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**Working Paper**

## How to Measure the Unobservable: A Panel Technique for the Analysis of TFP Convergence

Nota di Lavoro, No. 16.2005

**Provided in Cooperation with:**

Fondazione Eni Enrico Mattei (FEEM)

*Suggested Citation:* Di Liberto, Adriana; Mura, Roberto; Pigliaru, Francesco (2005) : How to Measure the Unobservable: A Panel Technique for the Analysis of TFP Convergence, Nota di Lavoro, No. 16.2005, Fondazione Eni Enrico Mattei (FEEM), Milano

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**How to Measure the Unobservable:  
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NOTA DI LAVORO 16.2005

**JANUARY 2005**

KTHC - Knowledge, Technology, Human Capital

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# How to Measure the Unobservable: A Panel Technique for the Analysis of TFP Convergence

## Summary

This paper proposes a fixed-effect panel methodology that enables us to simultaneously take into account both TFP convergence and the traditional neoclassical-type of convergence. We analyse a sample of Italian regions between 1963 and 1993 and find strong evidence that both mechanisms were at work during the process of aggregate regional convergence observed in Italy up to the mid-seventies. Finally, we find that our TFP estimates are highly positively correlated with standard human capital measures, where the latter is not statistically significant in growth regressions. This evidence confirms one of the hypotheses of the Nelson and Phelps approach, namely that human capital is the main determinant of technological catch-up. Our results are robust to the use of different estimation procedures such as simple LSDV, Kiviet-corrected LSDV, and GMM *à la* Arellano and Bond.

**Keywords:** TFP, Panel data, Regional convergence

**JEL Classification:** O47, O33, O18, C23

*We gratefully acknowledge financial support from MIUR Cofin “I fattori che influenzano i processi di crescita economica e di convergenza tra le regioni europee”. We would like to thank Pasquale Scaramozzino, Mark Rogers, Paul Cheshire, Anna Soci, Bernard Fingleton, Wendy Carlin and the participants at the ERSA Conference (Jyvaskila, 2003) and EEFS Conference (Bologna, 2003) for their helpful comments and suggestions.*

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## 1. Introduction

In early studies on growth differentials across countries, technology was regarded as a pure public good, freely available to all, with heterogeneity in technological levels ruled out by assumption. More recent papers show how misleading such an assumption can be: indeed, differences in total factor productivity (TFP) levels are often estimated as a major component of the observed large cross-country differences in per capita income.<sup>1</sup> Typically, these studies show that a large part of cross-country per capita income differences is left unexplained after taking account of differences in physical or human capital. Remarkable differences in TFP have also been detected in highly integrated areas or across regions of a single country<sup>2</sup> and, nowadays, few economists would dispute these findings<sup>3</sup>.

More controversial is the question of whether such differences in TFP are constant or not – that is, whether TFP convergence is taking place, at what speed and under what conditions. In fact, observed convergence may be the result of three different mechanisms: convergence due to capital accumulation, convergence due to technology transfer (catch up), and convergence due to both. Indeed, as Bernard and Jones (1996) put it, we often do not know “how much of the convergence that we observe is due to convergence in technology versus convergence in capital-labour ratios”<sup>4</sup>. Both theoretically and empirically, the problem is therefore to find a methodology able to discriminate between these three hypotheses.

Recently, things have improved on both the analytical and on the empirical side. On the

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<sup>1</sup> Among the most influential, see Klenow and Rodriguez Clare (1997) and Hall and Jones (1999). The role of TFP heterogeneity in cross-country analysis is also stressed in Parente and Prescott (2000), Easterly and Levine (2001), and Lucas (2000), among many others.

<sup>2</sup> Using a sample of 101 EU regions Boldrin and Canova (2001) find that per capita GDP is much more correlated with their measure of TFP than with capital-labour ratios. See also Aiello and Scoppa (2000), and Marrocu, Paci and Pala (2001) for the Italian regions, and De la Fuente (2002) for the Spanish regions.

<sup>3</sup> For a different viewpoint on the role of TFP, see Young [(1994) and (1995)] and, more recently, Baier, Dwyer and Tamura (2002). Using growth accounting techniques in a sample of 145 countries Baier *et al.* (2002) find that TFP growth is an unimportant part of average output growth (about 8%).

<sup>4</sup> This problem remains unsolved in another stream of research in a non neoclassical tradition, where technology diffusion is regarded as the crucial source of convergence [for instance, Dowrick and Nguyen (1989) and Fagerberg

analytical side, simple models in which technology-driven TFP convergence and capital-deepening can be studied within a common framework are now available. In these models, transitional dynamics is simple enough to be useful for empirical analysis.<sup>5</sup> On the empirical side, existing literature has identified a number of different methodologies to measure TFP at different points in time. In general, in this case difficulties arise since measuring TFP levels is not an easy task, and measuring it at different points in time is even more difficult, given the current availability of data in most of the existing cross-country and cross-region datasets. Within the available empirical literature it is possible to identify three different approaches. The first approach introduces per capita (or per worker) GDP levels in a standard growth-convergence regression, or a combination of GDP and other variables, as a proxy for technology. In this case the estimated coefficient on this variable is interpreted as a technological convergence coefficient. This methodology was used in two pioneering works on this subject, Dowrick and Nguyen (1989) and Benhabib and Spiegel (1994). However, early studies have been criticised<sup>6</sup> owing to the fact that the coefficient on this variable may capture the effect of diminishing returns as well as that of technological diffusion, so that distinguishing between them is something beyond the reach of this technique. This approach has been recently further developed by Dowrick and Rogers (2002): the use of data on capital per worker enables these authors to control for possible observational equivalence problems. A second approach computes technology levels as a residual once the contribution of factors of production to per capita GDP has been taken into account<sup>7</sup>. Here technology is measured indirectly, as the residual component of GDP growth that cannot be explained by the growth of the assumed inputs of production. Growth (or level) accounting techniques have been criticised since they require the imposition of too many assumptions such as cross-country parameter homogeneity and the absence of externalities, and may easily attribute to capital accumulation something that should be attributed

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and Verspagen (1996)]. Here the whole observed convergence is typically assigned to the catch-up mechanism, and other mechanisms (such as capital-deepening) are neglected on a priori grounds, rather than tested.

<sup>5</sup> See for instance, De la Fuente (1997) and (2002), and Pigliaru (2003).

<sup>6</sup> See De la Fuente (1995).

<sup>7</sup> See Klenow and Rodriguez-Clare (1997), Hall and Jones (1999) and Aiyar and Feyrer (2002).

to technological progress and *vice-versa*<sup>8</sup>. In particular, by assuming the absence of externalities this methodology tends to ascribe too much of the observed income variation to differences in productivity.

In this paper we build upon a third methodology based on Islam (2000) in which the presence of TFP heterogeneity in cross-country convergence analysis is tested by using an appropriate fixed-effects panel estimator. Originally, this methodology was designed to measure cross-country convergence in capital-labour ratios while controlling for TFP (or technology) heterogeneity<sup>9</sup>. In particular, we show that the same methodology can be extended to analyse cases in which TFP differences in levels are *not* constants, and therefore might be converging.<sup>10</sup> The methodology we use is as follows. First, we use data on GDP per worker to estimate the standard convergence equation with a fixed effects estimator over two sub-periods. Second, we use the values of the individual intercepts to compute our regional TFP levels. The robustness of our results is assessed by comparing the estimates obtained using different estimators – namely, a Least Square with Dummy Variable (LSDV) estimator, a biased-corrected LSDV estimator and a GMM (Arellano-Bond) estimator. Third, we analyse the two TFP series to test whether the observed pattern over time is consistent either with the technological catch-up hypothesis or with the hypothesis that the current degree of TFP heterogeneity is at its stationary value.

Finally, note that direct estimates of TFP enable us to test for one of the main hypothesis in catching-up literature and, in particular, of the Nelson and Phelps (1966) approach. In these models TFP growth is determined by the technological distance from the leader and by the level of human capital, where the latter influences the capacity for both discovering new technologies and adopting innovations from abroad. In other words, human capital levels determine the capacity of adopting new technologies from abroad and, thus, the possibility of a catch up process among countries. As a

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<sup>8</sup> This is certainly true if capital is endogenous and responds to technological progress or if improvements in educational attainment have indirect effects on output through changes in labour force participation or R&D and growth of TFP. On this, see in particular Barro and Sala-i-Martin (2004) pp.457-60, Temple (2001) and Islam (2000) for a direct comparison between growth accounting and panel methodologies.

