Taking seriously parameter heterogeneity: culture, market rules, returns to human capital and economic development

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Abstract

The paper analyzes the contribution of human capital and of the labour augmenting component to the observed levels of per capita GDP using a random coefficient finite mixture model. Our approach deals parsimoniously with three important problems of the empirical literature (parameter heterogeneity, omitted variable bias and departures from the normality assumption) and provides additional insights for the interpretation of the determinants of economic development. More specifically, we identify five clusters of countries in which the heterogeneous impact of human capital and of the labour augmenting component on per capita GDP depends on differences in latent variables (i.e. cultural and institutional factors, quality of the educational system). Our results seem to find support for theoretical hypotheses arguing that these latent variables are crucial to address talents to economically productive activities and to increase returns to schooling.

Keywords: human capital, economic development, semi-parametric mixture, parameter heterogeneity.

JEL:C14, 030, 040.

1 Introduction

Cross country regressions on the determinants of levels and growth of per capita GDP generally suffer from a relevant omitted variable bias since many variables affecting the dependent variable, the intercept and the magnitude of the impact of different regressors on the dependent variable are missing or non recordable. The widespread use of random or fixed effect panel estimates only partially solves this problem as it captures time invariant (in the fixed effect case), country specific hidden factors, but it is not capable of measuring how these factors affect magnitudes of the available regressor

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coefficients. Another main limit in the current growth literature is the restrictive assumption on the homogeneity of the effect of human capital on levels of per capita GDP. In the reality, the interplay of this factor with other country fundamentals (economic freedom, quality of institutions, quality of the educational system, culture, religion and social norms) makes it hard to believe that proxies of human capital may have the same effect on the dependent variable in markedly different country environments. In other words, the restriction, implicit in standard homogeneous coefficient estimates, which requires human capital investment to have the same effects on levels of per capita GDP in Sub-Saharan Africa and in OECD countries is clearly untenable. Hence, we should be extremely cautious in drawing policy conclusions for heterogeneous group of countries on the basis of a unique estimated coefficient (Temple, 2001a, 2001b). Durlauf (2001) emphasizes this point in his survey on the empirical growth literature by arguing that the development of proper methodologies tackling the heterogeneity issue is one of the main goals of the current research in this field.

With this respect, a promising path followed by the most recent literature is the development of threshold models (see, among others, Kalaitzidakis *et al.*, 2001; Liu and Stengos, 1999; Masanjala and Papageorgiou, 2004) in which the problem of parameters heterogeneity is solved through exogenous subsampling.

A further contribution in this direction may be the development of an approach which deals parsimoniously with the main issues at stake (omitted variable bias and parameter heterogeneity) and, at the same time, endogenizes the clustering process by identifying optimal subgroups of countries with homogeneous parameters. To justify the technical complexity required to implement such approach, the latter should demonstrate to provide additional insights with respect to both standard homogeneous parameter and threshold regression models.

To follow this path we devise a semi-parametric random coefficient (hereafter also RC) model which keeps into account the possibility of an unobserved heterogeneous impact of regressors in different macroareas. We use this approach to account for heterogeneity in the intercept and in the human capital effects on levels of per capita GDP in a Solow augmented specification, controlling for sources of variability arising from omitted environmental and socio-cultural factors.

Our non parametric maximum likelihood approach identifies clusters characterized by homogeneous values of the random components (Lindsay, 1983a, 1983b). The RC model, when compared to a panel fixed effect approach - since locations and corresponding probabilities are completely free to vary over the corresponding supports - has the advantage of restricting the individual effects to a small discrete set of possible values and of accommodating extreme and/or strongly asymmetric departures from the normality assumptions of most parametric estimators (Lindsay, 1983a, 1983b, Lindsay and Lesperance, 1995, MacLachlan and Peel, 2000).

To sum up, our paper faces two challenges: i) on the methodological point of view it aims to develop an approach which, as explained above, deals originally and parsimoniously with crucial issues in the empirical literature (omitted variable bias, parameters' heterogeneity and departures from normality assumption); ii) on the side of the empirical findings it aims to show that the extra complexity of the model is compensated by the advantage of providing new insights in the interpretation of the determinants of economic development.

Our approach is not entirely novel as it has been successfully applied by Paap *et al.*(2005) to the problem of the comparison of patterns of growth of Latinamerican, Asian and African countries. Our original contribution with respect to this paper is that we do not limit the approach to some world regions but try to extend it to all world countries and focus on the heterogeneity of the contribution of the human capital factor to economic development. The paper is divided into seven sections (including introduction and conclusions). The second section illustrates the random coefficient model, the third the EM algorithm used to estimate it, the fourth the choice of variables used for the empirical analysis, the fifth econometric findings and the sixth shows how the observed heterogeneity in the contribution of human capital and the labour augmenting component to per capita GDP levels is highly correlated with measures of quality of the educational system, institutional quality and country fundamentals. The seventh section concludes.

