Statistic	Critical Values at 10%
0	1.3709	123.611	90.168	89.371
1	1.3198	82.665	62.634	64.742

Note: The corrected trace statistic is the trace statistic divided by the Johansen´s correction factor. The asymptotic critical values of the trace statistics are from Johansen, S., and B. Nielsen (1993).

The estimated cointegration vector (the $\hat{a}_1$´s) is as follows (the t-statistics are the numbers in parenthesis):

$$\ln(A) = \underset{}{-5.888} + \underset{(6.027)}{0.677}\ln(p_x/p_m) + \underset{(4.523)}{1.097}\ln(K) + \underset{(2.745)}{2.455}\ln(T) - \underset{(4.558)}{2.6260}\ln(L) - \underset{(2.808)}{0.5946}\ln(G)$$

All the estimated coefficients are significantly different from zero at the usual significance levels. In particular, the coefficient of the size of the government (G) is negative and similar in absolute value to that of the supply price elasticity. As the coefficient of labor is negative, this could mean that the direct impact of labor endowment on agricultural output would dominate the effects of changes in factor prices if production of non traded goods is intensive in the use of labor. On the other hand, an increase in government employment would have the opposite effects on the output of the agricultural sector and a positive coefficient should thus be expected. A plausible explanation is that the negative efficiency effect of the government sector dominates the other effects of reducing the amount of labor available to the private sector on agricultural output.

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

Table 4: Estimates of the *á* vector

ln A	ln(p_x/p_m)	ln$K$	ln$T$	ln$L$	ln$G$
-0.256	0.560	0.003	0.124	0.020	-.023
(-4.079)	(3.802)	(0.285)	(3.568)	(0.968)	(-0.875)

The average speed of adjustment of the agricultural output is equal to (-0.256) and this indicates that convergence to its long-run equilibrium state is achieved in about four years. In their simulations, Mundlak, Cavallo, and Domenech (1989) implicitly estimate a similar value of the output response to relative prices that is achieved in about fifteen years[7].

Weakly exogeneity tests of the variables can be conducted under the null hypothesis of $á_i=0$ using the likelihood ratio test adjusted by degrees of freedom that is distributed as $÷^2(1)$. The value of the test statistic under the null hypothesis that the relative price of agriculture (P_x/P_m) is weakly exogenous is (3.39) with a *p*- value of about 7 percent and this amounts not to reject it at the usual significance levels. On the other hand, the hypothesis that land is weakly exogenous in the agricultural supply function is rejected with a *p*-value of the likelihood ratio test statistic of about 2.7 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). If the critical value of the rejection region is fixed at 10 percent, then the hypothesis of weakly exogenous relative price cannot be rejected. A plausible explanation 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 then an export tax would maximize welfare from the country standpoint in absence of import taxes.

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, and a variable measuring the size of government. 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 changes in relative factor prices that they cause.

Technological change is regarded a variable subject to economic choice by firms given the available technology and, hence, 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 show that there is a statistically significant price elasticity of the agricultural output of about (0.70) and that the convergence to the equilibrium state is achieved in a relatively short period of time of about four years. This means that the negative response of the agricultural sector cannot be underestimated in the pursuing of

import substitution policies. The price elasticity is about twice as much as the one simulated by Mundlak, *et. al*. (1989), with a structural model for the same time horizon.

There is also evidence that the size of the government has significant negative effects on agricultural output and that this could reflect the lack of efficiency of the government in providing services to the private sector. Furthermore, the size of the coefficient of the relative government size is close in absolute value to that of the supply price elasticity. This would mean that the implementation of trade liberalization policies along with reforms in the public sector aimed at a more efficient provision of services to the private sector could lead to a substantial response of the agricultural output.

Footnotes

* The author is indebted to Prof. S. Johansen for his helpful suggestions. Any mistake is of the responsibility of the author.

[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).

[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).

[4] For a description of the data, see Appendix.

[5] 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.

[6] A linear trend was included in the cointegration vector as a proxy of technological change along with a trend in the data, but the hypothesis of r=0 could not be rejected at 15 percent significance level.

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

References

Batra, R. and F. Casas, "A Synthesis of the Heckscher-Ohlin and the Neoclassical Models of International Trade", Journal of International Economics, 6, 1976.

Cheung, Y. and K. Lai, "Finite-Sample Sizes of Johansen´s Likelihood Ratio Tests for Cointegration", Oxford Bulletin of Economics and Statistics, 55, 3 (1993).

Elías, V., Sources of Growth, ICS Press, 1992.

Herrou Aragón, A., "How Far Have Trade Reforms in Argentina Gone?", *Annals of the Argentinean Economic Association,* 2001.

- , "Trade Liberalization, Income Distribution and Technological Change in Argentina", *Annals of the Argentinean Economic Association*, 2004.

IEERAL, "Estadísticas de la Evolución Económica de Argentina. 1913-1984", Estudios, Año IX, Nº 39, Julio/Septiembre 1986.

Johansen, S. and K. Juselius, "Maximum Likelihood Estimation and Inference on Cointegration – With Applications to the Demand for Money", Oxford Bulletin of Economics and Statistics, 52,2 (1991).

Johansen, S. and B. Nielsen, "Asymptotics of Cointegration Rank Tests in the presence of intervention dummies", Manuscript, Institute of Mathematical Statistics, University of Copenhagen, 1993.

Johansen, S., "A Small Sample Correction for the Test of Cointegration Rank in the Vector Autoregressive Model", Econometrica, Vol. 70, Nº 5 (September, 2002).

Johansen, S., H. Hansen, and S. Fachin, "A Simulation Study of Some Functionals of Random Walk", University of Copenhagen, discussion paper with programs, revised version, January 2005.

Jones, R., "The Structure of Simple General Equilibrium Models", Journal of Political Economy, Volume LXXIII, Nº 6, December 1965.

Jones, R. and S. Easton, "Factor Intensities and Factor Substitution in General Equilibrium", Journal of International Economics, 15 (1983), North-Holland.

Lucas, R., Studies in Business Cycle Theory, The MIT Press, Cambridge, Massachusetts, 1981.

MacKinnon, J. G., "Critical Values for Cointegration Tests" in R. F. Engle and C. W. Granger (eds.), Long-run Economic Relationships: Readings in Cointegration, Oxford University press, 1991.

Maddala, G. S., Introduction to Econometrics, Third Edition, Wiley, 2000.

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

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

Reinsel, G. and S. Ahn, "Asymptotic Properties of the Likelihood Ratio Tests for Cointegration in the Nonstationary Vector AR Model", Technical Report, Department of Statistics, University of Wisnconsin, Madison, 1988.

Rivera-Batiz, F., "Non traded Goods and the Pure Theory of International Trade with Equal Numbers of Goods and Factors", International Economic Review, Vol. 23, N° 2, June 1982.

Rodriguez, C., "Salarios Reales y Protección", Cuadernos de Economía N° 54-55, 1981.

Sjaastad, L., "Commercial Policy, 'True tariffs', and Relative Prices", in John Black and Brian Hindley (eds.), Current Issues in Commercial Policy and Diplomacy, London, MacMillan, 1980.

Appendix: Data Description

(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 : Employment in the agriculture and non agriculture sectors and excludes public sector employment in million of people. Source IEERAL, op. cit.
G : Ratio of government employment to total employment. Source: IEERAL, op. cit.