<sup>9</sup> See Islam (1995) and Caselli et al. (1996), among others.

consequence, we should expect estimated TFP levels to be positively correlated with human capital stocks.

Our case-study is Italy and its persistent regional divide. Thus, we follow the Klenow and Rodriguez-Clare (1997) suggestion of using within country data bearing on the process of technology diffusion<sup>11</sup> and use a panel dataset of Italian regions, 1963-93. Note that, from a methodological point of view, the use of regional data has two main advantages. First, various unobservable components are supposed to be far more homogeneous across regions than across countries. This feature makes the interpretation of fixed-effects as technology (or more broadly TFP) in panel regressions far easier, since a number of components such as culture and institutions are necessarily more homogeneous across regions than across countries. Second, data comparability is also easier. Consider human capital, a crucial variable for convergence analysis. One of the main criticism with cross-country datasets is that direct comparison of the data is not viable due to divergent quality in schooling institutions. The use of a regional dataset allows us to limit this type of problem. At the same time, unlike in many other regional samples, Italian data usually reveals a remarkable degree of regional heterogeneity in variables such as per capita income levels and human capital stocks,<sup>12</sup> and we have access to fairly long time series data. These characteristics render our sample best suited for this type of empirical analysis.

The remaining part of the paper is organized as follows. In section 2 we briefly review the literature on Italian regional inequalities. In section 3 we discuss how a panel data technique can be used to test for both the presence of TFP heterogeneity and convergence. In section 4 we discuss how to select the estimator that suits our case better. Section 5 presents our evidence on the degree of cross-region TFP heterogeneity, while section 6 shows how much TFP convergence can be detected in our dataset. Finally, in section 7 we test if our regional TFP estimates are positively correlated with the observed regional human capital endowments. Conclusions are in section 8.

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<sup>10</sup> See also Islam (2000).

<sup>11</sup> "...we think the insights gleaned from cross-country regressions have run into sharply diminishing returns. We would like to see more detailed country analysis a la Young (1995)...." Klenow and Rodriguez-Clare (1997) p.614.

## 2. Regional inequality in Italy: a brief outline

Although the Italian case is one of the best known and most analysed cases of a regional divide, the existing papers have usually introduced the standard “classical approach<sup>13</sup>” to convergence analysis<sup>14</sup>, while studies that focus on the role of TFP<sup>15</sup> do not examine how the cross-region TFP distribution evolves over time. In other words, our paper provides the first explicit analysis of TFP convergence across Italian regions.

We start with a brief summary of the main stylised facts about regional inequality and convergence in Italy. The degree of regional inequality in Italy is among the highest across all the EU countries.<sup>16</sup> Such high inequality reflects the old and persistent North-South divide within the country. Stylised facts show that the per capita income distances between top and bottom regions during the 1960s were remarkable. Valle d’Aosta, was 42% (and Lombardia 38%) wealthier than in the average Italian region. On the other hand, the poorest regions, Calabria and Basilicata, had a 38% disadvantage with respect to the Italian average. This is a very large regional gap by any European standard. This gap has diminished during the last three decades but the decrease was neither persistent nor uniform, and came to a halt in the mid-1970s. Relevant literature has put forward various explanations for this. There was a decrease in migration from the South to the North following a more uniform national wage rate imposed by law in 1969. Moreover, there were radical changes in the policies aimed at fostering the development of the more backward regions. In particular, the Italian Government’s efforts to boost industrial investment in the South during the 60s and part of the 70s is well documented<sup>17</sup>. After that period, there was a shift in policy from investments to income maintenance in the form of direct transfers and by means of a significant

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<sup>12</sup> Paci and Pigliaru (1995), Di Liberto (2001), Boltho, Carlin and Scaramozzino (1999), among several others.

<sup>13</sup> See Sala-i-Martin (1996).

<sup>14</sup> See Mauro and Podrecca (1994), Di Liberto (1994), Boltho, Carlin and Scaramozzino (1999), Paci and Pigliaru (1995) among others.

<sup>15</sup> See Aiello and Scoppa (2000), and Marrocu, Paci and Pala (2001).

<sup>16</sup> See Barro and Sala-i-Martin (2004).

<sup>17</sup> On this issue see Graziani (1978).



expansion of the Public Sector. Finally, the rapid increase of oil prices in 1973-74 is thought to have influenced investment patterns, technology and additional factors that may have affected the convergence process internationally.<sup>18</sup> Whatever the reason, the pre- and post-1970 pattern is strongly confirmed by numerous studies.

### 3. A Panel Data approach to estimate TFP convergence

Our aim is to test for the presence of TFP heterogeneity and convergence in cross-country convergence analysis by using an appropriate fixed-effect panel estimator. As said above, TFP levels are measured by means of two main methodologies, namely the level accounting approach and the GDP-proxy approach. Another approach to control for possible cross-country differences in technology is to test for the presence of technology heterogeneity in cross-country convergence analysis by using an appropriate fixed-effect panel estimator<sup>19</sup>. Islam (1995) was among the first to suggest this econometric solution to the problem of estimating TFP levels. In particular, he extends the standard Mankiw Romer and Weil (1992) structural approach (hereafter MRW) by allowing TFP levels to vary across individual economies, together with saving rates and population growth rates. The convergence equation is given by:

$$\ln Y_{it_2} = (1-\beta) \frac{\alpha}{1-\alpha} \ln(s_{it_1}) - (1-\beta) \frac{\alpha}{1-\alpha} \ln(n_{it_1} + g + \delta) + \beta \ln Y_{it_1} + (1-\beta) \ln A_{t_0} + g(t_2 - \beta t_1) \quad (1)$$

where  $Y_{it_1}$  is per capita GDP in economy  $i$  at time  $t_1$  (initial period, while  $t_2$  is the final one), and  $s$ ,  $n$ ,  $\delta$  and  $g$  are, respectively the saving rate, the population growth rate, the depreciation rate, and exogenous technological change; the latter is assumed to be invariant across individual economies. Moreover,  $\alpha$  is the usual capital share of a standard Cobb-Douglas production function.

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<sup>18</sup> Indeed, stops and goes in regional convergence have been detected in several other countries. See de la Fuente (1997) for Spanish regions and Sala-i-Martin (1996) for examples from other OECD countries.

<sup>19</sup> See Islam (1995) and (2000) and Caselli Esquivel and Lefort (1996) among others.

Finally,  $\beta \equiv e^{-\lambda\tau}$ , where  $\lambda = (1-\alpha)(n+g+\delta)$  represents the convergence parameter and  $\tau \equiv t_2 - t_1$ , is the time span considered.

Unlike in the MRW approach, Islam introduces the idea that the unobservable differences in TFP are correlated with other regressors<sup>20</sup>, and uses suitable panel techniques to estimate:

$$y_{it} = \beta y_{it-1} + \sum_{j=1}^2 \gamma_j x_{j,it} + \eta_t + \mu_i + v_{it}, \quad j=1,2 \quad (2)$$

where the dependent variable is the logarithm of per capita GDP (measured in terms of population working age),  $v_{it}$  is the transitory term that varies across countries, and the remaining terms are:

$$x_{it}^1 = \ln(s_{it}) \quad (3)$$

$$x_{it}^2 = \ln(n_{it} + g + \delta) \quad (4)$$

$$\gamma = (1-\beta) \frac{\alpha}{1-\alpha} \quad (5)$$

$$\mu_i = (1-\beta) \ln A(0)_i \quad (6)$$

$$\eta_t = g(t_2 - \beta t_1) \quad (7)$$

In this specification, technology is represented by two terms. The first is the time trend component that captures the growth rate of the technology frontier assumed constant across individuals. The second term,  $\mu_i$ , a time-invariant component that varies across economies, should control for various unobservable factors like institutions or climate, and – crucially to our study – technology. Since technology is likely to be correlated with other regressors, a fixed effect estimator

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<sup>20</sup> In their study MRW's assume  $\ln A_{i0} = \ln A_0 + \varepsilon_i$ , with  $\ln A_0$  constant across individuals, and  $\varepsilon_i$  representing a random shock, uncorrelated with the other explanatory variables. Notice that, if instead technology is correlated with the explanatory variables, MRW's OLS results are not consistent.

is appropriate.<sup>21</sup>

Once we have the estimated individual intercepts, we can easily compute a proxy of TFP by:

$$A(0)_i = \exp\left(\frac{\mu_i}{1-\beta}\right) \quad (8)$$

In other words, this methodology can be used to obtain a measure of the degree of cross-country technology heterogeneity.<sup>22</sup>

The main problem with this methodology is that, while it was designed to control for the presence of cross-country TFP heterogeneity, it rules out technology convergence by assumption and completely ignores possible problems of observational equivalence between technological catch up and capital deepening. More precisely, as shown by equations (6) and (7), equation (2) is obtained by log-linearizing the Solow model around the steady-state under the assumption of a *constant* degree of TFP heterogeneity. In other words, all economies are assumed to grow at the same technological rate according to the process  $\ln A_{it} = \ln A_{i0} + gt$ , whatever their level of technological knowledge.

