International R&D Spillovers and Productivity Growth in the Agricultural Sector A Panel Cointegration Approach

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by

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First Draft

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Abstract

This paper analyses, within the new growth theory framework and using panel cointegration techniques, the effect of agricultural international technological spillovers on total factor productivity growth for a sample of 47 countries during the period 1970-1992. The analysis shows that total factor productivity is strongly influenced by domestic as well as foreign public R&D spending in agricultural sector and geographical factors matters. Countries located in temperate zones benefit more than countries located in tropical zones from technological spillovers. Finally, the analysis shows that the rate of return to agricultural R&D spending is higher in tropical countries and this could justify new support and an even greater investment of funds for agricultural R&D for these countries.

Key words **:** Technology spillover, agricultural productivity, panel cointegration.

JEL classification: C14, O30, Q16.

1. Introduction

Much research has been done in recent years to assess the importance of research and development (R&D) and trade in influencing output growth and total factor productivity. There is now a large body of literature that provide theoretical as well as empirical models where cumulative R&D is the main engine of technological progress and productivity growth (see Aghion and Howitt (1998), Grossman and Helpman (1991) and Romer (1990)). The empirical evidence has been provided by Coe and Helpman's (1995) seminal contribution where they find that accumulated spending on R&D by a country and by its trade partners helps to explain the growth of total factor productivity.

R&D investments are still central to agricultural productivity growth. Alston *et al.* (1999) in the introduction of their recent book on the theme underline that "Throughout the twentieth century improvements in agricultural productivity have been closely linked to investments in agricultural R&D and to policies that affect agricultural R&D".

Given the importance of agricultural R&D to the growth of the sector, many works have been devoted to reporting measures of the returns to domestic agricultural R&D, see recently Esposti (2000) and for a survey Alston *et al.* (2000). But in a world where the international trade of agricultural products and the dissemination of knowledge are widespread, domestic agricultural productivity depends not only on domestic R&D but also on foreign R&D efforts. This point has been fully recognised, among others, by Hayami and Ruttan (1985) where they emphasise that a country can acquire substantial gains in agricultural productivity by borrowing advanced technology existing in other countries.

Recent works by Evenson and Singh (1997), Schimmelpfennig and Thirtle (1999) and Johnson and Evenson (1999) analyse the effects of international public and/or private agricultural R&D on domestic agricultural productivity growth. They find, firstly, the presence of strong international spillovers in the agricultural sector and secondly that, without recognising knowledge spillovers, researchers will end with biased estimates of R&D elasticities.

However, the international transfer of agricultural technology is more difficult than that of industrial technology, Hayami (1997), Hayami and Ruttan (1985) and Sachs (2001). Modern agricultural technology has mainly been improved in developed countries located in temperate zones. Thus, without appropriate adaptive research which helps to assimilate and exploit externally available information, countries located in other ecological zones, for example tropical zones, may not benefit from technological spillovers.

In the next section of the paper, we present a theoretical model, mainly derived from Grossman and Helpman (1991), that links total factor productivity to the cumulative spending on R&D. In the third section we introduce and review the recent results on estimation and inference in panel cointegration. As is well known, cointegrating regression enables us to exploit the relationship among the variables in levels without transforming the data, such as by differencing, to avoid spurious regression problems. In section four we estimate a simple Cobb-Douglas production function for a sample of 47 countries during the period 1970-1992 by using panel cointegration. We also split the sample and estimate two production functions, one for the countries in the sample located in temperate zones, and one for the countries in the sample located in tropical zones. The results indicate that both production functions show constant returns to scale but factor elasticities are quite different. Using these results and following Coe and Helpman's (1995) empirical model, we are able to utilise panel cointegrating regression in order to estimate the relationship between total factor productivity and domestic as well as foreign R&D capital stocks. Using the estimates we calculate the effect of a change in R&D spending in a country on the change of total factor productivity in that country as well as in partner countries. Summarising, we find strong R&D

spillovers between countries located in temperate zones and, inside this group, between EU countries. International spillovers are of less importance when analysing tropical countries. Finally, section five concludes.

2. Theoretical Framework

In the following section, we briefly present a simple innovation-based growth theoretical model that links total factor productivity to the spending on research and development in the sector. The model is mainly derived from Grossman and Helpman's (1991) work.

We assume that agricultural output is produced in a competitive environment and has a Cobb-Douglas production function of the form

$$Y = AK^{\alpha}L^{\beta}\sum_{j=1}^{N}\left(X_j\right)^{1-\alpha-\beta}, \quad \alpha,\beta > 0,\ \alpha+\beta<1, \tag{1}$$

where Y is agricultural output, A is a constant, K is capital and L is the amount of labour used to product the final agricultural output. Output is a function of the X_j non durable intermediate inputs, numbered from 1 to N, used in the production process. From (1) we note first that the production function shows diminishing marginal productivity for each input K, L and X_j and constant returns to scale in all inputs together. Second, the marginal productivity of intermediate input j is independent of the quantity employed of intermediate input j'. Thus the innovation of new types of intermediate inputs do not tend to make any existing types obsolete. In this environment, the technological progress can be seen as improvements in the number N of intermediate inputs and we assume that this advance requires purposive effort in the form of R&D.

Defining the price of intermediate input as p_j and setting output price $p_y = 1$, from profit function maximisation we can derive the demand for input j

$$X_j = \left[\left(1-\alpha-\beta\right)AK^{\alpha}L^{\beta} / p_j\right]^{\frac{1}{\alpha+\beta}} \tag{2}$$

In these models, the inventor of new intermediate goods is usually seen as a monopolist who retains a monopoly right over the production and sale of the goods that uses his/her design. Assuming a marginal unit cost to produce the intermediate goods, a monopolist will set the price maximising the following expression

$$\max_{p_j}\left(p_j - 1\right)X_j \tag{3}$$

Substituting (2) in (3), the solution for monopoly price is

$$p_j = p = \left[1/\left(1-\alpha-\beta\right)\right] > 1 \tag{4}$$

We can now introduce (4) in (2) and utilising the result in (1) we end with the following production function

$$Y = FK^{a}L^{b} \tag{5}$$

where $a = \alpha/\left(\alpha+\beta\right)$, $b = \beta/\left(\alpha+\beta\right)$ and by definition $\left(a+b\right)=1$, i.e. the production function shows constant returns to scale on the two inputs K and L. The variable F , usually defined as total factor productivity, can be written as

$$F = A^{1/(\alpha+\beta)}\left(1-\alpha-\beta\right)^{2(1-\alpha-\beta)/(\alpha+\beta)} N$$

Given α and β as well as A values, it is clear from the above expression that in this model total factor productivity depends on the available assortment of intermediate inputs N: the more intermediates are used in production, the higher is total factor productivity. If the flow of these intermediate goods is proportional to real spending on research and development *RD*, we have that $N(T) = \delta \int_{-\infty}^{T} RD(t)dt$, where δ is a parameter that links, in each period, the growth rate of the number of intermediate inputs to the R&D spending. We therefore have a relationship between current total factor productivity and cumulative R&D investment. This link is central to the innovation based endogenous model and to our empirical specification.

