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WORKING PAPERS

2006/06

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Total Factor Productivity Estimates: Some Evidence from European Regions

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Abstract

This paper provides total factor productivity estimates for a sample of 115 European Regions over the period 1976-2000. In particular, a set of Cobb-Douglas production functions is estimated using panel techniques and allowing for heterogeneity across regions. Moreover, on the basis of specific panel tests, the paper shows that there is empirical evidence which suggests the presence of unit roots in the series and panel cointegration tests are applied to guard against spurious regression.

Keywords: Total Factor Productivity, Panel Unit Root Test, Panel Cointegration.

JEL Classification: C23, D24, O47, O52

1 Introduction

A common feature of many empirical studies on international comparison of Total Factor Productivity (TFP) has been the assumption of identical aggregate production function for all countries. However, the empirical evidence

suggests that the production function may actually differ across countries but attempts at allowing for such differences have been limited by the fact that most of these studies have been conducted in the framework of single cross-country regressions. In this framework it is econometrically difficult to allow for differences in the production function as are not easily measurable.

Solow (1956) develops a production function in which output growth is a function of capital, labour, and knowledge or technology. Technology is Harrod neutral and it is assumed to be exogenous and homogenous across countries. Economists use growth accounting approach to test the neoclassical growth model, and to evaluate the effect of physical capital accumulation on output growth.

The growth accounting approach provides a breakdown of observed economic growth into components associated with changes in factor inputs and a residual that reflects technological progress and other elements. The basic of growth accounting were presented in Solow (1957).[1]

The results of the early growth accounting exercises raise questions about the large unexplained residual in Solow-model calculations. The neoclassical model emphasizes the role of factor accumulation, neglecting differences in productivity growth and technological change captured by the residual. By defining capital to include physical and human capital, Mankiw (1995) finds that the results more closely resemble the theoretical prediction of the

[1] Differentiation of the neoclassical production function $Y = F(A, K, L)$ with respect to time yields:

$$\frac{\dot{Y}}{Y} = g + (\frac{F_{K} K}{Y}) \cdot (\dot{K}/K) + (\frac{F_{L} L}{Y}) \cdot (\dot{L}/L) \tag{1}$$

where F_K and F_L are the factor marginal products and g is the technological progress, given by:

$$g \equiv (\frac{F_{A} A}{Y}) \cdot (\frac{\dot{A}}{A}) \tag{2}$$

$$g = \frac{\dot{Y}}{Y} - \frac{F_{K} K}{Y}) \cdot (\dot{K}/K) - (\frac{F_{L} L}{Y}) \cdot (\dot{L}/L) \tag{3}$$

If the technological progress is Hicks neutral then $F(A, K, L) = A \cdot \tilde{F}(K, L)$ and $g = \frac{\dot{A}}{A}$. The technological change can be calculated as a residual (*Solow residual*) from (1).

neoclassical model. The works of Barro and Sala-i-Martin (1995), Mankiw, Romer and Weil (1992) follow a similar perspective.

Easterly and Levine (2001) suggest that growth economists should focus on TFP and its determinants rather than factor accumulation. They point out that much of the empirical evidence accumulated to date indicates that factor accumulation explains only a portion of the observed cross-country growth. Solow (1956) himself finds that income growth is explained only in little part by capital accumulation while the rest is explained by productivity growth. Easterly and Levine (2001) also observe that exists a tendency of production factors to move to the same places, causing a concentration of economic activity. In such circumstances, to apply the neoclassical model with homogenous technology is not appropriate.

Endogenous growth theory, starting from Romer (1986) and Lucas (1988), departs from the standard neoclassical theory and considers the technological change as endogenous. The theory focus on explaining the Solow residual.

Going back to the growth accounting approach, it is important to point out that it presents two major shortcomings: first of all, a key assumption is that prices coincide with social marginal products. If this assumption is violated, then the estimated Solow residual deviates from the true contribution of technological change to economic growth. Moreover, this approach ignores consideration on market power and returns to scale.

Hall and Jones (1996, 1997) suggest the cross-section growth accounting approach to TFP level comparisons and they follow Solow (1957) to arrive at the standard growth accounting equation. The difference with respect to Solow is that while in Solow (1957) differentiation is conducted in the direction of time t, Hall and Jones propose to apply the procedure in the cross-sectional direction, i.e. in the direction of i. But this poses a problem because the movement on i depends on the particular way the countries are ordered. Hall and Jones order the countries on the basis of an index that is a linear combination of the individual country's physical and human capital

per unit of labor and its value of α, the share of physical capital in income. In order to get the country specific α, the authors make the assumption that price of capital (r) is the same across countries.

The cross-section growth accounting approach presents several advantages. First, it does not require any specific form of aggregate production function. Only constant returns to scale and differentiability are required to arrive at the growth accounting equation. Second, it allows factor income share parameters to be different across countries. However, the cross-section growth accounting approach has some weaknesses too. First, it requires prior ordering of countries and TFP measurement may be sensitive to the ordering chosen. Second, TFP indices are also sensitive to inclusion or exclusion of countries. Third, computation of α_i is made on the basis of the assumption of a uniform rate of return across countries. Finally, using capital stock data and account for human capital in cross-country TFP comparison, it is possible to pick up some noise.