This is clearly in sharp contrast with the technological catching-up hypothesis. The latter may be described by a process where the growth rate of technology is proportional to the current gap between the world technology frontier and the technology level currently adopted in an economy. Typically, during the transition towards the steady-state in which all economies share the common long-run technological growth  $g$ , the presence of technological catch-up enables the technology levels in the lagging economies to grow faster than  $g$ . As a consequence, during the transition, the technology gap between the leader and a follower should decrease. On the contrary, if

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<sup>21</sup> For more on this see Baltagi (2003).

<sup>22</sup> One of the main criticisms of this approach is that the estimated individual intercepts do not simply control for technology but include also the effect of other possible unobservable factors such as institutions or geography. As explained in the following section, one way to control for this problem is to apply this methodology to samples that are relatively homogeneous with respect to other factors such as institutions.

no systematic process of technology diffusion is at work, this gap should stay constant over time, since all economies grow at a common rate of technology growth. Note that, since productivity is certainly correlated with technology, equation (2) is plagued by problems of observational equivalence, as the  $\beta$  coefficient may still capture the effect of both neoclassical convergence and catch-up.

Hence, how can we use equation (2) to test for the presence/absence of technological convergence? First, notice that, the longer the time dimension of the panel, the higher the risk that differences in TFP levels are not constant within the sample period, since technological diffusion is more likely to be at work. As a consequence, in the presence of technological convergence equation (2) should be regarded as an approximation of the real process – an approximation that deteriorates as the length of the period under analysis increases.

Second, and consequently, the presence of technological convergence should be detected by comparing the TFP values obtained by estimating (8) over different periods. This type of comparison should reveal whether the observed pattern of TFP values is consistent either with the catching-up hypothesis or with the alternative hypothesis that the current degree of technology heterogeneity is at its stationary value (or constant). Indeed, in this case we are estimating the initial level of TFP at two different points in time. By doing this, we are able both to detect if the pattern over time of these TFP levels is consistent with the presence/absence of catch up and to control for observational equivalence problems.

The technique we use follows several, simple steps. First, we estimate equation (2) over different sub-periods, in order to obtain a sequence of estimated values of individual intercepts. Second, the latter values are used to compute the individual values of  $\ln A_i$ . Third, we analyse the evolution over time of the distribution of  $\ln A_i$  in order to test for the presence of technological convergence. Indeed, while technology convergence implies a variance of  $\ln A_i$  that decreases over time while approaching its stationary value, the alternative hypothesis implies that the variance of

In  $A_i$  is at its stationary value, and thus no significant trend in its value should be detected. A similar approach has been previously introduced by Islam (2000), where the author compares the distribution of the estimated fixed effects over two points in time. However, in his paper the possibility that technology convergence lies behind the observed changes in the distribution is neither discussed nor tested.

#### 4. Comparing the available estimation procedures

We use Italian data on regional GDP per worker 1963-1993<sup>23</sup> to estimate the following equation:

$$\tilde{y}_{it} = \beta \tilde{y}_{it-\tau} + \sum_{j=1}^3 \gamma_j \tilde{x}_{jit-\tau} + \mu_i + u_{it} \quad (9)$$

$$\tilde{y}_{it} = y_{it} - \bar{y}_t, \quad \tilde{x}_{it} = x_{it} - \bar{x}_t \quad (10)$$

where  $\bar{y}_t$ , and  $\bar{x}_t$  are the Italian average in period  $t$ : data are taken in difference from the Italian mean, in order to control for the presence of a time trend component  $\eta_t$  and of a likely common stochastic trend (the common component of technology) across regions. We use a suitable time span in order to control for business cycle fluctuations and serial correlation, which are likely to affect the data in the short run. Moreover,  $x_{1i,t}$  is the lagged saving rate proxied by the ratio of regional investment to GDP, and  $x_{3it}$  represents a measure of human capital stock, namely average years of schooling. Both these variables are taken at their  $t - \tau$  level, while  $x_{2i,t}$  represents the sum of  $n$ , population growth,  $\delta$  the depreciation rate and  $g$  the exogenous technology growth rate and is taken as an average over the five years preceding  $t$ .<sup>24</sup> As standard in this literature,  $(g + \delta)$  is assumed equal to 0.05. Note that equation (9) simply augments equation (2) to include a measure of the stock of human capital since, in order to identify TFP differences, it is essential to control for one of its

<sup>23</sup> The length of the dataset is constrained by the human capital variable. We use census data and more recent observations are not available yet.

<sup>24</sup> In fact, regional series may be characterised by high volatility and this was the case for  $x_{i,t}^2$ .

most likely determinants.<sup>25</sup> Finally, as shown in the previous section, the coefficient on the lagged dependent variable yields a measure of the speed of the conditional convergence (or *within* convergence), while individual effects reflect the degree of TFP heterogeneity.

The first problem we face when we estimate a dynamic panel data model such as the one represented by equation (9) is which estimator suits our case better. The answer is not simple. It is still true today what Kiviet wrote a few years ago: “As yet, no technique is available that has shown uniform superiority in finite samples over a wide range of relevant situations as far as the true parameter values and the further properties of the DGP are concerned.”<sup>26</sup> Indeed, the LSDV estimator, while consistent for large  $T$ ,<sup>27</sup> is characterised by small sample problems and, in particular, it is known to produce downward biased estimates in small samples. Conversely, the Arellano and Bond (1991) estimator (GMM-AB from now on) is becoming increasingly popular since it has both the advantage of producing consistent estimates in a dynamic panel regression with both (i) endogenous right hand side variables, and (ii) the presence of measurement error. Moreover, it is more efficient than other standard IV estimators such as the Anderson-Hsiao estimator. However, it has recently been shown<sup>28</sup> that, when  $T$  is small, and either the autoregressive parameter is close to one (highly persistent series), or the variance of the individual effect is high relative to the variance of the transient shock, then even the GMM-AB estimator is biased and, in particular, downward biased. Note that the presence of a relatively small number of time periods and persistent time series are typical features of macro-growth datasets like ours.

To control for this problem, Kiviet (1995) put forward a more direct approach to the problem of the LSDV finite sample bias by estimating a small sample correction to the LSDV estimator. Monte Carlo analysis<sup>29</sup> finds that for small  $T$  (such as the one we usually find in

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<sup>25</sup> For details on how this variable is constructed see Di Liberto (2001).

<sup>26</sup> See Kiviet (1995) page 72.

<sup>27</sup> See Amemyia (1967).

<sup>28</sup> See Blundell and Bond (1998) and Bond-Hoeffler-Temple (2001). We do not consider the system-GMM suggested in these studies here since this methodology requires that first-difference  $\Delta y_{it}$  are not correlated with  $\mu_i$ , thus implying the absence of technological catching-up.

<sup>29</sup> See Kiviet (1995) and Judson and Owen (1996).

convergence literature) LSDV estimates corrected for the bias (KIViet from now on) seem more attractive than GMM. In particular, these Monte Carlo studies explicitly analyse typical macro dynamic panels and find that for  $T \leq 20$  and  $N \leq 50$ , as in our case, the KIViet and Anderson-Hsiao estimators consistently outperform GMM-AB. Moreover, despite having a higher average bias, KIViet turns out to be more efficient than Anderson-Hsiao. Overall, these Monte Carlo analyses suggest that a reasonable strategy would be to use the KIViet estimator for smaller panels ( $T \leq 10$ ), while Anderson-Hsiao should be preferred for larger panels, as the efficiency of the latter improves with  $T$  (Anderson-Hsiao has the additional advantage of being computationally simpler than the former).

Let us now turn to our specific case. Our Italian regional panel includes the period 1963-93 for 19 regions.<sup>30</sup> Using the five-year time span (or  $\tau = 5$ ) implies that we are left with  $T=7$  observations for each of the  $N=19$  regions, corresponding to 1963, 1968, 1973, 1978, 1983, 1988, and 1993. Given the dimension of our panel and the above discussion, the Kiviet-corrected LSDV estimator should be preferred<sup>31</sup>. However, since there is still a lot of debate on which technique shows a clear superiority in finite samples, in the following analysis we will use several estimators and will compare their results in order to assess their robustness.