Until now, innovation has been associated with an expansion in the range of intermediate products used in the production process. We can think of this activity as basic innovation that amount to new kinds of goods or method of production. Aghion and Howitt (1992) and Grossman and Helpman (1991, Ch. 4) see innovation as improvements in the quality of intermediate inputs also. Clearly the two models identify different aspects of the innovation process and should be viewed as complements rather than as substitutes.

The production function (1) can now be written as

$$Y = AK^{\alpha}L^{\beta}\sum_{j=1}^{N}\left(\tilde{X}_j\right)^{1-\alpha-\beta}, \quad \alpha,\beta > 0,\ \alpha+\beta<1, \tag{6}$$

where $\tilde{X}_j = \sum_{k=0}^{k_j} \lambda^k X_{jk}$ is the amount of jth intermediate input used in the production process and λ^k is a coefficient, $\lambda > 1$, that adjusts intermediate goods for quality. For $k=0$, $\lambda = 1$, and the subsequent values are at the level $\lambda, \lambda^2, \ldots, \lambda^{k_j}$. The value k_j is the highest quality level in the sector j.[1] In the previous model, technological progress was seen as an improvement in the number *N* of intermediate inputs and we assumed that this advance required R&D investment. In this case, we assume that the value of k_j increases in response to the R&D investment aimed at quality improvement of the intermediate inputs.

The analysis is similar to the previous one. We assume for simplicity that only leading quality is used.[2] As before, in this case the monopoly price for the leading input X_{jk_j} is equal to $p_{jk_j} = 1/\left(1-\alpha-\beta\right)$ and the quantity of intermediate goods produced is given by

$$X_{jk_j} = \left[\left(1-\alpha-\beta\right)^2 AK^{\alpha}L^{\beta}\lambda^{k_j(1-\alpha-\beta)}\right]^{1/(\alpha+\beta)}$$

Substituting this result in the production function (6) we end with the production function (5) where now total factor productivity can be written as

$$F = A^{1/(\alpha+\beta)}\left(1-\alpha-\beta\right)^{2(1-\alpha-\beta)/(\alpha+\beta)} I$$

and $I = \sum_{j=1}^{N} \lambda^{k_j(1-\alpha-\beta)/(\alpha+\beta)}$. Note that increases in the k_j's influence total factor productivity. If we assume that in each period the improvements in the quality of products equals $\dot{I}(t)$ and as before this rise is proportional to real spending in R&D, I(T) will be proportional to cumulative R&D spending in the sector.

As we have seen, theory suggests that productivity depends on the domestic R&D capital stock. Recent developments in the theory of international trade and economic growth have, in addition, identified a number of channels through which a country's external relationships might affect its productivity performance. Grossman and Helpman (1991, Ch. 9) identify four distinct channels. First, international trade opens channels of communication that facilitate the transmission of technical information. This helps the spread of new production methods and the employment of domestic resources more efficiently. Second, trade reduces duplication of research by encouraging producers in each country to pursue new and distinctive ideas and technologies. Third, international trade enlarges the size of the market which influences the incentives to innovate. Finally, when countries' research experience differ, or when the composition of their endowment bundles differ, international trade induces patterns of specialisation that has implications for productivity growth in each of the trading partners. Thus, total factor productivity will be influenced not only by the domestic R&D spending but also by the foreign R&D spending of a country's trade partners.

However, international transfer of agricultural technology is not easy. The sector is strongly constrained by geographical conditions and consequently it is difficult, without adaptive research, to transfer advanced technologies developed in the temperate zones to the tropical zones. This issue is well known in economic literature. Hayami and Ruttan (1985, pg. 255) highlight that "Less developed countries can acquire substantial gains in agricultural productivity by borrowing advanced technology existing in developed countries…. (but) the direct transfer of agricultural technology from other agroclimatic regions have been largely unsuccessful". Recent works by Hayami (1997), Johnson and Evenson (2000), Gutierrez (2000) and Sachs (2001) analyse this point. In the empirical section we will address these issues by using Coe and Helpman's (1995) empirical model for a sample of 47 countries.

3. Panel unit roots and cointegration: theoretical background.

Several studies have examined whether the time series behavior of economic variables is consistent with a unit root (see for a survey Diebold and Nerlove, 1990; Campbell and Perron 1991). In general, the analysis has been carried out by using tests such as the augmented Dickey-Fuller's (ADF) (Dickey and Fuller, 1981) test or semi-parametric tests, as in the case of Phillips-Perron tests (Phillips and Perron, 1988). The main problem here is that, in finite sample, any unit roots process can be approximated by a trend-stationary process. For example, the simple difference stationary process $y_t = f y_{t-1} + e_t$ with $f = 1$ can be arbitrarily well approximated by a stationary process with f less than but close to one. The result is that unit root test statistics have limited power against the alternative. Campbell and Perron (1991) show that when 100 observations are generated by a stationary process but with a root close to unity, then the unit root tests have very little power. They compare the case $f = 1$ with the stationary case $f = 0.98$ and find that the rejection rate is no more that 1% greater for the stationary case than for the unit root case.

Recently, starting from the seminal works of Quah (1990, 1994), Breitung and Meyer (1991) and Levin and Lin (1992, 1993), many tests have been proposed which attempt to introduce unit root tests in panel data. They show that combining the time series information with that from the cross-section, the inference about the existence of unit roots can be made more straightforward and precise, especially when the time series dimension of the data is not very long and similar data may be obtained across a cross-section of units such as countries or industries. A second advantage when using panel unit root tests is that, whereas many of the estimators and statistics for unit root processes in time series are complicated distributions of Wiener processes, the former estimators are normally distributed. This result is still robust when heterogeneity is introduced across the units comprising the panel.

The problem now is that we need new multivariate central limit theorems in order to analyze the asymptotic properties of estimators and tests. Recently, Phillips and Moon (1999a) have presented the formal and general treatment of the asymptotic behaviour of a double indexed integrated process. The limit of the process may depend on which index, N (the units) or T (the time), tend to infinity. We can fix N and allow T to tend to infinity and then pass N to infinity or permit T and N to tend to infinity at a given controlled rate. For example, Levin and Lin (1992, 1993) show that their panel unit root statistics have limiting normal distributions as N and T tend to infinity with $N / T \to 0$ and Im *et al.* (1997) proposes a set of normally distributed test statistics for N and T sufficiently large and $N/T \rightarrow k$, where k is a positive constant. In the following, we shall present a short review of the Levin-Lin tests and their extension by Im, Pesaran and Shin (1997) which have been used in the empirical literature on panel unit root tests and will be proposed in the empirical section.