2 The Model

Consensus on the existence and relevance of parameter heterogeneity in cross-country level and growth regressions, observed under different estimating techniques, has significantly grown in the last decade (Bianchi, 1997; Bloom, *et al., 2003*; Brock and Durlauf, 2000; Desdoigts, 1999; Durlauf, 2001a; Durlauf *et al.*, 2001; Durlauf and Johnson, 1995; Kalaitzidakis *et al.*, 2001; Liu and Stengos, 1999; Masanjala and Papageorgiou, 2004; Paap and Van Dijk, 1998; Paap *et al.*, 2005 and Quah, 1996 and 1997). This evidence is in contrast with the standard homogeneity assumptions adopted when estimating the augmented Solow model. Furthermore, it is well known that empirical findings from the Solovian model can be affected by omitted variable and error-in-variable biases (Durlauf, 2000). The standard approach followed to solve these problems is to introduce additional factors generally incorporated into the labour augmenting component. In this way, though, parameter heterogeneity (and the role of latent variables on it) is not addressed and there is always the risk of not including all relevant explanatory variables. An alternative approach to solve jointly the omitted bias and the parameter heterogeneity problem is to handle random effects with a mixture model. In such model the random component captures the impact of the unobserved variables preventing the omitted variable bias, while the endogenous clustering of the mixture identifies subgroups of countries with homogeneous parameters, allowing us the possibility to explain parameter heterogeneity across subgroups by taking into account the role of measurable latent variables. To perform this task we conventionally define as $\{\mathbf{y}_1, \ldots, \mathbf{y}_n\}$ the realized random vectors of n conditional independent and identically distributed ($i.i.d$) levels of per capita GDP $\{\mathbf{Y}/\mathbf{L}_1, \ldots, \mathbf{Y}/\mathbf{L}_n\}$, where $(\mathbf{Y}/\mathbf{L})_i$ is a $T - dimensional$ random vector with probability density function $f(\mathbf{y}_i)$ on $\mathbf{R}^T$; i.e. $\mathbf{y}_i$ contains the realized random variables corresponding to the year t $(t = 1, ..., T)$ measurement made for the $i - th$ country in the sample. The $i - th$ country's output per worker in a Solow model augmented for the role of human capital (Mankiw, Romer and Weil, 1992) is typically

$$f(y_{it}|A_{it}, k_{it}, h_{it}) = A_{it}^{(1-\alpha-\beta)} k_{it}^{\alpha} h_{it}^{\beta} \tag{1}$$

Under the assumption of a Cobb-Douglas production function, the corresponding canonical parameters are modeled in the balanced growth path as a log-linear function of an outcome-specific set of predictors, as follows

$$\begin{aligned} \ln(y_{it}) &= \ln(A) + g + \frac{\alpha}{(1-\alpha-\beta)} \ln\left[\frac{sk_{it}}{(n_{it}+g+\delta)}\right] + \\ &+ \frac{\beta}{(1-\alpha-\beta)} \ln\left[\frac{sh_{it}}{(n_{it}+g+\delta)}\right] + \varepsilon_{it} \end{aligned} \tag{2}$$

where $\ln(sk_{it})$ is the log of output per worker, A captures the labor augmenting component, $\ln(sk_{it})$ and $\ln(sh_{it})$ are the fractions of income invested in physical and human capital, respectively, and $ln(n_{it} + g + \delta)$ is the sum of the domestic rates of change in population and technological progress plus depreciation.

The standard specification illustrated in **??** is criticised on several grounds. Brock and Durlauf (2000), Durlauf (2001), Durlauf, *et al* (2001), Durlauf and Johnson (1995), Liu and Stengos (1999), Kalaitzidakis *et al.* (2001), Masanjala and Papageorgiou (2004), under different approaches, reject the parameter homogeneity assumption implied by **??** due to the following problems: nonlinearity of the log transformed production function, parameters variability across countries and omitted variables. According to the first argument, the relationship between logs of output and logs of factors could be far from linear. In addition, the Cobb-Douglas function may not properly describe the efficient productive frontier, or each country or group of countries may have different production functions. A further issue which makes problematic the adoption of OLS is heteroskedasticity, even though this problem may be corrected through traditional parametric applications when standard OLS estimates are averages of the group specific parameters, iff no correlation exists among observations in the same group (Zellner, 1969). Unfortunately, the latter condition is unlikely to be met since economic series are seldom uncorrelated.[1] Furthermore, one of the most important problems in growth estimates is the unobserved heterogeneity which affects the estimated betas. The unobserved heterogeneity may not be a problem in a linear regression model if it is not correlated with the covariates and affects only the residual term. On the contrary, if it is correlated with covariates, estimated betas are biased. To cope with this problem Robinson (1988) proposes a quasi linear model in which the relationship between covariates and the response variable is affected by an unspecified nonlinear component. In the model this relationship is assumed to be only locally linear, but individual sources of heterogeneity are not modeled. Kalaitzidakis *et al.* (2001), adopt the partially linear regression approach to investigate if the relationship between human capital and per capita GDP is affected by unobservable heterogeneity.