The panel approach to international TFP comparison arose directly from recent attempts at better explaining cross-country growth regularities. Islam (1995) takes the work of Mankiw, Romer and Weil (1992) as its starting point and examines how the results change with the adoption of the panel data approach. The main usefulness of the panel approach with respect to the single cross-country regressions lies in its ability to allow for differences in the aggregate production function across economies. These leads to results that are significantly different from those obtained from single cross-country regressions. The panel approach makes it possible to allow for differences in the aggregate production function in the form of unobservable individual "country effects". To the extent to which the "country effects" (intercepts) are correlated with the regressors, the conventional cross-section estimates of Mankiw, Romer and Weil (1992) are biased. Harrigan (1995) shows that there are systematic differences across countries in industry output. One possible explanation for this result is that technology is not the

same across countries. This hypothesis has gained great attention from international economists: Trefler (1993, 1995), Dollar and Wolff (1993) and Harrigan (1997a). More recently, Harrigan (1999) compute TFP for eleven OECD countries in the 1980s and he finds large and persistent TFP differences among them.

In comparison with the cross-section growth accounting approach, the panel regression approach has some advantage. First, it does not require any prior ordering of countries. Second, it is not sensitive to inclusion or exclusion of countries. Third, the approach is flexible to the use of capital stock data or investment data and to inclusion of human capital. Finally, the econometric estimation can provide a check for th severity of noise in the relevant data. Of course, the panel approach also presents some weaknesses: it requires a specific form for aggregate production function, it imposes homogeneity of factor share parameters and, finally, it is subject to certain pitfalls of econometric estimation.[2]

Here I use the panel data approach to estimate production functions because, as discussed above, it allows for differences across countries. The aim of this paper is to analyze the economic performance of the European regions. In particular, I use a Cobb-Douglas specification for a sample of 115 European Regions over the period 1976-2000. Moreover, the paper provides estimates of TFP for each region.

The paper also shows, on the basis of specific panel tests, that there is empirical evidence which suggests the presence of unit roots in the series under study. I apply, then, the panel cointegration test, proposed by Pedroni (1999), to guard against spurious regression problems.

The paper is organized as follows. Section 2 describes the model. Section 3 describes the econometric methodology. Section 4 presents the empirical results. Section 5 concludes. Tables are in appendix.

[2]The cost of econometric analysis is that parameter estimation requires imposing a statistical model on the data (see Harrigan,1999)

2 The model

I estimate the parameters of production functions and calculate total factor productivity for a sample of European regions from Cobb-Douglas production function specifications:

$$Y_{it} = A_{it} K_{it}^{\alpha} L_{it}^{\beta} \tag{4}$$

where Y_{it} is the value added in region i at time period t, K_{it} is the stock of physical capital, L_{it} is the amount of labour used in production. A_{it} is the specification for Hicks-neutral technology and it introduces a stochastic component into the model. The knowledge production function for region i at time period t can be defined as follows:in region i at time period t

$$A_{it} = e^{a_i + \gamma_t + \varepsilon_{it}} \tag{5}$$

where A_{it} is the level of technology in region i at time t, a_i denotes a region specific constant which captures the efficiency in technology production, γ_t is a common time effect which captures the countrywide or worldwide knowledge accumulation and ε_{it} is a random shock. The common time effect γ_t allows to take account of cross-regional dependence in the estimation of the regional production function.

Rewriting equation (4) in natural logarithms yields the following:

$$\ln Y_{it} = a_i + \gamma_t + \alpha \ln K_{it} + \beta \ln L_{it} + \varepsilon_{it} \tag{6}$$

The panel model includes a regional specific effect a_i and a common time effect γ_t. The parameters α and β are the elasticities of capital and labour with respect to output, respectively. This paper estimates 6 by using a panel data of 115 European regions over the period 1976-2000. The list of the regions is given in Appendix (tables (2), (3) and (4)).

The stock of physical capital is determined by using the Perpetual Inven-

tory Method :

$$K_t = (1-\delta)K_{t-1} + I_{t-1} \tag{7}$$

where δ is the depreciation rate: it is assumed constant and equal to 8%, which is consistent with OECD estimates; I is the gross fixed capital formation.[3] The initial value of K is calculate as:

$$K_0 = \frac{I_0}{g+\delta} \tag{8}$$

where g is the average annual logarithmic growth of investment expenditure and I_0 is investment expenditure in the first year for which data on investment are available.

3 Econometric methodology

Non-stationarity issues on series have been often overlooked when the panel approach has been used to estimate production functions. At the best of my knowledge, no attempt has been made to asses the non-stationarity of the series used on the estimation of production functions for European regions. Because of non-stationarity problems, first step of this work is to investigate the properties of regional time series for value-added, capital stock and labour.

I start applying the panel unit root test proposed by Im, Pesaran and Shin (2003, IPS hereafter), while the spurious regression problem is analyzed through the cointegration test recently proposed by Pedroni (1999).

3.1 Panel unit root tests

Over the past decade a number of important panel data set covering different countries, regions or industries over long time spans have become available.

[3] See Machin and Van Reenen, (1998)

This raises the issue of the plausibility of the dynamic homogeneity assumption that characterizes the traditional analysis of panel data models. The inconsistency of pooled estimators in dynamic heterogeneous panel models has been demonstrated by Pesaran and Smith (1995), and Pesaran et al.(1996).