3.1 The Levin–Lin unit root tests.

Levin and Lin (1993), (LL), consider a sample of N cross-sections observed over T time periods. They suppose that the stochastic process $\{y_{it}\}$ for i=1,…,N and t=1,…,T can be generated by one of the following three models:

model 1 : $\Delta y_{it} = \beta_i y_{it-1} + \varepsilon_{it}$

model 2 : $\Delta y_{it} = \alpha_i + \beta_i y_{it-1} + \varepsilon_{it}$

model 3: $\Delta y_{it} = \alpha_i + \delta t + \beta_i y_{it-1} + \varepsilon_{it}$,

where $\Delta y_{it} \equiv y_{it} - y_{it-1}$ follows a stationary ARMA process for each cross-section unit and ε_{it} are independently and identically distributed both across i and t with finite variance. If we consider model 1, the null hypothesis of unit roots can be expressed as

$$\mathrm{H}_0 : \beta_i = 0 \text{ for all } i, \tag{7.1}$$

against the alternatives,

$$\mathrm{H}_A : \beta_i = \beta < 0 \text{ for all } i. \tag{7.2}$$

From (7.1) and (7.2) emerges one of the major drawbacks of the LL tests. They require β to be homogenous across i. For example, in testing convergence hypothesis in growth models, such as in Gutierrez (1999, 2000) this means that, under the null, none of the countries converge, against the alternative hypothesis that all countries converge and at the same rate.

Auxiliary assumptions under the null are required for the coefficients α_i and α_i, δ_i respectively for the model 2 and 3. In the first case the null is given by $\mathrm{H}_0 : \beta_i = 0$ and $\alpha_i = 0$ for all i, against the alternative $\mathrm{H}_A : \beta_i = \beta < 0$ and $\alpha_i \in \mathbb{R}$ for all i, while in the second case we have a null hypothesis $\mathrm{H}_0 : \beta_i = 0$ and $\delta_i = 0$ for all i against the alternative that $\mathrm{H}_A : \beta_i = \beta < 0$ and $\delta_i \in \mathbb{R}$ for all i. It is useful to underline that here, as for the univariate process, when a deterministic component is present in the observed data but it is not included in the regression procedure, the unit root test will be inconsistent, and when included in the regression analysis but not present in the observed data, the statistical power of the unit root test will be reduced. LL procedure to test panel unit root involves the following steps:

i) Remove cross-section averages from the observed data to eliminate the influence of the aggregate effects.

ii) Instead of applying the augmented Dickey-Fuller (ADF) test to each series i, perform two auxiliary regressions of Δy_{it} and y_{it-1} with respect to the p_i lagged first differences $\Delta y_{it-1}, \ldots, \Delta y_{it-p_i}$ and the appropriate deterministic variables, where the maximum lag p_i is permitted to vary across the units [3], and calculate the residuals, respectively $\hat{e}_{it}$ and $\hat{v}_{it-1}$, from these two auxiliary regressions. Now regress $\hat{e}_{it}$ on $\hat{v}_{it-1}$ to get the OLS values for β_i:

$$\hat{e}_{it} = \beta_i \hat{v}_{it-1} + \varepsilon_{it}.$$

To control for heterogeneity in ε_{it} they suggest the following normalisation

$$\hat{\sigma}_{e_i} = \sum_{t=p_i+2}^{T} \left(\hat{e}_{it} - \hat{\delta}_i \hat{v}_{it-1}\right)^2 / (T - p_i - 1)$$

$$\tilde{e}_{it} = \hat{e}_{it} / \hat{\sigma}_{e_i}$$

$$\tilde{v}_{it-1} = \hat{v}_{it-1} / \hat{\sigma}_{e_i}.$$

Now asymptotically $\tilde{e}_{it}$ are i.i.d for all the units i.

iii) Estimate the ratio of long-run to short-run standard deviation for each series i and calculate the average ratio for all units i as

$$\hat{S}_{NT} = \sum_{i=1}^{N} \left(\frac{\hat{\sigma}_{y_i}}{\hat{\sigma}_{e_i}} \right) \Big/ N$$

where the long-run variance $\hat{\sigma}_{y_i}$ is calculated as

$$\hat{\sigma}^2_{y_i} = \frac{1}{T-1} \sum_{t=2}^{T} \Delta y^2_{it} + 2 \sum_{L=1}^{\bar{K}} w_{\bar{K}L} \left(\frac{1}{T-1} \sum_{t=L+2}^{T} \Delta y_{it} \Delta y_{it-L} \right),$$

and $\bar{K}$ is the lag truncation parameter and $w_{\bar{K}L}$ is a lag window.

iv) Compute the panel test statistic. Under the null hypothesis the normalised residuals $\tilde{e}_{it}$ are independent of the lagged residual $\tilde{v}_{it-1}$ and writing $\tilde{e} = \left(\tilde{e}_{1t}, \tilde{e}_{2t}, \ldots, \tilde{e}_{Nt}\right)'$ and $\tilde{v} = \left(\tilde{v}_{1t-1}, \tilde{v}_{2t-1}, \ldots, \tilde{v}_{Nt-1}\right)'$, this hypothesis can be tested by running the following regression

$$\tilde{e} = \delta \tilde{v} + \varepsilon$$

where now all i and t observations are used. The t-statistics is given by

$$t_{\delta=0} = \frac{\hat{\delta}}{\mathrm{SE}\left(\hat{\delta}\right)} \tag{7.3}$$

where

$$\mathrm{SE}\left(\hat{\delta}\right) = \sigma_{\varepsilon} \left(\tilde{v}'\tilde{v}\right)^{-1/2}$$

$$\sigma^2_{\varepsilon} = \left(\varepsilon'\varepsilon\right) / N\tilde{T}$$

$$\tilde{T} = (T - \bar{p} - 1) \text{ and } \bar{p} = \sum_{i=1}^{N} p_i$$

is the average lag used in the individual ADF regressions.

LL show that the t-statistic (7.3) has a standard normal limiting distribution for model 1, but it is not centered at zero for model 2 and 3. Thus, they propose the following adjusted t-statistic:

$$t^*_{\delta} = \frac{t_{\delta=0} - N\tilde{T}\hat{S}_{NT}\sigma_{\varepsilon}^{-2}\mathrm{SE}\left(\hat{\delta}\right)\mu^*_{\tilde{T}}}{\sigma^*_{\tilde{T}}} \tag{7.4}$$

where all the terms have been previously defined and $\mu^*_{\tilde{T}}$ and $\sigma^*_{\tilde{T}}$ are the mean and standard deviation adjustment obtained from the Monte Carlo simulation and tabulated in their paper.[4]

3.2 The Im-Pesaran-Shin tests.

Im, Pesaran and Shin (1997) (IPS) introduce t-statistics as well as Lagrange Multiplier statistics for unit roots in panel where the alternative hypothesis allows for β_i to differ across groups. Then the hypothesis of unit roots becomes

$$\mathrm{H}_0 : \beta_i = 0 \text{ for all } i, \tag{7.5}$$

against the alternatives,

$$\mathrm{H}_\mathrm{A} : \beta_i < 0 \;\; i = 1,2,\ldots,N_1,\;\; \beta_i = 0,\;\; i = N_1 + 1, N_1 + 2, \ldots, N. \tag{7.6}$$