Within this literature we propose an additional contribution. We start from the point developed by Cameron and Trivedi (2005) who demonstrate that, in case of unobserved heterogeneity, the βs are biased due to the omitted variable bias, unless the covariance matrix between the observed covariates and the unspecified nonlinear component is equal to zero. We then follow the advice of Aitkin *et al.* (2005) who suggest that a simple way to unify the different sources of model mispecification is through omitted variables. We propose to model the output variable through a Generalized Linear Model by adopting a semi-parametric technique based on finite mixtures and relaxing the assumption of *i.i.d.* residuals. To simplify the mathematical notation we define

$$\gamma_1 = \alpha, \quad \gamma_2 = \beta, \quad x_{1it} = \ln(sk_{it}), \quad x_{2it} = \ln(sh_{it}), \quad x_{3it} = \ln(n_i + g + \delta).$$

and we can rewrite (2) by modeling $\ln(y_{it})$ as a function of a set of three covariates

$$\mathbf{X}_i^{\mathsf{T}} = [(x_{1i1}, \ldots, x_{1iT})^{\mathsf{T}}, (x_{2i1}, \ldots, x_{2iT})^{\mathsf{T}}, (x_{3i1}, \ldots, x_{3iT})^{\mathsf{T}}]$$

[1]As it is well known, this problem is particularly severe in growth models where the inclusion of a lagged dependent variable causes autocorrelation problems since the latter is correlated with the error term. This renders OLS estimates biased and inconsistent, even when error terms are not serially correlated (Arellano and Bover, 1995 and Blundell and Bond, 1998). First generation first-differenced (Holtz-Eakin, Newey and Rosen, 1988 and Arellano and Bond, 1991) and second generation system GMM (Arellano and Bover, 1995) and Blundell and Bond, 1998) are currently used to overcome these problems.

With this respect, given the design vector $\mathbf{X}$ for unit i, we assume that some fundamental covariates are not considered in the model and that their joint effect can be summarized by adding a set of unobserved variables, u_i, to the linear predictor:

$$\ln(y_{it}) = \mu_{it} = \gamma_0 + \sum_{l=1}^{p} x_{itl}\gamma_l + u_i = \mathbf{X}_i^{\mathsf{T}}\boldsymbol{\gamma} + u_i \tag{3}$$

In (3) u_i appears additively in the model. However, we relax the additive assumption by associating random parameters to some elements of the adopted covariates set. In this case the previous model can be generalized into the *random coefficient* model in which a random coefficient can be associated to each of the covariates (for a detailed discussion, see Alfó and Trovato, 2004). By this manipulation we explicitly adjust the estimate of model parameters for country specific omitted variables which affect the relationship between economic growth and selected regressors with random components (i.e. cultural backgrounds, institutional factors, quality of the educational system, etc.). We postulate that variables whose effects are assumed to be fixed across countries are collected in the $\ln(sk_{it})$ and $\ln(n_i + g + \delta)$ terms, while those which vary over units are in $\ln(sh_i)$ and the intercept.[2]

The previous model can then be easily generalized in the following *random coefficient* model:

$$\ln{(y_{it})} = \mu_{it} = \gamma_{0i} + x_{1it}\gamma_1 + x_{2it}\gamma_{2i} + x_{3it}\gamma_3 \tag{4}$$

where $\gamma_{0i} = \ln(A)_0 + gt + u_{1i}$ and $\gamma_{2i} = \beta + u_{2i}$.

Equation (**??**) can be estimated through the marginal likelihood we will describe in section 3 (equation **??**) where the $\hat{\gamma}_{0i}$ and $\hat{\gamma}_{2i}$ vary across countries in order to capture country specific effects through the random effects u_{1i} and u_{2i}, for $i = 1, \ldots, n$.

Our approach is not entirely novel. Paap *et al.* (2005) apply it to a limited group of countries to evaluate heterogeneity in the intercept of growth equation. Alfó *et al* (2006), in a semi-parametric bivariate model, impose country specific heterogeneity showing consistency of the Solovian augmented model but finding no evidence of convergence to a unique steady-state equilibrium. In the present work, we show that unobserved heterogeneity may be determined by latent factors which affect returns to schooling and therefore should be taken into account by imposing the human capital

[2]Our choice is based on the stronger and more articulated interaction with cultural and institutional factors and with country fundamentals of the "living" human input with respect to the "dead" physical capital input. The assumption is also consistent with a significant part of the literature viewing education as a crucial factor of economic growth (Temple 2002). Further evidence based justification for our assumption is provided in section 3.

coefficient as random. Indeed, by establishing that $\gamma_{2i} = \beta + u_{2i}$, we assume that i) the impact of human capital investment varies across countries; ii) undetermined factors behind this variability are captured by the u_{2i}. In this context the random components in γ_{0i} and γ_{2i} represent the zero mean deviations from their fixed parts $\ln(A)_0 + gt$ and β, respectively.

3 Parameter estimation

To estimate our RC model we adopt the following maximum likelihood (ML) approach. Consider that, conditionally upon the parameters $\theta_i = [\gamma_{0i}, \gamma_1, \gamma_{2i}, \gamma_3]^{\mathsf{T}}$, the probability density function of $\ln(y)_{it}$ is

$$f(y_{it}|\theta_i) = \frac{1}{\sqrt{2\pi\sigma^2}} \exp\left\{-\frac{1}{2\sigma^2}\left[y_{it} - \gamma_{0i} - \gamma_1 x_{1it} - \gamma_{2i}x_{2it} - \gamma_3 x_{3it}\right]^2\right\} \quad (5)$$

while that of $\mathbf{y}_i = [y_{i1}, ..., y_{iT}]^{\mathsf{T}}$ is

$$f(\mathbf{y}_i|\theta_i) = \prod_{t=1}^{T} f(y_{it}|\theta_i) \quad (6)$$

We assume that θ_i is a random variable with probability function g. The marginal (unconditional) distribution of $\mathbf{y}_i$ is then given by the following integral

$$f(\mathbf{y}_i) = \int f(\mathbf{y}_i|\theta_i)g(\theta_i)d\theta_i \quad (7)$$