## 5. Testing for TFP heterogeneity

Before moving into the estimate of the presence of TFP convergence, we firstly test for the presence of regional TFP heterogeneity using the whole sample 1963-93. We suggest using this strategy because the presence of TFP heterogeneity may be considered as a necessary but not sufficient condition for TFP convergence. Moreover, note that the analysis of TFP convergence requires a further reduction in our sample and, as we have seen above, all these estimators perform poorly in small samples. Thus, we only continue our analysis further if we find evidence in favour

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<sup>30</sup> There are 20 Italian regions. We have excluded Valle d'Aosta because it represents an outlier. Nevertheless, results do not change if we include this region.

<sup>31</sup> To implement the Kiviet's small-sample correction we use the STATA routine proposed by Adam (1998).

of TFP heterogeneity.

In Table 1 we compare the results obtained by using four different estimators, namely, a standard pooling OLS, LSDV, KIVIET and GMM-AB. For each regression we include the estimates obtained and the implied  $\hat{\lambda}$ , i.e. the speed of the convergence parameter.<sup>32</sup>

### TABLE 1

Let us start by the results obtained using the (uncorrected) LSDV estimator. When equation (9) is estimated with LSDV (Model 2) we find an AR(1) coefficient of 0.51 and a correspondingly relatively high speed of convergence of 13%. Among the regressors, only the coefficients on the lagged dependent variable and on population growth are significant, while both the coefficient on the investment share and on human capital are not significant. These results will be confirmed when other estimation procedures are introduced. The use of the Kiviet correction procedure increases the LSDV parameter from 0.51 to 0.67 (model 3), with a decline in the corresponding speed of convergence coefficient from 13% to 8%<sup>33</sup>. Note that KIVIET only corrects the bias in the LSDV estimated parameters but does not produce alternative or corrected standard errors. This is why they are not shown among the results.

Let us now extend our comparison to the other available estimators. The GMM-AB estimates are shown together with the p-value of the AB-2 statistic and the Sargan test as in Arellano and Bond (1991)<sup>34</sup>. The GMM-AB estimator may be performed under very different assumptions on the endogeneity of included regressors. In this study we specify two different

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<sup>32</sup> From  $\beta = e^{-\lambda t}$

<sup>33</sup> To compute the estimated bias the methodology requires the use of a consistent estimator, such as Anderson-Hsiao or GMM, and the routine used to calculate KIVIET cannot handle missing observations. See Adam (1998).

<sup>34</sup> The first statistic tests for the presence of serial correlation. In particular, since the final regression equation is in first differences, it tests for second order serial correlation in the error term. The latter must be absent for the assumption of no serial correlation in the model in levels to be accepted. The presence of second-order serial correlation would imply that the estimates are inconsistent. The second statistic tests the validity of overidentifying restrictions. The consistency of this estimation procedure crucially depends on the identifying assumption that lagged values of both income and other explanatory variables are valid instruments in these growth regressions.



hypotheses on the additional regressors  $x$ 's. First, Model 4 (or model GMM-AB1) in Table 1 assumes that all regressors are predetermined, that is  $E[x_{it}u_{is}] \neq 0$  for  $s < t$ , and zero otherwise. Second, Model 5 (or model GMM-AB2) assumes instead that all  $x$ 's are strictly exogenous, that is,  $E[x_{it}u_{is}] = 0$  for all  $t$  and  $s$ <sup>35</sup>.

The results in Table 1 show that both specifications are valid: the p-values of the AB-2 and Sargan tests say that it is not possible to reject the null hypothesis of absence of second-order autocorrelation and that the over-identifying restrictions are valid. Thus, the choice between these two specifications is not obvious, even if the increase of the p-value of the Sargan test in GMM-AB1 indicates that treating the included regressors as predetermined makes it more difficult to reject the null and, thus, that Model 4 should be preferred to Model 5. However, these estimates may be biased. To detect a possible bias in our GMM-AB models we follow the procedure suggested by Bond, Hoeffler and Temple (2001). Given that it is well known that OLS is biased upwards in dynamic panels and LSDV is biased downwards, these authors suggest that a consistent estimate should therefore lie between the two. As expected, when we introduce the pooling-OLS estimator (Model 1), the coefficient on the lagged dependent variable is high compared with other individual effects estimators, with a corresponding speed of convergence of 4%. All regressors are in this case significant and with the expected sign. Since we presume that the true parameter values lie somewhere between  $\hat{\beta}_{ols}$  and  $\hat{\beta}_{LSDV}$ , in this specific case we expect its value to be between 0.80 and 0.51. The estimated AR(1) coefficient on GMM-AB2 is higher than that obtained with OLS, where the latter should be characterized by upward bias. Consequently, we exclude GMM-AB2 from the following analysis, while we accept GMM-AB1 since its estimated AR(1) coefficient is 0.696, very similar to that obtained in KIVIET, 0.669. With a value ranging between  $\hat{\beta}_{ols}$  and  $\hat{\beta}_{LSDV}$ , this estimate does not suggest any obvious presence of bias.

With these estimates in hand we now move on to compute our regional TFP measures. In

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<sup>35</sup> See Baltagi (2003).

our LSDV estimates the regional dummy coefficients,  $\hat{\mu}_i$ , are almost invariably statistically significant. As for the GMM-AB1 estimator, note that it controls for fixed effects by transforming data in first difference, so the individual effects are therefore not directly estimated. Following Caselli, Esquivel and Lefort (1996), we obtain estimates of  $\mu_i$  and, through equation (8),  $\hat{A}(0)_i$ , by:

$$(\hat{\mu}_i + \hat{u}_{it}) = \tilde{y}_{it} - \hat{b}\tilde{y}_{it-1} - \sum_{j=1}^2 \hat{\gamma}_j \tilde{x}_{jit-1} - \hat{\xi}\tilde{h}_{it-1} \quad (11)$$

$$\hat{\mu}_i = \frac{1}{T} \sum (\hat{\mu}_i + \hat{u}_{it}) \quad (12)$$

The same procedure has been used to obtain  $\hat{\mu}_i$  using KIVLET. In all cases, the regional TFP estimates  $\hat{A}(0)_i$  are then used to compute the ratio  $\hat{A}(0)_i / \hat{A}(0)_{Lom}$ , with  $\hat{A}(0)_{Lom}$  being the estimated TFP value for Lombardia, currently the richest, most industrialised and arguably the most technologically advanced Italian region.

Table 2a includes the estimates of the relative (to Lombardia) levels of regional TFP obtained by applying each different estimator. To make the interpretation of results easier, Table 2b shows the ranking of each region obtained with the different methodologies.

### TABLE 2A & TABLE 2B

Overall, these results suggest that different econometric methodologies produce similar TFP estimates. This conclusion is confirmed by the analysis of the Spearman rank order coefficient (see Table 3).

### TABLE 3.

In particular, as expected, the estimates of the regional relative TFP levels obtained by LSDV and

KIVIET are almost identical, with a correlation coefficient equal to 0.98. A lower but still very high ranking correspondence is found between the LSDV, KIVIET and GMM-AB1 estimates. Thus, the close correspondence found in this section among the TFP estimates obtained with different estimation procedures support the conclusion that our results can be safely regarded as robust.

Moreover, the overall picture emerging from our estimates is an interesting one. In particular, it strongly confirms that TFP differences can be significantly large even across regions, and that an important part of the Italian economic divide seems to be due to such differences. We also find confirmation that the northern and richer regions are the most technologically advanced areas in the country, and that at the bottom end are the southern, less developed areas.

To sum up, the pattern and the magnitude of TFP heterogeneity as measured by our estimates suggest that a potential for technological catch-up does exist for the lagging regions. In turn, this implies that any analysis of aggregate convergence across Italian regions should take this potential source of convergence into account. This is what we will do in the next section.