Note that in this case the IPS testing approach allows, under the alternative hypothesis, for some of the individual series to have unit roots. To test the null hypothesis, IPS use separate unit root test t-statistics, t_{iT}, for each of N cross-section units and define a t-bar statistic as $\bar{t}_{NT} = \sum_{i=1}^{N} t_{iT} / N$. Under the assumption that the second moment of t_{iT} exists for all *i,* they propose the following group-mean t-bar statistic

$$\Gamma_{\bar{t}} = \frac{\sqrt{N}\,\{\bar{t}_{NT} - E(t_T \mid \beta_i = 0)\}}{\sqrt{Var(t_T \mid \beta_i = 0)}} \Rightarrow N(0,1) \tag{7.7}$$

where $E(t_T \mid \beta_i = 0)$ and $Var(t_T \mid \beta_i = 0)$ are the common mean and variance of t_{iT} obtained under $\beta_i = 0$. As previously introduced, consistency for $\Gamma_{\bar{t}}$ is guaranteed when *N* and *T* go to infinity and $N/T \rightarrow k$. In a similar way IPS show that a group-mean LM-bar statistic, under the null hypothesis of unit roots and as *N* and *T* go to infinity, is normally distributed. IPS also report in their paper sample critical values for both statistics. [5]

3.3 Panel cointegration tests.

One difficulty that can arise when regressing two non-stationary series is the problem of *spurious regression*: when using two unrelated integrated series, regressing one on the other tend to produce a not consistent but apparently significant $\hat{\beta}$ parameter, Granger and Newbold (1974).

By contrast with the pure time series spurious regression, in the case of non-stationary panel data, Phillips and Moon (1999a) show that for the spurious panel regression, and under quite weak regularity conditions, the pooled least squares estimator of the coefficient β is consistent and has a limiting normal distribution. The reason is that independent cross-section data in the panels introduce information and this leads to a stronger signal than the pure time series case. The problem here is that while the coefficient β converges to its true value, t-statistic diverges so that inferences about $\hat{\beta}$ are wrong with probability that goes to one asymptotically, Kao et al. (1999).

In the empirical analysis we will use two sets of cointegration tests. The first set of tests has been proposed by Kao et al. (1999), and can be seen as a generalisation of the Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) tests in the context of panel data. The tests consist of taking as null the hypothesis of no cointegration and using the residuals derived from a panel static regression to construct the test statistics and tabulate the distributions. Defining $\hat{e}_{it}$ as the estimated residuals from the static regression, the DF tests can be derived from the following regression

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + \nu_{it}, \tag{8.1}$$

The null of no cointegration can be written as $H_0 : \rho = 1$.

Kao et al. (1999) propose four DF type tests

1. $\mathrm{DF}_{\gamma} = \dfrac{\sqrt{N}T(\hat{\gamma}-1)+3\sqrt{N}}{\sqrt{10.2}}$

2. $\mathrm{DF}_{t} = \sqrt{1.25}t_{\gamma} + \sqrt{1.875}N$

3. $\mathrm{DF}^{*}_{\gamma} = \dfrac{\sqrt{N}T(\hat{\gamma}-1)+\left(3\sqrt{N}\hat{\sigma}_{v}^{2}/\hat{\sigma}_{0v}^{2}\right)}{\sqrt{3+\left(7.2\hat{\sigma}_{v}^{4}/\hat{\sigma}_{0v}^{4}\right)}}$

4. $\mathrm{DF}^{*}_{t} = \dfrac{t_{\gamma}+\left(\sqrt{6N}\hat{\sigma}_{v}/\hat{\sigma}_{0v}\right)}{\sqrt{\left(\hat{\sigma}_{0v}^{2}/2\hat{\sigma}_{v}^{2}\right)+\left(3\hat{\sigma}_{0v}^{2}/10\hat{\sigma}_{v}^{2}\right)}}$

where stars indicate tests for cointegrating regression with endogenous x_{it} regressors, t_{γ} is the t-statistic for γ and finally $\hat{\sigma}_{v}^{2} = \Sigma_{u} - \Sigma_{u\varepsilon}\Sigma_{\varepsilon}^{-1}$ and $\hat{\sigma}_{0v}^{2} = \Omega_{u} - \Omega_{u\varepsilon}\Omega_{\varepsilon}^{-1}$ where Σ and Ω are respectively the covariance and long-run covariance matrices of errors in the cointegrating static regression. Kao et al. (1999) proposes an ADF type test. In this case regression (8.1) is augmented in order to include p differenced lags of the static cointegrating regression error terms $\hat{e}_{it}$

$$\hat{e}_{it} = g\hat{e}_{it-1} + \sum_{j=1}^{p} d_{j}\Delta\hat{e}_{it-j} + n_{it}, \tag{8.2}$$

and in this case the ADF test, with the null hypothesis of no cointegration, is given by

5. $\mathrm{ADF} = \dfrac{t_{\gamma}+\left(\sqrt{6N}\hat{\sigma}_{v}/\hat{\sigma}_{0v}\right)}{\sqrt{\left(\hat{\sigma}_{0v}^{2}/2\hat{\sigma}_{v}^{2}\right)+\left(3\hat{\sigma}_{0v}^{2}/10\hat{\sigma}_{v}^{2}\right)}}$

Note that the ADF test is equal to the DF test (4.) except that in this case t_{γ} is the t-statistic for γ in the regression (8.2). All the tests have asymptotic distributions that converge to a standard normal distribution N(0,1).

Pedroni (1999), enlarging on the results in Pedroni (1995), proposes seven panel cointegration statistics for the null of no cointegration in dynamic panels with multiple regressors.[6] The tests allow for heterogeneity among individual units of the panel and, by contrast with Kao's (1999) DF tests 1. and 2., no exogeneity requirements are imposed on the regressors x_{it} in the cointegrating regression. Four of Pedroni's (1999) seven tests are defined as pooling along the within-dimension and three are based along the between-dimension. Within the first set of tests, three require the use of non-parametric corrections as in Phillips and Perron (1988), and the fourth is a parametric ADF test. In the second set of tests, two use non-parametric corrections while the third is an ADF test.

Denoting g_{i} as the autoregressive coefficient of the residuals in the ith cross-section, the first category of tests requires the following specification of null and alternative hypotheses

$$H_{0}: g_{i} = 1, \text{ for all } i,\ H_{A}: g_{i} = g < 1 \text{ for all } i,$$

and the second set of tests uses

$$H_{0}: g_{i} = 1, \text{ for all } i,\ H_{A}: g_{i} < 1 \text{ for all } i.$$

So all Pedroni's tests can be constructed using the residuals of the cointegrating regression and various nuisance parameter estimators which can be obtained from these and, in most cases, by the specific long run conditional variance for the residuals. Finally Pedroni (1999) shows that, after appropriate standardisation, all tests have asymptotic distributions that converge to a standard normal distribution N(0,1).