Treating the $\mathbf{u}_i$'s as nuisance parameters and integrating them out, we obtain for the likelihood function the following expression

$$L(\cdot) = \prod_{i=1}^{n}\left\{\int_{\mathcal{U}} f(\mathbf{y}_i|\mathbf{X}_i, \mathbf{u}_i)\mathrm{d}G(\mathbf{u}_i)\right\} \quad (8)$$

where $\mathcal{U}$ represents the support for $G(\mathbf{u}_i)$, the distribution function of $\mathbf{u}_i$. Due to the assumption of conditional independence among outcomes we have

$$f_i = f(\mathbf{y}_i|\mathbf{X}_i, \mathbf{u}_i)) = \prod_{t=1}^{T} f(y_{it}|\mathbf{X}_i, \mathbf{u}_i) = \prod_{t}^{T} f_{it} \quad (9)$$

Model parameters are estimated by adopting a non parametric maximum likelihood (NPML) approach (Laird, 1978)[3]

[3]On the consistency of the NPML estimators, Kiefer and Wolfowitz (1956) show that, by letting the probability density function of $\mathbf{u}_i = (u_{1i}, u_{2i})$ undetermined, we can correctly estimate the correlation between the two random effects. Furthermore, Lindsay

We do not assume a particular specification for the p.d.f. g, but we estimate it together with the other parameters. As demonstrated by Lindsay (1983a, 1983b), since the NPML estimate of a mixing distribution is a discrete distribution on a finite number of K locations, the likelihood function can be expressed as:

$$L(\cdot) = \prod_{i=1}^{n} \left\{ \sum_{k=1}^{K} f(y_i \,| \mathbf{X}_i, \mathbf{u}_k) \pi_k \right\} = \prod_{i=1}^{n} \left\{ \sum_{k=1}^{K} [f_{ik} \pi_k] \right\} \tag{10}$$

where $f(y_i \,| \mathbf{X}_i, \mathbf{u}_k) = f_{ik}$ denotes the response distribution in the k-th component of the finite mixture (which is assumed to be Normal). Locations $\mathbf{u}_k$, and corresponding masses π_k, represent unknown parameters, as well as the number of locations K, which is treated as fixed and estimated via formal model selection techniques.

The maximum likelihood estimates of model parameters are computed by using an EM algorithm (Dempster *et al.*, 1977 and McLachlan and Krishnan, 1997), which consists of two (Expectation and Maximization) steps. As it is well known (see, among others, Aitkin, 1996 and Wang *et al.*, 1996), the univariate EM algorithm maximizes the complete likelihood of (8) in the M-step. The EM algorithm starts by denoting with $\mathbf{z}_i = (z_{i1}, \ldots, z_{iK})$ the unobservable vector of components, where $z_{ik} = 1$, if the observation has been sampled from the component of the mixture, and 0 otherwise. Since the vector of components $\mathbf{z}$ is unobservable, it has to be treated as missing data. We therefore denote as incomplete the observed random sample $\mathbf{y} = (\mathbf{y}_1^\mathsf{T}, \ldots, \mathbf{y}_n^\mathsf{T})^\mathsf{T}$, while the complete-data vector is $\mathbf{y}_c = (\mathbf{y}\mathsf{T}, \mathbf{z}^\mathsf{T})^\mathsf{T}$. On this basis in the M-step we maximize the complete likelihood

$$L(\cdot) = \prod_{i=1}^{n} \prod_{k=1}^{K} \{\pi_k f(\mathbf{y}_i | \mathbf{X}_i, \mathbf{u}_i)\}^{z_{ik}} \tag{11}$$

Since the z_{ik} components are treated as missing data, in the E-step they are estimated by their expectations

$$\hat{z}_{ik} = w_{ik} = \frac{\pi_k f_{ik}}{\sum_{k=1}^{K} \pi_k f_{ik}}. \tag{12}$$

where $\hat{z}_{ik} = w_{ik}$ is the posterior probability that the i-th unit belongs to the $k - th$ component of the mixture. It can be shown that, at each step (E or M) the likelihood (11) increases. The complete likelihood is maximized with respect to a sub-set of parameters given the current values of the others. Hence, the log of the complete likelihood

(1983a, 1983b) stresses that, the NPML approach generates clusters characterized by homogeneous values of the random components, avoiding an exogenous sub-sampling process such as that used in threshold regression models. It has to be noted that this kind of classification is possible and successful only if country heterogeneity exists (Lindsay, 1983a and 1983b).

$$\ell_c(\cdot) = \sum_{i=1}^{n}\sum_{k=1}^{K} \hat{z}_{ik}\left(\log(\pi_k) + \sum_i \log(f(y_i|\mathbf{u}_k))\right) \tag{13}$$

is maximized with respect to the π's and it reaches a maximum when

$$\pi_k = \frac{1}{n}\sum_i \hat{z}_{ik} \tag{14}$$

which represents a well known result of maximum likelihood in finite mixtures. Since closed form solutions of maximization of complete likelihoods are unavailable, we use a standard Newton-Rapson algorithm. The E and M-steps are alternatively repeated until the following relative difference

$$\frac{|\ell^{(r+1)} - \ell^{(r)}|}{|\ell^{(r)}|} < \epsilon, \qquad \epsilon > 0 \tag{15}$$

changes by an arbitrarily small amount if the adopted criterium is based on the sequence of likelihood values $\ell^{(r)}$, $r = 1, \ldots$. Since $\ell^{(r+1)} \geq \ell^{(r)}$, convergence is obtained with a sequence of likelihood values which are upward bounded. Penalized likelihood criteria (such as AIC, CAIC or BIC) have been used to choose the exact number of mixture components.