## **6. Detecting technological convergence: Empirical results**

To investigate whether the observed degree of hereogeneity is either constant or is the source of a process of TFP convergence we follow Islam (2003). He suggests testing for the presence of TFP convergence generating TFP-level indices for several consecutive time periods, so that the TFP dynamics can be directly analysed. The indices produced by panel methods may be used to this end as "...they contain ordinal as well as cardinal information, which can both be helpful in answering questions regarding TFP-convergence".<sup>36</sup> The main difficulty with this procedure is that, in order to generate different TFP-level indices for consecutive time periods we need to further reduce the time dimension of the estimated samples, thus worsening the problems associated with small sample bias discussed above. The KIVIET estimator should certainly be preferred with very small samples. However, note that using a time span equal to 5 implies that we are left with  $T=7$  observations,

which in turn implies  $T=5$  with Kiviet since it uses the Anderson-Hsiao estimates to calculate the bias.<sup>37</sup>

With such a dataset, the implementation of the KIVJET correction procedure to these short sub-samples becomes unfeasible. A possible alternative is to gain degrees of freedom by using a shorter time span. For example, a time span equal to three ( $\tau = 3$ ) yields  $T=11$  (i.e.,  $T=10$  with LSDV and  $T=9$  with KIVJET). Clearly, using a shorter time span has the obvious disadvantage that it increases the problems related to short term disturbances and serial correlation of the error term.<sup>38</sup> Given these problems, we firstly estimate TFP levels KIVJET to a sample with  $\tau = 3$ . Moreover, to control if the results obtained are robust, we apply the same procedure using LSDV since this estimator may be applied with a standard time span  $\tau = 5$ <sup>39</sup>.

When we apply KIVJET to the data with the non-standard time span of three years we obtain two sub-samples, each including 5 observations with one overlapping year. The first sub-sample includes 1969, 1972, 1975, 1978, 1981, the second 1981, 1984, 1987, 1990, 1993. As before, we estimate equation (9) and save the two different series of  $\hat{\mu}_i$ . Results are shown in Table 4.

#### TABLE 4

Models 1 and 2 in Table 4 suggest that the convergence coefficient is significant in both sub-periods. In the first subsample analysed, the other variables included are never significant, while in the second both  $\ln(s)$  and  $\ln(n + \delta + g)$  are negative and significant. Human capital is never significant. Moreover,  $\hat{\mu}_i$ , are almost invariably significant. Again, we use equation (6) to obtain

$\hat{A}(0)_i$ , and transform the data as  $\hat{A}(0)_i / \hat{A}(0)_{Lom}$ , with  $\hat{A}(0)_{Lom}$  being the estimated fixed effect of

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<sup>36</sup> See Islam (2003), p.349.

<sup>37</sup> Where data are taken in first difference and levels of the dependent variable (lagged twice and further) are introduced as instruments. For more details see Baltagi (2003).

<sup>38</sup> In this case we need to assume that measurement error is not three-order serially correlated. Moreover, it has been argued that short time spans may not be appropriate for studying growth convergence. See Islam (1995) p. 1140 and Caselli Esquivel and Lefort (1996) among others.

Lombardia. This procedure yields two different estimates of regional effects, the first corresponding mainly to the 1970s and the second mainly to the 1980s. Table 5 shows the relative TFP values for the two sub-periods. A well-defined dynamic pattern emerges as relative TFP values increase for most southern laggard regions in the second period, while they decrease for most northern regions. Moreover, regional TFP dispersion decreases: the variance of (relative) TFPs is higher in the first period (with a value of 0.020) than in the second (0.011).

These results are strongly consistent with the hypothesis that a process of *technological convergence* does exist and represents an important component of the aggregate convergence observed across these two sub-periods<sup>40</sup>. This conclusion finds clear confirmation in Figure 1, which shows the relationship existing between the TFP estimated for the first sub-sample and the subsequent one.

### FIGURE 1

The dotted 45 degree line shows the locus where the relative TFP level in each region would be unchanged between the two periods. Most southern regions are clearly above the 45 degree line<sup>41</sup> as they performed consistently better in terms of relative TFP growth, while six (mostly northern) regions are below the 45 degree line. This pattern is consistent with the hypothesis of TFP convergence.

The same data may be rearranged to analyse this result in terms of a typical *growth-initial level* convergence relationship. In Figure 2 the Y-axis represents the rate of growth of relative TFP, while the X-axis represents the initial relative TFP level. Convergence implies a negative correlation between the initial level of TFP and its subsequent growth rate. This is exactly what our

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<sup>39</sup> We do not lose one observation due to a first-difference transformation as with KIVIET.

<sup>40</sup> These results are qualitatively in line with those obtained by Maffezzoli (2004) for the 1980-2000 period using the Data Envelope Analysis approach

data reveal.

## FIGURE 2

Note that, since our TFP measures are interpreted as initial period TFP levels, our evidence shows that technology convergence took place *between* the two sub-periods of our analysis (1963 and 1981), while nothing can be inferred on what has happened, in terms of technology diffusion, *within* the second sub-period. Thus, during the period of observed per capita GDP convergence we find evidence that both neoclassical and TFP convergence forces were at work while nothing can be inferred for the second period. Indeed, as shown by Table 5, in spite of this clear convergence pattern during the first sub-period, the distance between northern and southern regions in terms of relative TFP remains significant in the second sub-period. So, one possibility is that the halt of aggregate convergence observed in this sub-period is due to a halt of technology diffusion.

Let us now go back to Table 4, columns 3 and 4, in order to assess the robustness of our previous result by applying the LSDV estimator using the standard time-span. Our LSDV estimates in columns 1 and 2 are based on two sub-samples of 4 observations each, with one overlapping year: the first sub-sample includes 1963, 1968, 1973 1978, and the second 1978, 1983, 1988, 1993. In this case, the convergence coefficient is significant only in the first subsample, 1963-78, while we see no evidence of convergence during 1978-93. Apart from this, our previous results based on the LSDV estimator are confirmed. Other explanatory variables are never significant while the regional dummies coefficients,  $\hat{\mu}_i$ , are almost invariably significant. Again, regional TFP dispersion decreases since the variance of (relative) TFPs is higher in the first period (with a value of 0.027)

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<sup>41</sup> These are Molise, Basilicata, Calabria, Puglia, Sicilia and Campania.

than in the second  $(0.010)^{42}$ . In sum, our previous result pointing to the presence of a clear pattern of TFP convergence appears to be robust to the use of LSDV with a different time-span.

## **7. Technology convergence and the role of human capital**

Finally, our measures of regional levels of TFP may also be used to test one of the main hypotheses of literature on catching-up. Nelson and Phelps (1966) and Benhabib and Spiegel (1994) show how TFP growth can be determined by the technological distance from the leader and by the level of human capital, where the latter influences the capacity for both discovering new technologies and adopting innovations from abroad. In other words, human capital levels determine the capacity of adopting new technologies from abroad and, thus, the possibility of a catch up process among countries. As a consequence, we should expect TFP levels to depend on human capital stocks, while the role of human capital on GDP growth is only indirect. This is exactly what our results seem to suggest. The correlation coefficient between the regional human capital level in 1969 and TFP levels estimated using KIVIET ( $\tau = 5$ ) is equal to 0.93, while that calculated between the level of human capital in 1978 and the subsequent TFP levels is 0.89<sup>43</sup>. The same relationship may be observed in Figures 3 and 4, where we include regional relative TFP estimates in the Y-axis and the initial level of human capital in the X-axis. In general, this evidence, together with the absence of significance of the human capital variable in our regressions, corroborate the hypothesis of a relationship between human capital and TFP as described by the Nelson and Phelps approach.

## **FIGURES 3 AND 4**

## **8. Conclusion**

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<sup>42</sup> Using these new TFP estimates we would obtain the same pattern shown by Figures 1 and 2. The only difference is observed for a group of three regions, Abruzzo, Umbria and Sardegna: these regions are among the (relative) losers in Figure 1 while in this case they would be among the (relative) winners.

<sup>43</sup> When we use LSDV, the results are almost identical (0.93, 0.87).

The aim of this paper was to assess the existence of technology convergence across the Italian regions between 1963 and 1993. Different methodologies have been proposed to measure TFP heterogeneity across countries, but only a few of them try to capture the presence of technology convergence as a separate component from the standard (capital-deepening) source of convergence. To distinguish between these two components of convergence, we have proposed and applied a fixed-effect panel methodology.

Our results show, first of all, the presence of TFP heterogeneity across Italian regions. This result is robust to the use of different estimation procedures such as simple LSDV, Kiviet-corrected LSDV, and GMM *à la* Arellano and Bond (1991). Second, we find strong evidence to support the hypothesis that a significant process of TFP convergence has been a key factor in the observed aggregate regional convergence that took place in Italy up to the mid-seventies.

Moreover, our results show that a period of significant convergence in TFP has not generated a significant, persistent decrease in the degree of cross-region inequality in per capita income. The solution to this puzzle may be a simple one. Our evidence shows that technology convergence took place *between* the two sub-periods of our analysis (1963-78 and 1978-93), while nothing can be inferred on what has happened, in terms of technology diffusion, *within* the second sub-period. So, one possibility is that the halt of aggregate convergence in this sub-period is due to a halt of technology diffusion. More data and research are needed to test this additional hypothesis.