4. Empirical Results.

4.1 Total factor productivity estimation.

The purpose now is to provide some basic estimate of total factor productivity in agriculture in a relatively wide range of countries using the methodology presented in the previous section. We follow a simple and widely used approach where total factor productivity is computed as

$$TFP_{i,t} = Y_{i,t} / \left(K_{i,t}^{\alpha} L_{i,t}^{\beta} T_{i,t}^{\delta} \right), \tag{9}$$

where $Y_{i,t}$ is the value added in the agricultural sector for country i in time period t, $K_{i,t}$ is the capital stock, $L_{i,t}$ is the quantity of labour, $T_{i,t}$ is the quantity of land, α, β and δ are respectively the elasticity of capital, labour and land with respect to value added and $TFP_{i,t}$ is the total factor productivity variable. Naturally we have that when $\alpha + \beta + \delta = 1$ the production function shows constant returns to scale (which later will be tested) and constancy of factor elasticities across countries and over time. The assumption of constant returns has recently received empirical support from Mundlak *et al.* (1997) and has extensively been used by Bernard and Jones (1996) when testing for productivity convergence between countries and sectors and Gutierrez (2000) for EU and US agricultural sectors. The assumption of constancy of factor elasticities may be too restrictive. We try to correct for this problem by estimating α, β and δ for a different set of countries, in our case temperate and tropical countries.

The data for output comes from the World Bank and are given by the gross value added in the agricultural sector in constant 1990 US dollars. Fixed capital stock measures, in constant 1990 US dollars, were kindly supplied by Donald Larson and referenced in Crego *et al.* (1997). Hectares of arable and permanent cropland are used for land input and labour is given by the economically active population in agriculture. Both variables come from FAO data set.

We start the analysis determining whether the variables included in (9) are stationary or non-stationary, i.e. whether the series contain unit roots. We use the Levin and Lin (LL) tests and the Im, Pesaran and Shin (IPS) tests presented in the previous section. The results are reported in Table 1.

Table 1 about here

Both IPS and LL tests are one-sided tests from a N(0,1) distribution, thus a statistic less than -1.65 or -2.33 would case rejection respectively at 5 percent and 1 percent of the null hypothesis of non-stationarity. Looking at Table 1 we note that all the variables, with one exception, fail to reject the null of non-stationarity. The exception is the labour variable when using the LL test without a time trend but including individual specific effects. In any case LL test fails to reject the null when including a time trend in the model and when using IPS tests. Thus, given the presence of non-stationary variables, we proceed by estimating the production function and testing for cointegration.

The factor elasticity estimates for the full sample of countries and for a subset of twenty-five countries located in temperate area and twenty-two countries located in tropical area are given in Table 2. We decide whether to include a country in the tropical or temperate subset depending on whether more than 50% of land area is located inside or outside the tropics.

The asymptotic properties of the estimators and associated statistical tests in cointegrated panel models are quite different from those of the time series regression models. Kao and Chen (1995) and Chen, McCoskey and Kao (1999) show that the OLS estimator is asymptotically normal but asymptotically biased and propose a method to correct the estimates. Secondly, they found that different estimators based on fully modified (FM) estimator or dynamic OLS (DOLS) estimator can be more promising in cointegrated panel regressions. The first estimator is a panel generalisation of the Phillips and Hansen (1990) time series estimator and has been proposed for the first time by

Pedroni (1996). The second one has been used in Kao, Chiang and Chen (1999) and was built as panel generalisation of the Stock and Watson (1993) time series estimator. In Table 2 we present biased-corrected OLS, FM and DOLS estimates. Finally, Phillips and Moon (1999b, pg.12) suggest detrending variables in order to obtain consistent estimation of long-run average estimates, so that all our variables have been previously detrended using OLS regression. [7]

Following these results, in Table 2 each column reports the factor price elasticities obtained from the bias-corrected OLS estimator, FM estimator and finally DOLS estimator. All the estimates have been carried out under the assumption of homogeneous long-run covariance across cross-sectional units.[8]

Table 2 about here

Many interesting results emerge from Table 2. The production elasticities and their levels of statistical significance are satisfactory and the three methods provide quite similar results. Capital elasticities are generally higher than labour and land estimates both for the total sample of countries as well as for the sample of temperate and tropical countries. The two subsets of countries seem to have the same values for capital elasticities but show differences when comparing labour and land elasticities. Table 2 reports an higher value for labour elasticity and a lower value for land elasticity in temperate countries. Looking at the elasticities for the total sample of countries it is interesting to note that their sum is near one, revealing the possible presence of constant returns to scale. We tested the null hypothesis of constant returns to scale by using the Wald test proposed in Kao and Chiang (1995).[9] They show that in cointegrated panel regressions the test converges in distribution to a chi-squared random variable with *m* degrees of freedom, where *m* is the total number of restrictions, in our case one. All the test statistics do not reject the null hypothesis of constant returns to scale for the total sample of countries and for the subsets of tropical countries. Some evidence of increasing return to scale is testified to by the Wald statistics for the sample of temperate countries when using FM and DOLS method. Finding increasing return of scale in the agricultural sector is not new. Griliches (1963) reports increasing return in cross-regions analysis for the United States. Hayami and Ruttan (1985) provides evidence of increasing returns for a sample of developed countries and they find that for a sample of less developed countries the sum of conventional input coefficients is not significantly different from one. This finding may support our results. The sample of countries located in temperate regions is mainly constituted by developed countries whereas many countries which we label tropical countries are defined by the World Bank as less developed countries.

In order to compare our results with some previous ones, in Table 3 we collect factor input elasticities obtained from DOLS estimation and four previous attempts to estimate intercountry production functions. All the estimates have been scaled by their sums in order to obtain comparable values.

Table 3 about here

Looking at the table, it emerges that capital elasticity estimates are usually higher than the labour or land elasticities . Only Mundlak *et al.* (1997) propose a land elasticity estimate higher than labour elasticity. Finally, our labour estimates is halfway between the highest value of 0.45 proposed by Hayami and Ruttan (1985) and the lowest value of 0.09 proposed by Mundlak *et al.* (1987).

We are now ready to test whether estimated equations are actually cointegrated. Table 4 reports cointegration test results using the previously reported panel cointegration tests.

Table 4 about here

All the test statistics are in general significant so that the null hypothesis of no cointegration is rejected. Some doubts arise when analysing test statistics for the sample of tropical countries. In this case three of Kao's (1999) statistics do not reject the null hypothesis of no cointegration.

Using the estimated elasticities in Table 5 we are now able to highlight, for the period 1970-1992, the average annual growth rate of the total factor productivity, labour productivity, capital and land intensities for the full sample of countries as well as for the two subsets of temperate and tropical countries.

Table 5 about here

During the period of analysis, labour productivity in countries located in temperate zones was mainly influenced by the strong increase in the capital-labour ratio whereas total productivity growth was quite low (0.33 per cent). At the same time, tropical countries saw labour productivity growth mainly influenced by total factor productivity growth. This means that whereas labour productivity growth in temperate countries is explained by the substitution of other factors for labour, in tropical countries, where the use of labour increased during the period at an annual average rate of 0.8 per cent, labour productivity growth was mainly affected by change in total factor productivity.