Panel based unit root tests have been advanced by Quah (1990, 1994), Breitung and Meyer (1991), Levin and Lin (1992), Phillips and Moon (1999), Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003), among others. Quah uses the random field methods to analyze a panel with i.i.d. disturbances, and demonstrates that the Dickey-Fuller test statistic has a standard normal limiting distribution as both cross-section and time series dimensions grow arbitrarily large. Unfortunately, the random field method does not allow for individual specific effects. Breitung and Meyer approach allows for time specific effects and higher-order serial correlation, but cannot be extended to panel with heterogeneous errors. Levin and Lin test allows for heterogeneity only in the intercept and is based on the following model

$$\Delta y_{it} = \beta y_{i,t-1} + \alpha_{mi} d_{mt} + u_{it} \tag{9}$$

$i = 1, ..., N$; $t = 1, ...T$; $m = 1, 2, 3$

where d_{mt} contains deterministic variables; $d_{1t} = \{0\}, d_{2t} = \{1\}, d_{3t} = \{1, t\}$.

The Levin and Lin test requires the strong condition $N/T \to 0$ for its asymptotic validity. A revised version of Levin and Lin's (1992) earlier work is proposed by Levin, Lin and Chu (2002). The panel-based unit root test proposed in this paper allows for individual-specific intercepts, the degree of persistence in individual regression error and trend coefficient to vary freely across individuals. This test is relevant for panels of moderate size. However, this test has its limitations. First, there are some cases in which contemporaneous correlations cannot be removed by simply substracting cross-sectional averages. Secondly, the assumption that all individuals are identical with respect to the presence or absence of a unit root is in some sense restrictive.

Im, Pesaran and Shin (2003) propose unit root tests for dynamic heterogeneous panels based on the mean of individual unit root test statistics. In particular they propose a standardized t-bar test statistic based on the (augmented) Dickey-Fuller statistics averaged across the groups.

Consider a sample of N cross-section observed over T time periods. IPS suppose that the stochastic process, y_{it}, is generated by the first-order autoregressive process:

$$y_{it} = (1 - \phi_i)\mu_i + \phi_i y_{i,t-1} + \varepsilon_{it} \tag{10}$$

$i = 1, ..., N$, $t = 1, ..., T$,

where initial values, y_{i0}, are given. The null hypothesis of unit roots $\phi_i = 1$ can be expressed as

$$\Delta y_{it} = \alpha_i + \beta_i y_{i,t-1} + \varepsilon_{it} \tag{11}$$

where $\alpha_i = (1 - \phi_i)\mu_i$, $\beta_i = -(1 - \phi_i)$ and $\Delta y_{it} = (y_{it} - y_{i,t-1})$. The null hypothesis of unit roots then becomes

$$H_0 : \beta_i = 0 \tag{12}$$

for all i, against the alternatives

$H_1 : \beta_i < 0$, $i = 1, ..., N_1$, $\beta_i = 0$, $i = N_1 + 1, N_2 + 1, ..., N$.

This formulation of the alternative hypothesis allows for $\beta_i$to differ across groups, and is more general than the homogeneous alternative hypothesis, namely $\beta_i = \beta < 0$ for all i, which is implicit in the testing approaches of Quah and Levin-Lin.

The IPS group-mean t-bar statistic is given by:

$$t - bar_{NT} = N^{-1} \sum_{i=1}^{N} t_{iT_i}(p_i) \tag{13}$$

where t_{iT_i} is the individual t statistic for time series with different lag length.

3.2 Panel cointegration test

Methods for nonstationary panels have been gaining increased acceptance in recent empirical research. Initial theoretical work on nonstationary panels focused on testing for unit roots in univariate panels.[4] However, many applications involve multi-variate relationships and a researcher is interested to know whether or not a particular set of variables are cointegrated. Pedroni (1999) proposes a method to implement tests for the null of cointegration for the case with multiple regressors. The tests allow for a considerable heterogeneity among individual members of the panel.[5]

3.2.1 Testing for cointegration in heterogeneous panels: the multivariate case

Here I provide a complete description of the test proposed by Pedroni. The first step is to compute the regression residuals from the hypothesized cointegrating regression. The general case is:

$$y_{it} = \alpha_i + \delta_i t + \beta_{1i} X_{1_{it}} + \beta_{2i} X_{2_{it}} + ... + \beta_{Mi} X_{M_{it}} + e_{it} \tag{14}$$

for $t = 1, ..., T$; $\quad m = 1, ..., M$.

where T refers to the number of observation over time, N refers to the number of individual members in the panel, and M refers to the number of variables. The parameter α_i is the fixed effects parameter and β_{1i}, β_{2i},..., β_{Mi} are the slope coefficients. Both the fixed effects parameter and slope coefficients are allowed to vary across individual members. $\delta_i t$ represents a

[4] See for instance, Levin and Lin (1993) and Quah (1994).

[5] Pedroni cointegration tests include heterogeneity in both the long run cointegrating vectors as well as in the dynamics associated with short run deviations from these one.

deterministic time trend, which might be included in some applications.

To capture disturbances, which may be shared across the different members of the panel, common time dummies can be included.