4 Selected variables and empirical findings

As specified above, our model allows to consider explicitly country specific components of economic growth which affect heterogeneity in the response of output to the standardly adopted observable inputs. By applying a random coefficient approach, we do not need to specify other regressors beyond those considered in the standard human capital augmented version of the Solow model, since we implement our model with random parameters which take into account the effects of these latent variables. In other terms, the model, by introducing a random distribution of parameters, allows us to estimate an unbiased coefficient of the intercept and the human capital regressor, conditional to the effects of additional unobserved environmental variables, even though these variables are not specified in the model. Moreover, since each country belongs to a given space on the log-likelihood with a proper posterior probability $\widehat{z}_{ik}$, the i-th country can be classified in the l-th group (component of the estimated mixture) if $\widehat{z}_{il} = \max(\widehat{z}_{i1}, \ldots, \widehat{z}_{iK})$. It is worth noting that each group is characterized by homogeneous values of the estimated random effects, i.e. conditionally on the observed covariates, countries assigned to that group have a similar structure. Furthermore, the assumption of conditional independence, which is at the basis of the finite mixture model approach, implies that the *global* production function may be obtained by weighting, say, K different functions corresponding to different groups in the analyzed samples. This implies that the coefficients estimated by assuming a common production function, as in the Generalised Least Square (GLS) approach, do not coincide with the coefficients of the mixture model which are weighted averages of those of the identified clusters with homogeneous production functions.

We estimate the model for non-overlapping 5-year periods between 1960 and 1995, with regressors being lagged four years with respect to the dependent variable. Five year periods are standard in the panel growth literature, given the trade-off between having enough degrees of freedom and avoiding the negative effects of the strong autocorrelation of dependent variables (Bond et al. 2001). Data are drawn from the Summers-Heston Penn World Tables (PWT). We estimate the model for the 1960-1995 period and for the sub-set of non oil countries. Our benchmark is the parametric analysis performed by Feasible Generalised Least Squares (hereafter also FGLS). In estimating (**??**) we use as dependent variable $ln(Y/L)$, that is, the logarithm of the real gross domestic product (real GDP) per total labor force. Among regressors $ln(sk)$ is the log of gross domestic investment over GDP, $ln(sh)$ is the log of the schooling years of the working population and $ln(ngd)$ is the log of the sum of the rates of change in population and in technological progress plus depreciation. One of the most important assumptions in our methodology is that countries' unobserved heterogeneity can be measured by random factors in the intercept and in the human capital coefficient. This assumption can be questionable. In principle, there are no reasons why the impact of physical capital investment should be fixed across countries and only those of human capital investment and the labour augmenting component should be considered as variable. To find support for our assumption, as a first step, we need to evaluate whether the above mentioned parameters are indeed heterogeneous when compared across countries. To this purpose we perform single country estimates and we show that the variability in parameters among the selected countries is quite strong (Table **??**). If we let all model parameters free to vary, the model becomes untractable given the complex structure of the correlation between latent effects and measured covariates. We therefore try to evaluate whether the hypothesis of the restriction on the homogeneity of the physical capital coefficient is acceptable. Following Aitkin (1997) and McLachlan and Peel (2000), we compare the fitted mixture distribution with the empirical distribution function. As it is well known, we cannot use standard parametric tests to investigate the goodness of fit of a mixture model. As stressed among others by Aitkin (1997) and McLachlan and Peel (2000), we can use as a diagnostic tool a plot comparing the fitted mixture distribution with the observed distribution function. In Figure **??**, 95% bands for the observed Cumulative Density Function (CDF) based on the usual binominal interval and the two fitted CDFs are shown. The estimated CDF based on the mixture model provides a close fit to the observed data. In contrast, the Feasible GLS-based

CDF shows substantial and significant departures from the observed CDF for several values of the log of per capita GDP. In Figure **??** we report the empirical density of log per capita GDP levels against those predicted by the GLS and RC models respectively. The empirical density reveals the likely presence of population heterogeneity. Moreover, by comparing the estimated and empirical densities, the mixture model seems to fit nicely and better the data generating process. The inspection of the twin-peaked shape of the density function of our dependent variable (Figure **??**) seems to support our hypothesis that GDP levels may be better modelled with a mixing distribution, where the observed density of the dependent variable is a linear combination of K different densities. We test this hypothesis by estimating equation (**??**) with a normal mixture using a semiparametric specification and comparing these results with those of the standard GLS panel fixed effect estimates which assume that the dependent variable is the outcome of a unique density function. Estimate results are presented in Table **??**. In identifying locations and prior probabilities of the mixture, we find that the Bayesian Information Criterion is maximized when considering the optimal number of five mass points. The locations of the five components of the mixture (mass-points) for the two random effects (intercept and human capital coefficient) are finally reported in Table **??**. In Table **??** the estimated within-country variance is .056 and its standard error .003. The hypothesis of parameter homogeneity is therefore rejected. The intercept random variance is 1.010 while the random variance of the human capital coefficient is .023. The Likelihood Ratio Test (LRT) shows that the effect of human capital varies significantly among countries with p-values equal to 0.001. [4] The hypothesis of heterogeneity in the impact on levels of per capita GDP is therefore not rejected also for this variable. After this first indication supporting the choice of a random coefficient model we compare its results with those obtained from the standard GLS estimates, where the effects of the intercept and the human capital coefficient on levels of per capita GDP are regarded as homogeneous across countries. With this respect, we observe that mean response coefficients of the independent variables under the random coefficient method are different from the coefficients obtained under the GLS method (the three regressor coefficients are smaller while the intercept is higher). Considering also that the semiparametric specification fitted much best the density function of the dependent variable rather than the parametric approach (Figure **??**), these findings confirm that the impact of human capital is heterogeneous across different countries or macroareas.