4.2 International R&D spillovers and productivity growth in agricultural sector.

The panel cointegration approach can be usefully used in estimating the long-run relationship between total factor productivity and the domestic and foreign R&D capital stocks. The aim of the section is twofold. First, we estimate the effects of a rise in a country's R&D capital stock on the country's total factor productivity. As seen in the introductory section, this issue has been long debated in the literature as testified by the large number of works published on this theme, which are however mainly devoted to calculating the rates of return of agricultural R&D. Second, we are interested in analysing the effect of foreign R&D capital stock on total factor productivity in order to introduce new evidence on the effects of new technology from one country on its trade partners and between different climate zones. This point has been debated at length in literature. For example, Thirtle and Bottomley (1989, pg. 1082), studying the effect of public UK R&D on total factor productivity, recognised that "spillover of new technology from one jurisdiction to others is an even more insoluble problem".

In this section, using Coe and Helpman's (1995) empirical model we attempt to provide evidence on this issue. We find that foreign agricultural R&D capital stock has a strong effect on a country's total factor productivity. This effect is stronger for countries located in homogeneous climate zones. For example an increase of US agricultural R&D has a larger effect on countries located in temperate zones and less on tropical countries. Once more these results paint the agriculture sector as strongly constrained by environmental conditions where, by contrast industrial sector, transferring technologies developed in the temperate zones to tropical zones is difficult.

As previously underlined, Coe and Helpman's (1995) empirical model provides a source for analysing the relationship between a country's own R&D as well as the R&D efforts of its trade partners and productivity growth. They estimate the following log linear equation:

$$\log\ TFP_i = \alpha_{0i} + \alpha_d SRD_{di} + \alpha_f \left(m_i \log\ SRD_{fi} \right), \tag{10}$$

where α_{0i} are country-specific constants that can differ across countries, SRD_{di} represents the domestic R&D capital stock of country *i*.

Domestic capital stock is built following a perpetual inventory model,

$$SRD_{dt} = (1-\delta) SRD_{dt-1} + RD_{t-1}$$

where RD_t are the agricultural R&D expenditures at the time t and δ is the depreciation rate. The starting value for SRD_{di} was calculated following Griliches (1980) as

$$SRD_{d0} = RD_0 / \left(\delta + g \right)$$

where g is the average annual logarithmic growth of R&D expenditures over the period of analysis.[10]

The variable SRD_{fi} in (10) represents the foreign R&D capital stock, defined as a weighted average of the domestic R&D capital stock of trade partners. Coe and Helpman (1995) use as weights the bilateral total import share provided by the IMF's *Direction of Trade*. Agricultural bilateral imports for our full sample of countries and period are unavailable, so we construct two measures of the foreign R&D capital stock. The first is obtained for each period and country as a simple sum of domestic R&D capital stock in the other countries and the second measure defines foreign R&D capital stock in country *i* as the bilateral total import-share-weighted average of agricultural domestic R&D capital stock of the remaining $i-1$ countries.

In equation (10) the log of foreign R&D is multiplied by the variable m_i, which in this case stands for the fraction of agricultural imports relative to agricultural GDP for country *i*. The hypothesis is that the country where agricultural imports are higher relative to its GDP may benefit more from foreign R&D. Therefore, the composite variable $m_i \log SRD_{fi}$ can account for the interaction between foreign agricultural R&D capital stocks and the international level of agricultural trade.

Public agricultural R&D expenditures were collected from various sources. The main ones are Alston *et al.*(1999) for OECD countries, Cremers and Roseboom (1997) for Latin America countries and Tabor *et al.* (1998) for other African and Asian countries.

Table 6 reports the bias-corrected OLS, FM, DOLS pooled cointegrating regressions based on equation (10), all of which include unreported country-specific constants.[11] This time, given the strong significance, DOLS estimates for the heterogeneous case are reported. In order to reduce the endogeneity problem, which may be relevant when estimating equation (10), we introduce both R&D variables with a lag.

Table 6 about here

All the estimates have the expected sign and are significant. The long-run estimated elasticities of TFP with respect to domestic R&D capital stock varies within a range of 0.236 for FM coefficient and 0.391 for DOLS (heterogeneous) estimate. These values are similar to those found in single-country studies. For example, Thirtle and Bottomley (1989) propose long-run estimates ranging from 0.293 and 0.441.[12] Foreign R&D capital has a strong effect. We report only the estimated elasticities when SRD_{fi} is defined using a bilateral-import-scheme because they are better in terms both of adaptation and significance. The elasticity values range from 0.415 and 0.521. The adjusted R^2 coefficients highlight that a substantial reduction of unexplained variance is obtained for the DOLS estimation when error variances are permitted to vary across cross-section units. In this case the model explains more than 40 per cent of the variance of the 1034 observations. Finally, all cointegration test statistics, not reported here for brevity, are significant so the null of no cointegration is strongly rejected.

Table 7 reports the estimated elasticities of total factor productivity with respect to the foreign R&D capital stock[13] for 1970, 1980 and 1990 and a sample of developed and developing countries. Two facts seem to emerge. The first is that the estimated impact of foreign R&D rose during the period, with the exception of India and Zimbabwe. Second, countries located in temperate zones benefit more than tropical countries from foreign R&D. The unweighted elasticity averages for the two sample of countries show a stronger impact of foreign R&D on temperate countries relative to tropical countries but for both, and still on average, the domestic R&D has a larger impact on total factor productivity than does foreign R&D. Note that Table 7 shows the notable exception of the Netherlands and UK to the last assertion. In these countries a rise of domestic R&D in trade partners strongly influences total factor productivity and thus agricultural output.

Table 7 about here

We can make further progress by analysing the international spillovers in the agricultural sector. Each entry in Table 8 presents the estimated elasticity of total factor productivity in the countries indicated in the row with respect to the R&D capital stock in the country indicated in the column.[14]

Table 8 about here

The United States R&D capital stock has the strongest effect on total factor productivity of its trade partners. A 1 per cent increase in the R&D capital stock in this country increases total factor productivity by an average of 0.058 per cent for the full sample of 47 countries. The effect is stronger for the subset of countries located in temperate zones, where the elasticity rises to 0.081, whereas tropical countries are less influenced by R&D in the United States. Looking at the values for single countries, the United States has the strongest effect on the Netherlands and UK (the elasticity are 0.47 and 0.34, respectively). European countries are well integrated. A 1 per cent increase in the R&D capital stock in France increases total factor productivity in Italy by 0.06 per cent, in the Netherlands by 0.09 per cent, in UK by 0.05 per cent. Japan and the USA are less influenced, with elasticities respectively of 0.002 and 0.003 per cent. Similar effects are easily verifiable for an increase in R&D capital stock in Italy, in the Netherlands and in UK. As in the Table 7, we compute elasticities also for a set of developing countries. Note how a rise in the R&D capital stock in the column countries has a lower effect on total factor productivity in India, Pakistan, Philippines, Kenya and Zimbabwe. Similar values can be reported for the other countries located in tropical zones.