Pedroni derives the asymptotic distributions of seven different statistics: four are based on pooling along the within-dimension, and three are based on pooling along the between-dimension. Pedroni calls the within-dimension based statistics as panel cointegration statistics, and the between-dimension based statistics as group mean panel cointegration statistics. The first of the panel cointegration statistics is a type of nonparametric variance ratio statistic. The second is a panel version of nonparametric statistic analogous to the Phillips and Perron rho-statistic. The third statistic is also nonparametric and analogous to the Phillips and Perron t-statistic. The fourth of the panel cointegration statistics is a parametric statistic analogous to the augmented Dickey-Fuller t-statistic.

The other three statistics are based on a group mean approach. The first and the second ones are analogous to the Phillips and Perron rho and t-statistic respectively, while the third one is analogous to the augmented Dickey-Fuller t-statistic.

Table (1) presents the seven statistics.

Pedroni panel cointegration test computes the seven statistics following a procedure in steps:

1. Estimate the panel cointegration regression (14) and collect residuals $\hat{e}_{it}$;
2. Estimate (14) in difference and collects residuals (η_{it});
3. Compute the long run variance of $\hat{\eta}_{it}$ using a kernel estimator, such as the Newey-West (1987) estimator, and calculate $\hat{L}_{11i}^{-2}$;
4. Use the residuals $\hat{e}_{it}$ and :

Table 1: Panel Cointegration Statistics

Panel Statistics (within)	
v	$T^2 N^{3/2} Z_{\hat{v}_{NT}} \equiv T^2 N^{3/2} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1}$
ρ	$T\sqrt{N} Z_{\rho_{NT}-1} \equiv T\sqrt{N} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} (\hat{e}_{it-1} \Delta\hat{e}_{it} - \hat{\lambda}_i)$
t nonparametric	$Z_{t_{NT}} \equiv \left(\tilde{\sigma}_{NT}^{2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} (\hat{e}_{it-1} \Delta\hat{e}_{it} - \hat{\lambda}_i)$
t (parametric)	$Z_{t_{NT}}^{*} \equiv \left(\tilde{s}_{NT}^{2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*} \Delta\hat{e}_{it}^{*}$
Group Statistics (between)	
ρ	$TN^{-1/2} Z_{\tilde{\rho}_{NT}-1} \equiv TN^{-1/2} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{t=1}^{T} (\hat{e}_{it-1} \Delta\hat{e}_{it} - \hat{\lambda}_i)$
t (nonparametric)	$N^{-1/2} \tilde{Z}_{t_{NT}}^{*} \equiv N^{-1/2} \sum_{i=1}^{N} \left(\hat{\sigma}_i^{2} \sum_{t=1}^{T} \hat{e}_{it-1}^{2} \right)^{-1/2} \sum_{t=1}^{T} (\hat{e}_{it-1} \Delta\hat{e}_{it} - \hat{\lambda}_i)$
t (parametric)	$N^{-1/2} \tilde{Z}_{t_{NT}}^{*} \equiv N^{-1/2} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{s}_i^{*2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum_{t=1}^{T} \hat{e}_{it-1}^{*} \Delta\hat{e}_{it}^{*}$

where $\hat{\lambda}_i = \frac{1}{T} \sum_{s=1}^{k_i} (1 - \frac{s}{k_i+1}) \sum_{t=s+1}^{T} \hat{\mu}_{it} \hat{\mu}_{it-s}$,
$\hat{s}_i^2 \equiv \frac{1}{T} \sum_{t=1}^{T} \hat{\mu}_{it}^2$,
$\hat{\sigma}_i^2 = \hat{s}_i^2 + 2\hat{\lambda}_i$,
$\tilde{\sigma}_{NT}^2 \equiv \frac{1}{N} \sum_{i=1}^{N} \hat{L}_{11i}^{-2} \hat{\sigma}_i^2$,
$\hat{s}_i^{*2} \equiv \frac{1}{T} \sum_{t=1}^{T} \hat{\mu}_{it}^{*2}$,
$\tilde{s}_{NT}^{*2} \equiv \frac{1}{N} \sum_{i=1}^{N} \hat{s}_i^{*2}$,
$\hat{L}_{11i}^{-2} = \frac{1}{T} \sum_{t=1}^{T} \hat{\eta}_{it}^2 + \frac{2}{T} \sum_{s=1}^{k_i} (1 - \frac{s}{k_i+1}) \sum_{t=s+1}^{T} \hat{\eta}_{it} \hat{\eta}_{it-s}$
and where $\hat{\mu}_{it}$, $\hat{\mu}_{it}^{*}$ and $\hat{\eta}_{it}$ are obtained from the following regressions:
$\hat{e}_{it} = \hat{\rho}_i \hat{e}_{it-1} + \hat{u}_{it}$, $\quad \hat{e}_{it} = \hat{\rho}_i \hat{e}_{it-1} + \sum_{k=1}^{K_i} \hat{\gamma}_{ik} \Delta\hat{e}_{it-k} + \hat{u}_{it}^{*}$,
$\Delta y_{it} = \sum_{m=1}^{M} \hat{b}_{mit} \Delta X_{mit} + \hat{\eta}_{it}$

a) compute the non parametric statistics estimating the following regression:

$$\hat{e}_{it} = \hat{\rho}_i \hat{e}_{it-1} + \hat{u}_{it}$$

The residuals $(\hat{u}_{it})$ are used to calculate the long run variance, denoted by $\hat{\sigma}_i^2$, while $\hat{s}_i^2$ is the simple variance of $\hat{u}_{it}$ and the term λ_i is calculated as $\lambda_i = \frac{1}{2}(\hat{\sigma}_i^2 - \hat{s}_i^2)$;

b) compute the parametric statistics estimating the following regression:

$$\hat{e}_{it} = \hat{\rho}_i \hat{e}_{it-1} + \sum_{k=1}^{K_i} \hat{\gamma}_{ik} \Delta \hat{e}_{it-k} + \hat{u}_{it}^*$$

and use the residuals $(\hat{u}_{it}^*)$ to compute the simple variance $\hat{s}_i^{*2}$.

Pedroni (1995, 1997a) shows that each of the seven statistics presented in table (1) will be distributed as standard normal after an appropriate standardization. This standardization depends only on the moments of certain Brownian motion functionals.[6] In Pedroni (1999) the moments of the vector of Brownian motion functionals are computed by Monte Carlo simulation for the case of multiple regressors.

The asymptotic distributions for each of the seven panel and group mean statistics can be expressed in the form

$$\frac{\varkappa_{NT} - \mu\sqrt{N}}{\sqrt{\nu}} \rightarrow N(0,1)$$

where $\varkappa_{NT}$ is the standardized form of the statistics as described in table (1), and the value for μ and ν are functions of the moments of Brownian motion functionals.

[6] A Brownian motion is a continuous-time stochastic process with three important properties. First, it is a Markov process and it means that the probabilty distribution for all future values of the process depends only on its current value. Second, the Brownian process has indipendent increments. Finally, changes in the process over any finite interval of time are normally distributed.

3.3 Panel estimation of long-run relationship

The main theme of this paper is to analyse the economic performance of a sample of European regions. But it is worth emphasising that only if the cointegration test provides evidence of long run dynamics in the series, although they are nonstationary, it is possible to proceed with the analysis. I have in mind a particular form of normalization among variables (a production function relation) and in this case, as pointed out by Pedroni, the interest is in knowing whether the variables are cointegrated, not how many cointegrating vectors exist.

The model I use is a two error component model, with $u_{it} = a_i + \gamma_t + \varepsilon_{it}$, and ε_{it} is assumed homoskedastic. If the assumption fails, the estimates are still consistent but inefficient. It is possible investigate about the validity of this assumption by performing a groupwise likelihood ratio heteroskedasticity test. This test is performed on the residuals of the model estimated by OLS. The test is chi-squred distributed with $N - 1$ degrees of freedom, where N is the number of groups in the sample

Baltagi and Li (1995) suggest an LM test for serial correlation in fixed effects models. They propose two version of the test, depending on the assumption for the autocorrelation structure, namely AR(1) and MA(1). The test is asymptotically distributed as $N(0, 1)$ under the null.

4 Data and empirical results

In my analysis I use a panel of 115 European regions over the period 1976-2000. Annual data on value added and labour units are from Cambridge Econometrics dataset. The stock of capital is determined by using the Perpetual Inventory Method and is measured at 1995 constant prices, as value added.

I analyze the time series properties of my data, applying the IPS panel root test to control for stationarity of the three variables included in the panel

used to estimate the production function. Table (5) report the results of the test for the logarithm of value added (Y), capital stock (K) and labour (L). The test is performed both on levels and first differences ($\Delta Y, \Delta K, \Delta L$) of the variables. The null hypothesis refers to nonstationarity behavior of the time series, connection admitting the possibility that the error terms are serially correlated with different serial correlation coefficients in cross-sectional units. Under the null of nonstationarity the test is distributed as $N(0,1)$, so that large negative numbers means stationarity.

The test is performed with constant but not trend ($t-bar$), constant and heterogeneous trend ($t-bar^*$) in the test regression. I introduce up to five lags of the dependent variable for serial correlation in the errors.

Table (5) show the $t-bar$ and the $t-bar^*$statistics values. The variables are integrated of order one or $I(1)$ process: they are nonstationary in levels but are stationary in first differences.[7]

Because of nonstationarity of the series, next step of this work is to determine if all three variables are cointegrated in order to avoid spurious regression problem. In the absence of cointegration I can simply first difference the data and work with these transformed variables. However, in the presence of cointegration the first differences do not capture the long run relationships in the data.

The cointegrating regression that I estimate is

$$\ln Y_{it} = a_i + \gamma_t + \alpha_i \ln K_{it} + \beta_i \ln L_{it} + \varepsilon_{it} \tag{15}$$

so that each region has its own relationship among Y_{it}, gross value added, K_{it}, capital stock, and L_{it}, total employment. The variable ε_{it} represents a stationary error term . Table (6) presents the results of cointegration test on (15) with a lag length of up to 5 years in order to check the robustness of results with respect to different structure dynamics. The slopes (α_i, β_i) of the cointegrating relationship are allowed to vary across regions. The com-

[7]The exact critical values of the t-bar statistic are given in IPS (2003)

mon time factor γ_t, capture any common effects that would tend to cause the individual region variables to move together over time. These may be short term business cycle effects or longer run effects. All reported values are *normally distributed* under null of no cointegration. Panel statistics are weighted by long run variances. Under the alternative hypothesis, the panel variance statistic diverge to positive infinity, and consequently large positive values imply that the null of no cointegration is rejected. To the contrary, the other six statistics diverge to negative infinity under the alternative hypothesis and large negative values imply that the null of cointegration is rejected.