Finally, our human capital measures have been found to be highly positively correlated with TFP levels. This result confirms one of the hypothesis of the Nelson and Phelps approach, namely that human capital is the main determinant of technological catch-up. This latter result suggests an explanation for the existence of persistent differences in regional GDP per worker which might be due to the fact that the backward regions never caught-up with the the northern and richest ones in terms of their human capital endowments.



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## APPENDIX

**Table 1: Estimation of the augmented Solow model**

<b>Sample: Italian regions (1963-93)</b>					
<b>Dependent Variable</b>					
	<b>1</b>	<b>2</b>	<b>3</b>	<b>4</b>	<b>5</b>
	<b>OLS</b>	<b>LSDV</b>	<b>KIVIET</b>	<b>GMM-AB1</b>	<b>GMM-AB2</b>
<i>Observations</i>	95	95	95	95	95
$\ln(y_{i,t-1})$	.800 (.051)	.514 (.090)	.669	.696 (.097)	.834 (.135)
$\ln(s_{i,t-1})$	-.063 (.020)	-.019 (.030)	-.022	-.047 (.043)	-.044 (.037)
$\ln(n_{i,t-1} + g + d)$	-.108 (.025)	-.074 (.028)	-.089	-.125 (.031)	-.108 (.028)
<i>human capital</i>	.024 (.011)	-.010 (.019)	0.00	.001 (.021)	-.028 (.022)
$\lambda$	0.04	0.13	0.08	0.07	0.04
<i>Sargan test (p-value)</i>				0.79	0.35
<i>AB-2 test (p-value)</i>				0.31	0.20

**Notes:**

19 Italian regions included (Val d'Aosta excluded)

Standard errors in parentheses.

LSDV is the Least Squares with Dummy variables estimator.

GMM-AB1 is the Arellano-Bond estimator under the assumption that  $x$ 's are predetermined

GMM-AB2 is the Arellano-Bond estimator under the assumption  $x$ 's strictly exogenous

KIVIET is the LSDV estimator with the Kiviet (1995) correction. KIVIET only corrects the bias in the LSDV estimated parameters, but does not produce alternative or corrected standard errors.

This is why they are not shown among results.

The figures reported for the Sargan test are the  $p$ -values for the null hypothesis, valid specification.

The figures reported for the AB-2 test are the  $p$ -values of the Arellano-Bond test that average autocovariance in residuals of order 2 is 0.

Lambda is the corresponding convergence coefficient

**Table 2a: Regional relative TFP levels by different estimation procedures**

<i>ln(y) 1968</i>		LSDV		GMM-AB1		KIVIET	
<i>Lazio</i>	3.69	<i>Lombardia</i>	1.00	<i>Lombardia</i>	1.00	<i>Lombardia</i>	1.00
<i>Liguria</i>	3.67	<i>Lazio</i>	0.97	<i>Lazio</i>	0.99	<i>Lazio</i>	0.98
<i>Lombardia</i>	3.61	<i>Liguria</i>	0.95	<i>Friuli</i>	0.94	<i>Emilia Romagna</i>	0.95
<i>Toscana</i>	3.55	<i>Emilia Romagna</i>	0.94	<i>Emilia Romagna</i>	0.94	<i>Friuli</i>	0.95
<i>Trentino</i>	3.54	<i>Friuli</i>	0.92	<i>Veneto</i>	0.94	<i>Liguria</i>	0.95
<i>Emilia Romagna</i>	3.51	<i>Veneto</i>	0.90	<i>Liguria</i>	0.93	<i>Veneto</i>	0.92
<i>Veneto</i>	3.49	<i>Piemonte</i>	0.89	<i>Trentino</i>	0.89	<i>Piemonte</i>	0.90
<i>Piemonte</i>	3.47	<i>Toscana</i>	0.86	<i>Sardegna</i>	0.89	<i>Trentino</i>	0.88
<i>Sardegna</i>	3.45	<i>Trentino</i>	0.86	<i>Piemonte</i>	0.87	<i>Toscana</i>	0.86
<i>Friuli</i>	3.40	<i>Marche</i>	0.82	<i>Toscana</i>	0.86	<i>Abruzzo</i>	0.85
<i>Campania</i>	3.39	<i>Sardegna</i>	0.81	<i>Abruzzo</i>	0.85	<i>Marche</i>	0.85
<i>Sicilia</i>	3.38	<i>Abruzzo</i>	0.81	<i>Marche</i>	0.85	<i>Umbria</i>	0.84
<i>Marche</i>	3.36	<i>Umbria</i>	0.81	<i>Umbria</i>	0.84	<i>Sardegna</i>	0.83
<i>Umbria</i>	3.35	<i>Sicilia</i>	0.76	<i>Sicilia</i>	0.81	<i>Puglia</i>	0.78
<i>Abruzzo</i>	3.24	<i>Puglia</i>	0.75	<i>Puglia</i>	0.79	<i>Sicilia</i>	0.78
<i>Puglia</i>	3.23	<i>Campania</i>	0.73	<i>Campania</i>	0.77	<i>Campania</i>	0.75
<i>Calabria</i>	3.21	<i>Molise</i>	0.68	<i>Basilicata</i>	0.72	<i>Molise</i>	0.73
<i>Basilicata</i>	3.17	<i>Calabria</i>	0.65	<i>Molise</i>	0.70	<i>Calabria</i>	0.68
<i>Molise</i>	3.06	<i>Basilicata</i>	0.64	<i>Calabria</i>	0.68	<i>Basilicata</i>	0.67

**Notes:**

*ln(y) 1968* is the logarithm of GDP per worker in 1968

LSDV includes the regional individual effects estimated using the LSDV estimator (95 observations)

GMM-AB1 includes the regional individual effects calculated using the GMM-AB1 estimator

KIVIET includes the regional individual effects calculated using the LSDV estimator with the KIVIET correction

**Table 2b: Regional TFP ranking**

<b>Regions</b>	<b>LSDV</b>	<b>GMM-AB1</b>	<b>KIVIET</b>
<i>Abruzzo</i>	12	11	10
<i>Basilicata</i>	19	17	19
<i>Calabria</i>	18	19	18
<i>Campania</i>	16	16	16
<i>Emilia Romagna</i>	4	4	3
<i>Friuli</i>	5	3	4
<i>Lazio</i>	2	2	2
<i>Liguria</i>	3	6	5
<i>Lombardia</i>	1	1	1
<i>Marche</i>	10	12	11
<i>Molise</i>	17	18	17
<i>Piemonte</i>	7	9	7
<i>Puglia</i>	15	15	14
<i>Sardegna</i>	11	8	13
<i>Sicilia</i>	14	14	15
<i>Toscana</i>	8	10	9
<i>Trentino</i>	9	7	8
<i>Umbria</i>	13	13	12
<i>Veneto</i>	6	5	6

**Notes:**

*LSDV* includes the rank of regional individual effects estimated using the *LSDV* estimator.

*GMM-AB1* includes the rank of regional individual effects calculated using the *GMM-AB1* estimator.

*KIVIET* includes the rank of regional individual effects calculated using the *LSDV* estimator with the *KIVIET* correction.

<b>Table 3: Spearman rank order correlation coefficient</b>			
	LSDV	GMM-AB1	KIVIET
LSDV	1		
GMM-AB1	0.94	1	
KIVIET	0.98	0.96	1

**Table 4: Estimation of the augmented Solow model (two subsamples)**

<b>Sample: Italian regions 1963-978 and 1978-93</b>				
<b>Dependent Variable</b>	<b><math>\ln(y_{i,t})</math></b>			
	<b>1</b>	<b>2</b>	<b>3</b>	<b>4</b>
	<i>time span=3</i>	<i>time span=3</i>	<i>time span=5</i>	<i>time span=5</i>
	<b>1963-1981</b>	<b>1981-1993</b>	<b>1963-1978</b>	<b>1978-1993</b>
	<b>KIVIET</b>	<b>KIVIET</b>	<b>LSDV</b>	<b>LSDV</b>
<i>Observations</i>	95	95	57	57
$\ln(y_{i,t-1})$	.760	.386	.480 (.141)	.015 (.124)
$\ln(s_{i,t-1})$	-.068	-.061	.055 (.049)	-.051 (.036)
$\ln(n_{i,t-1} + g + d)$	-.047	-.062	-.061 (.052)	-.069 (.027)
<i>human capital</i>	-.30	.013	-.051 (.052)	.006 (.044)
$\lambda$	0.09	0.32	0.15	0.84

**Notes:**

19 Italian regions included (Val d'Aosta excluded)

Standard errors in parentheses.