Finally, we can estimate the rates of return on investment in R&D. Instead of calculating the rate of return for the full set of countries we concentrate attention on the average rates of return for the two groups of countries: countries located in temperate zones and countries located in tropical zones. The average rate of return for the first set of countries in 1990 was 114 per cent in the temperate countries and 217 per cent in the tropical countries.[15] Values above 100 percent for the rate of return of agricultural R&D are not new (as summarised in Alston *et al*. (2000)) but our results must be treated with care. They are sensitive to the level of R&D capital stock which is influenced, for example, by the depreciation rate used to compute the initial value for R&D capital stock. A lower value of the depreciation rate increases the R&D capital stock and thus reduces the rate of return. Elasticities are less influenced by this problem due to the presence of country dummies in the regressions.

5. Conclusions

This paper investigates the question of how R&D spending and trade affects total factor productivity in the agricultural sector. Although this is not a new question, only recently has the new economic growth literature provided theoretical as well as empirical models to analyse this field of research.

This paper addresses this problem computing total factor productivity in the agricultural sector for a sample of 47 countries during the period 1970-1992 and uses this variable to analyse the relationship with domestic and foreign public R&D spending in agriculture. New panel cointegration econometrics has been adopted to compute sound long-run estimates.

Many interesting results emerge from the analysis. First a country's total factor productivity is positively and significantly influenced not only by its domestic R&D capital stock but also by the foreign R&D capital stock of its trade partners. Second, geographic factors influence international spillovers in the agricultural sector. Countries located in temperate zones benefit more than countries located in tropical zones from technological spillovers. Thus, temperate countries need a lower effort in technological capability, i.e. less investments in adaptive research are needed to make effective use of technological knowledge and generate sizeable spillover benefits. Third, the USA is the country that exerts the major impact in transferring agricultural R&D world-wide. A 1 per cent increase in the R&D capital stock in this country increases total factor productivity by an average of 0.058 per cent for the full sample of 47 countries. The effect is stronger for the countries located in temperate zones, 0.081 per cent, than for the countries located in tropical zones, 0.017 per cent. The Netherlands and UK are the two countries in Europe that benefit most from agricultural R&D spending in the US. R&D investment in EU countries mainly influences agricultural productivity and output in these countries and a lesser impact is shown on US or Japanese total factor productivity. Finally, the average rate of returns for agricultural R&D spending is higher in tropical countries than temperate countries. This finding could provide evidence to justify new support and even a greater investment of funds for agricultural R&D in tropical zones.

Note

[1] Note that in the previous model, where quality improvements were not considered, $k_j = 0$ was applied for each jth intermediate inputs.

[2] A more general specification, where all inputs of different qualities are used, gives substantially the same results. See Grossman and Helpman (1991, Ch.4).

[3] Levin and Lin (1993) recommend using the method proposed by Hall (1994) for selecting the appropriate lag order.

[4]The procedure has been implemented in GAUSS 3.2 by Chiang and Kao (2000).

[5] See note 4.

[6] We do not discuss the seven tests here. Pedroni collects the tests in Table 1. (pg. 660) of Pedroni's (1999) paper and further discussion can be found in Pedroni (1997).

[7] Phillips and Moon (1999) provide evidence that the OLS method gives better results than the GLS methods.

[8] We also ran panel regression with heterogeneous cross-section variance but we obtained worse results.

[9] The Wald test is computed under the hypothesis of homogeneous long-run covariance structure. See Kao and Chiang (1995) remark 9 pg.13.

[10] In order to compare our results with Coe and Helpman's (1995) finding, we assume a value for $\delta = 0.05$. In any case different values as $\delta = 0.01$ or $\delta = 0.10$ do not strongly alter the regression results. Possible pitfalls using the perpetual inventory model when estimating R&D capital stock are discussed in Esposti and Pierani (2000).

[11] For reason of brevity, we do not report unit root tests. The unit root tests on the panel data confirm that all the variables are non-stationary with unit roots.

[12] They use a second degree Almon distribution to model the shape of elasticity. The values that we report are the sum of lag coefficients.

[13] The elasticity is given by the estimated coefficient obtained from DOLS (heterogeneous) method multiplied by agricultural import share.

[14] We use the same formula as in Coe and Helpman (1995). When the R&D capital stock of country *i*, SRD_{dj}, increase by 1%, the foreign R&D capital stock for country *j*, SRD_{fj}, rises by $m_i^j SRD_{di} / \sum_{k \neq j} m_k^j SRD_{dk}$ per cent and county *j*'s TFP rises by $m^j \alpha_f m_i^j SRD_{di} / \sum_{k \neq j} m_k^j SRD_{dk}$ per cent, where m^j is country *j*'s import share and m_i^j is the fraction of *j*'s imports coming from country *i*.

[15] The average rate of return for a set C of homogeneous country equals $\rho_C = \alpha_{dC}\left(\sum_{j\in C} Y_j / \sum_{j\in C} SRD_{dj}\right)$,

where α_{dC} is the domestic elasticity and Y_j is the output (value added in our case) in country *j*.

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Table 1. Panel Data Unit Roots Tests Results

Tests	Variables	Value	Results
Im et al. (1997) tests (a)			
Without a time trend	Y	10.85	Fail to reject
	K	10.72	Fail to reject
	L	10.73	Fail to reject
	T	10.51	Fail to reject
With a time trend	Y	16.12	Fail to reject
	K	16.09	Fail to reject
	L	16.16	Fail to reject
	T	16.16	Fail to reject
Levin and Lin (1993) tests (b)			
Without a time trend	Y	12.08	Fail to reject
	K	9.75	Fail to reject
	L	-5.54	Reject at 0.01
	T	8.67	Fail to reject
With a time trend	Y	41.64	Fail to reject
	K	26.92	Fail to reject
	L	57.71	Fail to reject
	T	23.70	Fail to reject

Notes : (a) One lag included in the ADF process
(b) Include individual specific effect
Source: Author's calculation using Chiang and Kao's (2000) NPT 1.1 package

Table 2. Production Function Estimation Results : pooled sample 1970-1992 (a)(b)

Variables	OLS (c)	FM	DOLS
Total sample (47 countries)			
Fixed Capital	**0.6010**	**0.5918**	**0.5818**
	(27.38)	(25.79)	(21.90)
Labour	**0.2683**	**0.2792**	**0.2902**
	(6.71)	(6.67)	(5.99)
Land	**0.1276**	**0.1334**	**0.1366**
	(2.96)	(2.96)	(2.62)
Sum elasticities	**0.9969**	**1.0034**	**1.0086**
Wald $\chi^2(1)$ test: $(\alpha+\beta+\gamma)=1$	0.026	0.056	0.220
p-value	0.87	0.81	0.64
N	1081	1081	1081
R^2 adjusted	0.99	0.99	0.99
Temperated countries (25 countries)			
Fixed Capital	**0.5629**	**0.5496**	**0.5212**
	(18.59)	(17.36)	(15.71)
Labour	**0.3922**	**0.4121**	**0.4495**
	(7.08)	(7.12)	(7.41)
Land	**0.0930**	**0.0982**	**0.1145**
	(2.13)	(2.15)	(2.39)
Sum elasticities	**1.0481**	**1.0599**	**1.0852**
Wald $\chi^2(1)$ test: $(\alpha+\beta+\gamma)=1$	2.847	4.403	8.917
p-value	0.09	0.036	0.002
N	575	575	575
R^2 adjusted	0.99	0.99	0.99
Tropical countries (22 countries)			
Fixed Capital	**0.5556**	**0.5726**	**0.5508**
	(43.29)	(16.02)	(14.71)
Labour	**0.2337**	**0.2309**	**0.2341**
	(7.38)	(2.50)	(2.42)
Land	**0.2328**	**0.2057**	**0.2390**
	(6.43)	(1.97)	(2.19)
Sum elasticities	**1.0221**	**1.0092**	**1.0239**
Wald $\chi^2(1)$ test: $(\alpha+\beta+\gamma)=1$	0.046	0.120	0.822
p-value	0.83	0.729	0.365
N	506	506	506
R^2 adjusted	0.99	0.99	0.99

Notes: (a) conventional *t*-statistics are reported in parentheses.
(b) all equations include unreported, country-specific constants.
(c) bias-corrected OLS estimates.