The results suggest that the null of no cointegration is rejected by five out of seven statistics: only panel rho and group rho statistics do not reject the null hypothesis. Except for panel rho and group rho statistics, it is worth noting that the statistics are highly significant even at lower lags. Test results provide evidence in favour of a long-run production function relationship.

Table (7) presents a groupwise likelihood ratio eteroskedasticity test performed on the residuals of the production function estimates by fixed effects. The test is chi-squred distributed with $N-1$ degrees of freedom, where N is the number of groups in the sample (115 in my case). The null hypothesis of homoskedasticity is rejected.

Table (8) presents the two versions of the Baltagi and Li (1995) test for serial correlation in fixed effects models. The test presents two alternative specifications for autocorrelation in the errors: $AR(1)$ and $MA(1)$. Under both assumptions, the null hypothesis of no serial correlation is rejected.

Test results justify the adoption of a GLS fixed effect estimator, in order to control for region unobservables and to correct for heteroskedasticity across regions and residual serial correlation. The common time factor instead captures the contemporaneous correlations across regions.

I first estimate the (6) for all sample. I do not impose the assumption of constant returns to scale: the production function can display increasing, constant, or decreasing returns to scale as $\alpha+\beta$ is greater than, equal to,

or less than one, respectively. Table (9), presents in the first column the results of a two-dimension panel where individuals are represented by 115 European Regions over the period 1976-2000. I use a fixed effects GLS model accounting for heteroskedasticity and serial correlation. Time dummies are included in the specification to capture disturbances which may be shared across the different regions. These may be business cycle effects or long run effects such as changes in technology.

The coefficient of capital stock (0.39) is very close to the findings of the accounting approach, where it is found in the range [0.35,0.38].

Results show that the production function for European regions exhibits increasing returns to scale and this may suggest a dynamic and innovative production organization on the European scene.

Finally, I divide the sample into sub-samples (seven), grouping regions on the basis of the country they belong (see tables (9) and (10).[8] The results point out a considerable degree of heterogeneity: the capital and labour elasticities range from low values for some countries to high values for some others.

From the estimated fixed effects I calculate the antilogarithms, which represents the parameter of technological efficiency for each region (see table in appendix, (11), (12), (13). The results show remarkable differences among regions in the technological knowledge levels. In particular, Corse, Alsace, Haute-Normandie (France) and Salzburg, Vorarlberg (Austria) exhibit the most high levels of TFP. On the other hand, the lowest parameters are those of regions of Greece, Spain and United Kingdom. Looking at the results for Italian regions, the highest values are those of the northern regions; the leader region is Valle d'Aosta, with a technological parameter of (4.14). On the other hand, the lowest value are those for southern regions, with Puglia exhibiting the lowest parameter (2.24). This finding broadly confirms the

[8] The sub-samples are for german (DE), british (UK), french (FR), italian (IT), spanish (ES), greek (GR) and austrian (AT) regions.

long-lasting dualism between North and South.

5 Conclusion

This paper has analyzed the economic performance of a sample of European Regions. It has provided estimates of Cobb-Douglas production functions over the period 1976-2000. The sample was composed by 115 European Regions of 12 Countries: Austria, Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Luxembourg, Nethelands, Spain, United Kingdom. Great attention has been devoted to the estimation procedures. Because problems of non-stationarity may arise when panel data approach is used to estimate production function, first step of this work has been to investigate the properties of regional time series for value added, capital stock and labour. The presence of unit roots in the series has been found and, consequently, I applied panel cointegration tests to guard against the spurious regression problem. It has been clearly shown that in the given panel all the variables share long-run relationship and this imply evidence in favour of a long-run production function relationship.

I have reported results for a fixed effects GLS estimator to take account of heteroskedasticity and serial correlation.

I have found a coefficient for capital stock very close to the findings of the accounting approach.

This paper also reports the estimated Total Factor Productivity for each region. The results show remarkable differences among regions in the technological knowledge levels. In particular, some regions of France and Austria exhibit the most high levels of TFP. On the other hand, the lowest parameters are those of regions of Greece and Spain.

Looking at the results for Italian regions, the highest values are those of the northern regions; the leader region is Valle d'Aosta, with a technological parameter of (4.14). On the other hand, the lowest value are those for

southern regions, with Puglia exhibiting the lowest parameter (2.24). This finding confirms the well-known dualism between North and South of Italy.

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6 Appendix : Data and tables

Data are mainly from Cambridge Econometrics, a validated database of economic indicators for cities and regions. The database draws on the available official data at European and national levels and has undergone a substantial process of updating and quality checks to improve its consistency, timeliness and coverage. The current database includes output, employment, household expenditure, investment expenditure, demographic indicators (total and working population).