[4]The p-values are obtained by comparing the maximum likelihood of the full and constrained model. The latter is obtained by estimating the model without the random coefficient of the schooling variable

5 Interpretation of heterogeneous country parameters

We classify countries into different latent groups by allocating them into the group with the largest posterior probability. In other terms, we assign each y_{it} (the realized values of the GDP per capita for the i-th country) to the component of the mixture to which it is more likely to belong. The classification is made on the basis of the posterior probability estimates $\widehat{z}_{ik}$, which represent an important by-product of the adopted semiparametric approach. It is worth noting that each group is characterized by homogeneous values of the estimated random effects, i.e. conditionally on observed covariates, countries assigned to that group share the same estimated productive structure.

We identify five different clusters for which we report in Table 4 deviations of coefficients of the human capital and labour augmenting component from sample average values. Locations for the intercept and the schooling random coefficients are similar to those obtained with the prior probabilities. After correcting for the observed deviations, all of the five subgroups maintain positive coefficients for both considered variables. The number of clusters is roughly consistent with what found in the literature (Liu and Stengos, 1999, Kalaitzidakis *et al.*, 2001, Masanjala and Papageorgiou, 2004; Paap *et al.*, 2005) and the differences are reasonable considering also that estimation periods and number of countries do not coincide. The approach closer to ours (Paap *et al.* 2005) identifies three groups but limits its analysis to Africa, Asia and Latinamerica. Note also that we exactly have three clusters for the countries included in the Paap *et al.*(2005) sample plus other two which contain almost only high income OECD countries which are not considered in the above mentioned paper.

In order to evaluate our findings it is important to consider that, according to our endogenous clustering process, a groups exists only if $\widehat{z}_{il} = \max(\widehat{z}_{i1}, \ldots, \widehat{z}_{iK})$. This implies that countries with similar unobservable economic structure, represented by the random component of the intercept and the human capital coefficient, have been classified in the same group if they have a higher posterior probability to belong to it.

Among the five identified groups, two groups (the first and the third) exhibit significant positive deviation from the average sample coefficient for both the human capital and the labour augmenting component. The first group, entirely composed by high income OECD countries, has a relatively higher intercept deviation and a relatively lower human capital deviation with respect to the third group. The latter is a mix of high income OECD countries of relatively more recent development (Finland, Spain, Ireland, Portugal) and of some emerging and Latinamerican countries (Brazil, Botswana, Argentina, Paraguay, Venezuela).

All the remaining three groups include only non OECD high income countries and exhibit negative deviations from sample averages for both the intercept and the human capital parameters (2, 4 and 5). Groups 2 and 4 are those with the highest negative deviation of the human capital coefficient, while group 4 has a much milder negative deviation in the intercept than the other two. Group 5 has the highest negative intercept deviation.

6 Factors associated to the observed model heterogeneity

To interpret these differences among groups, and to obtain richer policy insights from our approach, we calculate in Table 5 pairwise correlation coefficients between deviations from the average sample coefficients for the human capital and the labour augmenting component, on the one side, and measurable latent variables which we believe may affect these differences, on the other side. With this respect, our background theoretical references are represented by those approaches which have something to say on factors which may influence not only the Solow residual but also the output elasticity of human capital. Among them, we remember the role of a) economic freedom and quality of market rules on returns to human capital and contribution of the latter to growth (see, among others, Knack and Keefer, 1995; Temple, 1999; Rodrik, 1999; Barro and Sala-i-Martin, 1992 and Frankel, 2002). With this respect we believe that a crucial set of rules are those enhancing freedom of business, credit and labour markets which are crucial to the flexibility and reorganisation needed to transform technological innovation into higher returns to human capital [5]; b) cultural backgrounds affecting the relative social praise for productive versus rent seeking activities (Murphy et al., 1991); c) quality of education.

Given these a priori assumptions we consider as latent variables [6]: i) measures of political rights and civil liberties and qualitative indicators of freedom of credit, labour and business taken form the Economic Freedom of the World Annual Report (see Table 7 legend); ii) the share of protestant, catholic and muslim as measures of cultural background induced by religious beliefs related to point b; iii) the Hanushek and Kimko's (2000)

[5]The point is well illustrated by Brynjolfsson et al. (2000) when they describe the "Macromed" case study where "eventually the management concluded that the best approach was to introduce the new equipment in a "greenfield" site with a handpicked set of young employees who were relatively unencumbered by knowledge of the old practices given that old line workers still retained many elements of the now obsolete old work practices". In this example flexibility of the labour market facilitates the adoption of the new technology and higher output is likely to be obtained via higher returns to more skilled human capital.