LSDV is the Least Squares with Dummy variables estimator.

KIVIET is the LSDV estimator with the Kiviet (1995) correction and only corrects the bias in the estimated parameters.

It does not produce alternative or corrected standard errors. This is why they are not shown among results.

Lambda is the corresponding convergence coefficient

**Table 5: Estimated Relative regional TFP levels**

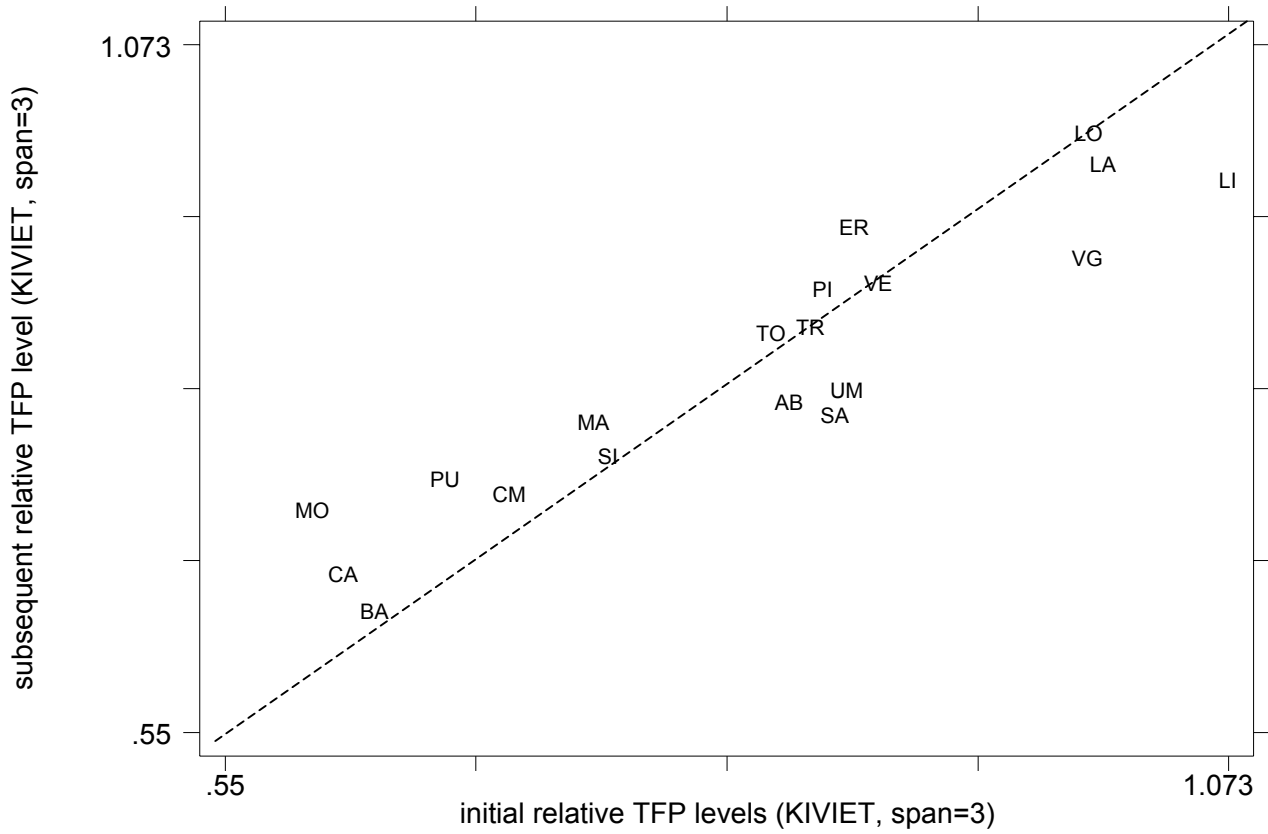
<b>Regions</b>	<b>Relative TFP levels: initial period</b>	<b>Relative TFP levels: subsequent period</b>
<i>Piemonte</i>	0.86	0.88
<i>Lombardia</i>	1.00	1.00
<i>Trentino Alto Adige</i>	0.85	0.85
<i>Veneto</i>	0.89	0.89
<i>Friuli Venezia Giulia</i>	1.00	0.91
<i>Liguria</i>	1.07	0.96
<i>Emilia Romagna</i>	0.88	0.93
<i>Toscana</i>	0.83	0.85
<i>Umbria</i>	0.87	0.80
<i>Marche</i>	0.74	0.78
<i>Lazio</i>	1.01	0.98
<i>Abruzzo</i>	0.84	0.80
<i>Molise</i>	0.60	0.71
<i>Campania</i>	0.70	0.72
<i>Puglia</i>	0.66	0.74
<i>Basilicata</i>	0.63	0.64
<i>Calabria</i>	0.61	0.66
<i>Sicilia</i>	0.75	0.75
<i>Sardegna</i>	0.87	0.79
Variance	0.020	0.011

**Notes:**

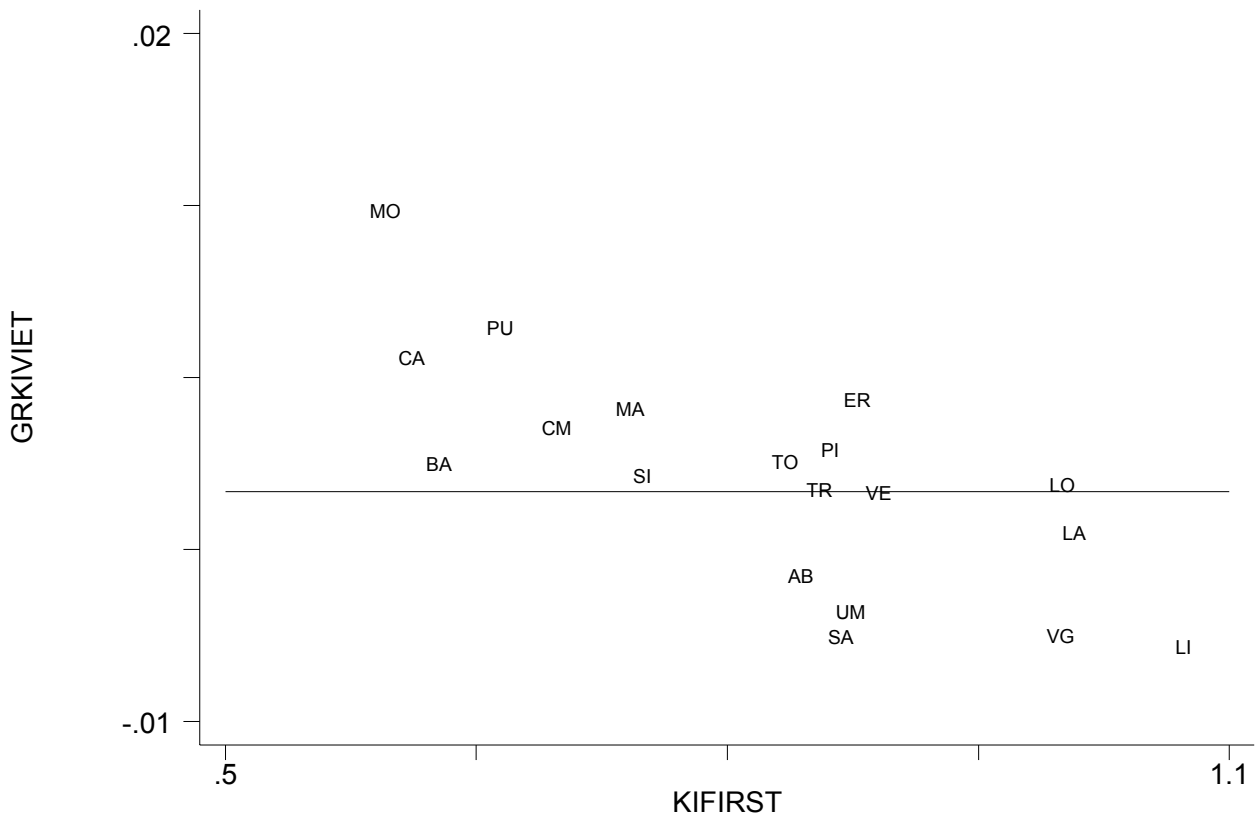
*Initial TFP level correspond to the TFP estimated using the sample 1969-1981;*

*Subsequent TFP level correspond to the TFP estimated using the sample 1981-1993.*