The regions in the sample are presented in tables (2), (3) and (4)

Y Gross Value Added in constant prices (base year 1995)

L Total Employment (source: Cambridge Econometrics)

K Real capital stock. The calculation of the capital stock is made according to the Perpetual Inventory Method. Data on investment expenditure are from Cambridge Econometrics

A Total Factor Productivity (TFP). It is obtained by estimating production functions by performing a GLS fixed effects estimator.

Table 2: Sample I
Regions

Bruxelles-Brussel (Be)	Voreio Aigaio (Gr)
Vlaams Gewest (Be)	Notio Aigaio (Gr)
Region Walonne (Be)	Kriti (Gr)
Denmark	Galicia (Es)
Baden-Wurttemberg (De)	Asturias (Es)
Bayern (De)	Cantabria (Es)
Berlin (De)	Pais Vasco (Es)
Bremen (De)	Navarra (Es)
Hamburg (De)	Rioja (Es)
Hessen (De)	Aragon (Es)
Niedersachsen (De)	Madrid (Es)
Nordrhein-Westfalen (De)	Castilla-Leon (Es)
Rheinland-Pfalz (De)	Castilla-la Mancha (Es)
Saarland (De)	Extremadura (Es)
Schleswig-Holstein (De)	Cataluna (Es)
Anatoliki Makedonia (Gr)	Com. Valenciana (Es)
Kentriki Makedonia (Gr)	Baleares (Es)
Dytiki Makedonia (Gr)	Andalucia (Es)
Thessalia (Gr)	Murcia (Es)
Ipeiros (Gr)	Ceuta y Melilla (Es)
Ionia Nisia (Gr)	Canarias (Es)
Dytiki Ellada (Gr)	Ile de France (Fr)
Sterea Ellada (Gr)	Champagne-Ard (Fr)
Peloponnisos (Gr)	Picardie (Fr)
Attiki (Gr)	Haute-Normandie (Fr)

Table 3: Sample II

Regions	
Centre (Fr)	Fr.-Venezia Giulia (It)
Basse-Normandie (Fr)	Emilia-Romagna (It)
Bourgogne (Fr)	Toscana (It)
Nord-Pas de Calais (Fr)	Umbria (It)
Lorraine (Fr)	Marche (It)
Alsace (Fr)	Lazio (It)
Franche-Comte (Fr)	Abruzzo (It)
Pays de la Loire (Fr)	Molise (It)
Bretagne (Fr)	Campania (It)
Poitou-Charentes (Fr)	Puglia (It)
Aquitaine (Fr)	Basilicata (It)
Midi-Pyrenees (Fr)	Calabria (It)
Limousin (Fr)	Sicilia (It)
Rhone-Alpes (Fr)	Sardegna (It)
Auvergne (Fr)	Luxembourg
Languedoc-Rouss. (Fr)	Noord-Nederland (Nl)
Prov-Alpes-Cote d'Azur (Fr)	Oost-Nederland (Nl)
Corse (Fr)	West-Nederland (Nl)
Ireland	Zuid-Nederland (Nl)
Piemonte (It)	Burgenland (At)
Valle d'Aosta (It)	Niederosterreich (At)
Liguria (It)	Wien (At)
Lombardia (It)	Karnten (At)
Trentino-Alto Adige (It)	Steiermark (At)
Veneto (It)	Oberosterreich (At)

Table 4: Sample III

Regions
Salzburg (At)
Tirol (At)
Vorarlberg (At)
North East (GB)
North West (GB)
Yorkshire and the Humb (GB)
East Midlands (GB)
West Midlands (GB)
Eastern (At)
London (GB)
South East (GB)
South West (GB)
Wales (GB)
Scotland (GB)
Northern Ireland (GB)

Table 5: Panel Unit Root Test

Variables	$t - bar$	$t - bar^*$
Y	0.95 (0.83)	−1.20 (0.12)
K	0.43 (0.67)	1.45 (0.93)
L	0.27 (0.61)	−4.07 (0.00)
ΔY	−8.87 (0.00)	−4.82 (0.00)
ΔK	−2.43 (0.01)	−1.77 (0.04)
ΔL	−8.48 (0.00)	−2.63 (0.00)

Notes: p-values are in brackets. All variables are in logs.
The test statistics are asymptotically distributed as N(0,1) under the null hypothesis of non-stationarity.

Table 6: Panel Cointegration Test

lags	1	2	3	4	5
panel v-stat	2.20**	2.20**	2.20**	2.20**	2.20**
panel rho-stat	−0.20	−0.20	−0.20	−0.20	−0.20
panel pp-stat	−2.76*	−2.76*	−2.76*	−2.76*	−2.76*
panel adf-stat	−4.08*	−4.10*	−4.15*	−3.38*	−3.43*
group rho-stat	2.84	2.84	2.84	2.84	2.84
group pp-stat	−1.88**	−1.88**	−1.88**	−1.88**	−1.88**
group adf-stat	−5.33*	−6.28*	−6.34*	−5.11*	−5.65*

The test statistics are distributed as N(0,1) under the null hypothesis of no co-integration. *, **, *** represent the rejection of null hypothesis at 1%, 5%, and 10% significance level. The critical values for 1%, 5%, and 10% level are -2.328, -1.645, and -1.285, respectively.