[6]Average subgroup values for the identified latent variables are provided in Table 5 and a detailed legend is in Table 6.

educational quality index, conveniently normalized by Wossmann (2003) for each country relative to the measure for the United States and divided by domestic schooling years. As it is well known this index includes a weighted average of schooling years corrected for their estimated returns and further weighted with an average of cognitive test results on country students. The variable therefore combines microeconomic estimates of returns to human capital with an indicator (cognitive test results) which is a mix of innate abilities and quality of the education system. It can therefore be seen both as a test of the relevance of the quality of the education system and as an indirect test of the validity of our approach in which returns to human capital are derived from an aggregate point of view through the mixture model.

Table 5 shows that our indicators of institutional quality and economic freedom are all strictly and significantly correlated with deviations of both considered parameters (human capital and intercept) from their sample averages. This evidence is not in contrast with the hypothesis that returns to schooling and skill premia are typically enhanced in institutional frameworks which ensure freedom of credit, labour and business (the Credlabbus variable)[7], security of property rights (the legstrupro variable), political rights (civlib and polrights variables) and domestic credit development (domcrepri). All these variables are also correlated with intercept deviations and therefore their impact on economic development does not work only through the enhancement of returns to schooling[8]. Strong and significant, as expected, is also the correlation of the quality of education index with deviations from the human capital component. This implies that the additional flexibility of our mixture with respect to the standard homogeneous production function approach allows us to capture the richness of microeconomic evidence on returns to human capital and quality of education.

Religious variables are significantly correlated with both intercept and human capital deviations. This may indicate that the effect of cultural backgrounds on economic development is more pervasive and abstracts from

[7]To this point consider the contribution of Brynjolfsson et al. (2000) on the complex pattern of effects of ICT on productivity in which it is shown that ICT is "only a small fraction of a much larger complementary system of tangible and intangible assets". From this perspective ICT potential contribution to productivity and returns to schooling, to be fully exploited, requires a thorough reorganization of the firm and therefore freedom of credit, labour and business has a crucial role on it.

[8]Quality of institutions is widely acknowledged as a crucial determinant of economic growth (Knack and Keefer, 1995; Temple, 1999; Rodrik, 1999; Barro and Sala-i-Martin, 1992). Frankel (2002) considers that the success of market based economies crucially depends on good institutions and, more specifically, institutions which are crucial in protecting property rights, fighting corruption, supporting macroeconomic stabilization and promoting social cohesion. In a recent contribution in which within country variation in growth regimes is measured through a Markov-switching approach, Jerzmanowski (2006) finds that the role of institutions is that of making growth episodes persistent rather than ruling out growth take-offs.

the impact on returns to schooling. The unit scale of the three religious variables is the same so that we can compare the magnitude of the impact. Protestant and catholic shares are significant and positive, while the muslim share is significant and negative. In this respect, without entering into delicate value issues, we simply observe that our results are consistent with the hypothesis that a clash of cultural values with market values has a significant impact on economic development.

To sum up, the new perspective on economic development provided by our approach may give original insights also in terms of development policies. The most important of them is that economic development crucially depends on the role of "side" factors affecting returns to human capital, and, also, on deep cultural factors which do not operate only via returns to schooling.

The lesson for emerging and LDCs could be that higher quality of institutions and economic freedom and quality of education may crucially foster economic development by bringing returns to human capital to the level observed in more industrialised countries. However, more slowly varying cultural factors, and a lack of correspondence between cultural and market values, may still slow down economic development. With reference to our findings, if we go back to our three clusters which exhibit only negative deviations (2, 4 and 5) and relate these deviations to our latent factors we realize that the problem of "cultural distance" appears to be stronger for group 5 and much milder for group 4 in the observed sample period. This implies that countries in group 4 should have a relatively higher comparative advantage due to the proximity between their culture and the market values than countries in group 5.

7 Conclusions

Omitted variable bias and heterogeneity in the contribution of different inputs to levels and growth of per capita GDP are on top of the agenda among problems which need to be tackled in the empirical literature on economic development. The standard estimation approaches followed (fixed effects, first differenced and system GMM estimates, threshold regression models) only partially deal with these problems. In this paper we propose an approach, based on a non-parametric maximum likelihood estimation of a discrete mixing distribution on a finite number of mass-points, which deals with these issues. First, we show that our discrete mixing distribution may jointly address the omitted variable and parameters heterogeneity problems. Second, we provide evidence for the existence of hetereogeneity by testing and rejecting the homogeneity of some of the coefficients of standard factors affecting levels of per capita GDP in a human capital augmented Solow growth model. Third, we identify the "hidden submodels" by discovering five clusters of countries with their average random components for the human capital and the labour augmenting variables, expressed as deviations from the average overall sample coefficients. Fourth, we show that subgroup deviations from non random average parameters provide useful insights in the picture of economic development, have sensible interpretations and are significantly correlated with institutional indicators and other country fundamentals which the growth literature typically identifies among crucial factors of economic development. Our findings support the hypothesis that education has a crucial role in economic development and that economic policies affecting latent variables which enhance returns to human capital (such as flexibility in market rules, quality of education) may have relevant effects on economic development. In the same time, slowly varying factors such as scarce compatibility of cultural background with market values may significantly slow down this process. To sum up, we believe that the approach proposed in the paper has provided a unique and original perspective on economic development. By dealing more parsimoniously with available information than a simple homogeneous coefficient interaction model in which identified latent factors are introduced as multiplicative dummies on standard regressors, it has shown that the structure emerging by data is consistent with several theoretical explanations of economic development. Among them the ones who fit better the data seem those centered on the crucial role of human capital and on the effects that cultural background, institutions and quality of the educational system may have on its returns.