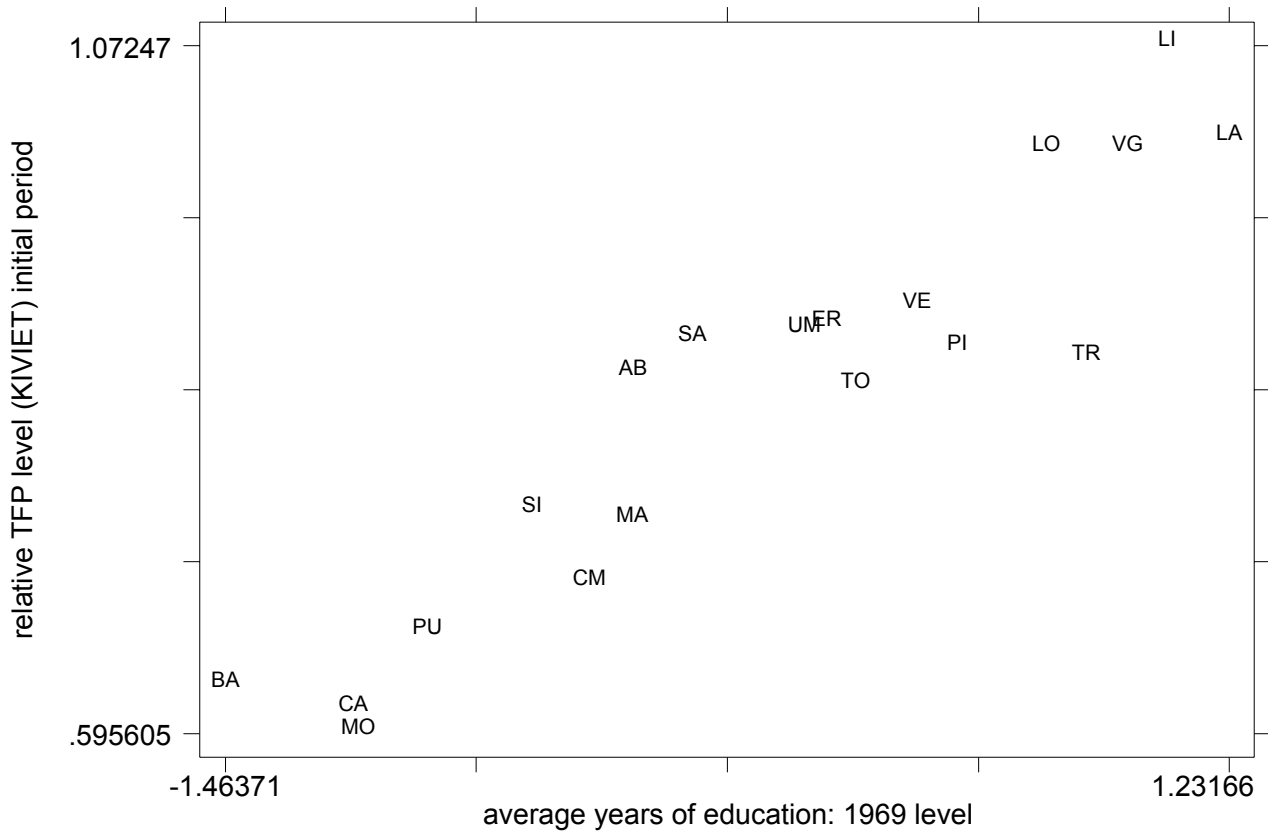




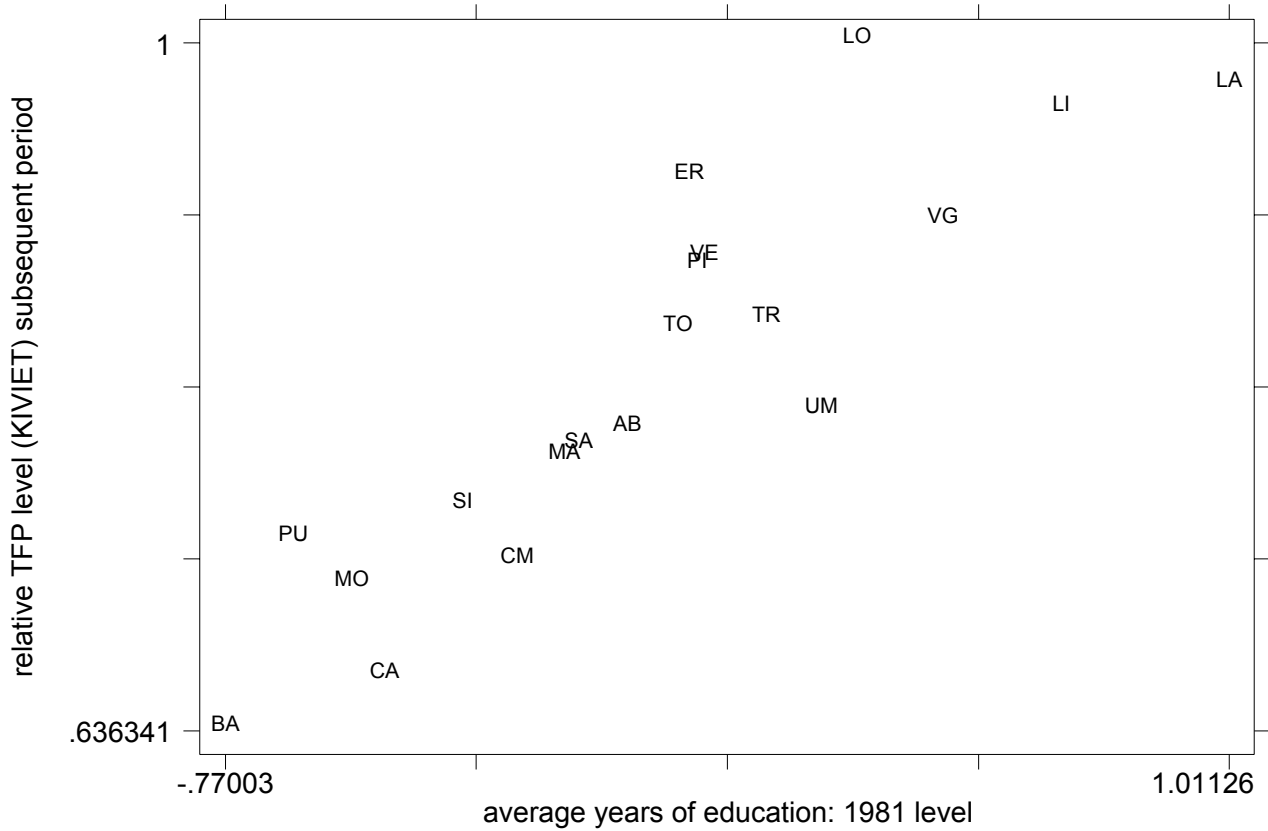
**Figure 1: Productivity dynamics (KIVIET)**



**Figure 2: Productivity dynamics (KIVIET), growth-initial level relationship**



**Figure 3:** human capital and TFP, initial period



**Figure 4:** human capital and TFP, subsequent period

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(lxv) This paper was presented at the EuroConference on “Auctions and Market Design: Theory, Evidence and Applications” organised by Fondazione Eni Enrico Mattei and sponsored by the EU, Milan, September 25-27, 2003

(lxvi) This paper has been presented at the 4th BioEcon Workshop on “Economic Analysis of Policies for Biodiversity Conservation” organised on behalf of the BIOECON Network by Fondazione Eni Enrico Mattei, Venice International University (VIU) and University College London (UCL), Venice, August 28-29, 2003

(lxvii) This paper has been presented at the international conference on “Tourism and Sustainable Economic Development – Macro and Micro Economic Issues” jointly organised by CRENoS (Università di Cagliari e Sassari, Italy) and Fondazione Eni Enrico Mattei, and supported by the World Bank, Sardinia, September 19-20, 2003

(lxviii) This paper was presented at the ENGIME Workshop on “Governance and Policies in Multicultural Cities”, Rome, June 5-6, 2003

(lxix) This paper was presented at the Fourth EEP Plenary Workshop and EEP Conference “The Future of Climate Policy”, Cagliari, Italy, 27-28 March 2003

(lxx) This paper was presented at the 9<sup>th</sup> Coalition Theory Workshop on "Collective Decisions and Institutional Design" organised by the Universitat Autònoma de Barcelona and held in Barcelona, Spain, January 30-31, 2004

(lxxi) This paper was presented at the EuroConference on “Auctions and Market Design: Theory, Evidence and Applications”, organised by Fondazione Eni Enrico Mattei and Consip and sponsored by the EU, Rome, September 23-25, 2004



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