Table 7: Test for groupwise heteroskedasticity

GH Test	$\chi^2_{(114)} = 18037.50$
P-value	0.00

The test is χ^2 distributed with $N-1$ degrees of freedom. The null hypothesis of homoskedasticity is rejected

Table 8: Test for groupwise heteroskedasticity

LM Test, AR(1) $v_{it}=\rho v_{it-1}+\varepsilon_{it}$	$\chi^2_{(1)} = 2038.19$ ($p-value \simeq 0.000$)
$H_0 : \rho = 0$	
LM_5Test, MA(1) $v_{it}=\varepsilon_{it}+\lambda\varepsilon_{it-1}$	$N(0,1) = 45.15$ ($p-value \simeq 0.000$)
$H_0 : \lambda = 0$	

Table 9: Production Function Estimate

Dependent Varable:Y_{it}	All Sample	DE	UK	FR
K_{it}	0.39 (0.01)	0.37 (0.07)	0.14 (0.04)	0.10 (0.05)
L_{it}	0.73 (0.02)	0.89 (0.11)	0.54 (0.05)	0.60 (0.05)
Year Dummies	yes	yes		yes
Fixed Effects	115	11	12	22
N.obs	2875	275	330	550

The estimation method is a feasible fixed effect GLS estimator, accounting for heteroskedasticity and serial correlation. Standard errors are in brackets.

Table 10: Production Function Estimate

Dependent Varable:Y_{it}	IT	ES	GR	AT
K_{it}	0.44 (0.08)	0.27 (0.04)	0.54 (0.05)	0.15 (0.04)
L_{it}	0.19 (0.03)	0.54 (0.03)	0.17 (0.03)	0.06 (0.06)
Year Dummies	yes	yes	yes	yes
Fixed Effects	20	18	13	9
N.obs	550	450	325	225

See note on table 9.

Table 11: Total Factor Productivity I

Regions	TFP	Regions	TFP
Bruxelles-Brussel	3.69	Voreio Aigaio	2.92
Vlaams Gewest	3.20	Notio Aigaio	3.29
Region Walonne	3.11	Kriti	2.59
Denmark	3.04	Galicia	1.71
Baden-Wurttemberg	2.74	Asturias	2.25
Bayern	2.66	Cantabria	2.55
Berlin	2.68	Pais Vasco	2.38
Bremen	3.59	Navarra	2.84
Hamburg	3.60	Rioja	2.77
Hessen	3.04	Aragon	2.34
Niedersachsen	2.77	Madrid	2.32
Nordrhein-Westfalen	2.59	Castilla-Leon	2.00
Rheinland-Pfalz	3.00	Castilla-la Mancha	2.12
Saarland	3.36	Extremadura	2.09
Schleswig-Holstein	3.05	Cataluna	2.09
Anatoliki Makedonia	2.30	Com. Valenciana	2.04
Kentriki Makedonia	2.26	Baleares	2.82
Dytiki Makedonia	2.94	Andalucia	1.95
Thessalia	2.32	Murcia	2.29
Ipeiros	2.34	Ceuta y Melilla	2.94
Ionia Nisia	2.62	Canarias	2.47
Dytiki Ellada	2.14	Ile de France	3.19
Sterea Ellada	3.60	Champagne-Ard.	3.44
Peloponnisos	2.49	Picardie	3.47
Attiki	2.21	Haute-Normandie	3.61

TFP is estimated over the period 1976-2000.

Feasible fixed effects GLS.

Table 12: Total Factor Productivity II

Regions	TFP	Regions	TFP
Centre	3.31	Fr.-Venezia Giulia	2.91
Basse-Normandie	3.24	Emilia-Romagna	2.61
Bourgogne	3.45	Toscana	2.59
Nord-Pas de Calais	3.08	Umbria	3.02
Lorraine	3.27	Marche	2.73
Alsace	3.74	Lazio	2.65
Franche-Comte	3.58	Abruzzo	2.76
Pays de la Loire	3.06	Molise	3.21
Bretagne	3.00	Campania	2.27
Poitou-Charentes	3.24	Puglia	2.24
Aquitaine	3.28	Basilicata	3.04
Midi-Pyrenees	3.20	Calabria	2.49
Limousin	3.41	Sicilia	2.52
Rhone-Alpes	3.07	Sardegna	2.78
Auvergne	3.27	Luxembourg	3.98
Languedoc-Rouss.	3.36	Noord-Nederland	3.92
Prov-Alpes-Cote d'Azur	3.39	Oost-Nederland	2.67
Corse	4.37	West-Nederland	2.61
Ireland	2.67	Zuid-Nederland	2.84
Piemonte	2.66	Burgenland	4.28
Valle d'Aosta	4.14	Niederosterreich	3.59
Liguria	2.91	Wien	3.65
Lombardia	2.56	Karnten	3.63
Trentino-Alto Adige	3.34	Steiermark	3.46
Veneto	2.62	Oberosterreich	3.56

See note on table 11

Table 13: Total Factor Productivity III

Regions	TFP
Salzburg	4.29
Tirol	4.14
Vorarlberg	4.42
North East	2.17
North West	1.90
Yorkshire and the Humb	1.97
East Midlands	2.05
West Midlands	1.90
Eastern	2.33
London	1.93
South East	2.12
South West	2.06
Wales	2.15
Scotland	2.07
Northern Ireland	2.21

See note on table 11

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