Source: Author's calculation using Chiang and Kao's (2000) NPT 1.1 package.

Table 3. Comparison of the estimates of the intercountry agricultural production function

Source	Capital	Labour	Land
Hayami and Ruttan (1985)	0.46	0.45	0.09
Mundlak *et al.* (1997)	0.47	0.09	0.45
Mundlak and Hellinghausen (1982)	0.40	0.40	0.20
Evenson and Kislev (1975)	0.65	0.20	0.10
DOLS estimates	0.58	0.29	0.13

Sources : Hayami and Ruttan (1985) Table 6-4; Mundlak *et al* . (1997) Table 4.

Table 4. Panel Cointegration Tests

Tests	Total sample	Temperated countries	Tropical countries
Pedroni (1996) tests :			
Panel ν-statistics	-11.70(0.00)	-8.35(0.00)	-8.05(0.00)
Panel ρ-statistics	11.09(0.00)	7.50(0.00)	7.83(0.00)
Panel *t*- statistics (c)	14.77(0.00)	8.93(0.00)	10.83(0.00)
Panel *t*- statistics	-155.94(0.00)	-155.36(0.00)	-90.52(0.00)
Group ρ-statistics	12.08(0.00)	8.87(0.00)	8.44(0.00)
Group *t*- statistics (c)	-11.20(0.00)	-8.47(0.00)	-7.29(0.00)
Group *t*- statistics	14.82(0.00)	9.87(0.00)	-10.75(0.00)
Kao (1999) tests :			
DF_ρ	-3.83(0.00)	-3.27(0.00)	-1.48(0.07)
DF_t	55.59(0.00)	26.99(0.00)	24.85(0.00)
DF^*_ρ	-6.43(0.00)	-6.14(0.00)	-3.43(0.00)
DF^*_t	-2.51(0.01)	-2.36(0.01)	-1.28(0.10)
ADF	-2.41(0.01)	-3.50(0.00)	-0.44(0.32)

Notes : (a) The critical probabilities are reported in parentheses

(b) Cointegration test statistics are calculated through the residuals from the DOLS estimation

(c) Nonparametric test

Source: Author's calculation using Chiang and Kao's (2000) NPT 1.1 package

Table 5. Unweighted average annual growth rates labour productivity, capital intensity, land intensity and multifactor productivity 1970-1992.

	Labour productivity	Capital Intensity	Land intensity	Total factor productivity
Total sample	2.13	2.83	0.14	0.71
Temperate countries	2.63	4.50	1.34	0.33
Tropical countries	1.56	0.93	-0.36	1.14

Note : Total factor productivity is equal to a weighted average of the growth in labour, capital and land productivity. The DOLS estimates are used as weights.

Source: Author's calculation based on World Bank, Mundlak *et al.* (1997) and FAO database.

Table 6. Total Factor Productivity Estimation Results pooled data 1970-1992 for 47 countries (a)(b)

Variables	**OLS (c)**	**FM**	**DOLS (homogeneous)**	**DOLS (heterogeneous)**
Total sample (47 countries)				
$\log SRD_{d(t-1)}$	**0.2392**	**0.2356**	**0.3552**	**0.3911**
	(2.719)	(2.557)	(3.304)	(107.617)
$m_{i(t-1)} \log SRD_{f(t-1)}$	**0.5210**	**0.4800**	**0.5612**	**0.4150**
	(2.605)	(2.292)	(2.296)	(45.58)
N	1034	1034	1034	1034
R^2 adjusted	0.140	0.140	0.133	0.43

Notes: (a) conventional *t*-statistics are reported in parentheses.
(b) all equations include unreported, country-specific constants.
(c) bias-corrected OLS estimates.

Source: Author's calculation using Chiang and Kao's (2000) NPT 1.1 package.

Table 7 Elasticity of total factor productivity with respect to foreign R&D capital stock

Countries	1970	1980	1990
France	0.275	0.333	0.331
Italy	0.169	0.212	0.262
Japan	0.098	0.145	0.164
Netherlands	0.572	0.794	0.858
U.K.	0.539	0.681	0.598
U.S.A.	0.094	0.086	0.103
India	0.019	0.008	0.006
Pakistan	0.019	0.036	0.064
Philippines	0.018	0.027	0.058
Kenya	0.035	0.048	0.041
Zimbabwe	0.030	0.055	0.039
Unweighted Average:			
Temperated countries	0.168	0.243	0.247
Tropical countries	0.083	0.109	0.126

Notes: Author's calculation based on DOLS (heterogeneous) estimates.

Table 8 **Cross-Countries estimated elasticity of Total Factor Productivity with respect to R&D capital stock – 1990.**

	France	Italy	Japan	Netherlands	U.K.	U.S.A.
France	-	0.0258	0.0474	0.0171	0.0398	0.1757
Italy	0.0611	-	0.0245	0.0127	0.0213	0.1218
Japan	0.0017	0.0006	-	0.0003	0.0020	0.1358
Netherlands	0.0912	0.0192	0.1000	-	0.1192	0.4703
U.K.	0.0548	0.0165	0.1039	0.0276	-	0.3426
U.S.A.	0.0030	0.0014	0.0662	0.0007	0.0041	-
India	0.0002	0.0001	0.0020	0.0001	0.0004	0.0023
Pakistan	0.0012	0.0009	0.0320	0.0005	0.0005	0.0238
Philippines	0.0005	0.0001	0.0156	0.0002	0.0006	0.0387
Kenya	0.0027	0.0007	0.0161	0.0007	0.0064	0.0097
Zimbabwe	0.0016	0.0007	0.0059	0.0006	0.0045	0.0209
Avg. Temperated Countries	0.0124	0.0039	0.0413	0.0049	0.0108	0.0811
Avg. Tropical Countries	0.0013	0.0005	0.0085	0.0003	0.0011	0.0168
Avg. 47 Countries	0.0083	0.0027	0.0292	0.0032	0.0072	0.0575

Notes : Estimated elasticity of total factor productivity in the row countries with respect to the R&D capital stock in the column country, based on DOLS (heterogeneous) estimates. Averages are calculated using agricultural GDP weights.