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Table 1: Summary Statistics of estimated regressors in sample country estimates

	Obs	Mean	Std. Dev.	Min	Max
Intercept	90	7.882	8.775	-45.175	33.033
$ln(sh)$	90	0.212	0.648	-1.978	1.99
$ln(sk)$	90	-0.026	0.857	-3.051	2.362
$ln(ngd)$	90	-0.527	3.730	-24.364	11.161

Legend: ln(sh): coefficient the schooling years of the working population; ln(sk): coefficient of the Summers-Heston corrected investment/GDP ratio; ln(ngd): coefficient of the sum of the rates of change in population and in technological progress plus depreciation.

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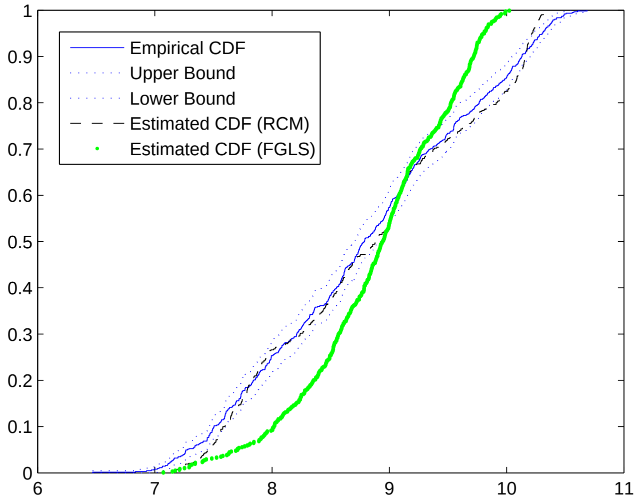

Figure 1: Cumulative Density Function (CDF) limits and RCM and Feasible GLS CDFs.

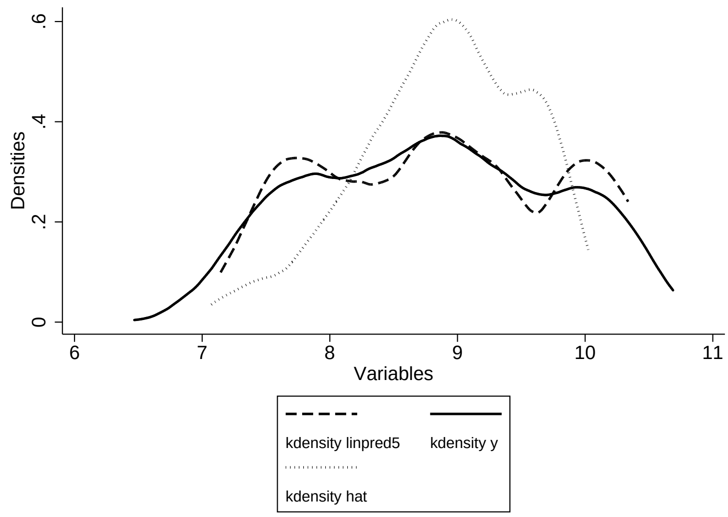

Figure 2: Observed and fitted kernel density distribution responses.

Table 2: A comparison of panel level estimates with Feasible Generalised Least Squares (FGLS) and Random Coefficient Finite Mixture Models (RCFMM)

	FGLS		RCFMM	
ln(y)	Coef.	Std. Err.	Coef.	Std.Err.
$ln(sh)$	0.350	0.023	0.264	0.028
$ln(sk)$	0.400	0.031	0.159	0.021
$ln(ngd)$	-1.238	0.093	-0.510	0.106
$Const.$	7.538	0.251	8.579	0.303
ℓ	22.581		-125.459	
Number of observations	530		530	
Number of countries	90		90	
Nc			5	
σ^2			0.056 (0.003)	
$\sigma^2\gamma_0$			1.010**	
$\sigma^2_{\gamma_{sh}}$			0.023**	
Penalized criterias				
AIC			261.920	
BIC			298.844	
CAIC			309.844	

Legend: ln(y): log of per capita GDP; ln(sh): log of schooling years of the working population; ln(sk): log of the Summers-Heston corrected investment to GDP ratio; ln(ngd): log of the sum of the rates of change in population and in technological progress plus depreciation ; Nc: number of mixture components which are selected by the BIC criteria; ℓ: log-likelihood;σ^2: within-countries (residual) variance; $\sigma^2\gamma_0$: variance of the random intercept; $\sigma^2_{\gamma_{sh}}$: variance of the random slope of $ln[sh]$. Time intervals in panel estimates: non-overlapping 5-year periods between 1960 and 1995, with regressors being lagged four years with respect to the dependent variable. ** 95 percent significance in a Likelihood ratio test in which p-values are obtained by comparing the maximum likelihood of the full and constrained model. The latter is obtained by estimating the model without the random coefficient of the schooling variable.

Table 3: Deviation from overall sample average values for coefficients of subgroups identified by the random coefficient finite mixture model (see Table 3)

Groups	Um_{γ_0}	$Um_{\gamma_{sh_i}}$	p	Freq.
1	1.392	0.181	0.199	12
2	-0.572	-0.117	0.198	16
3	1.061	0.216	0.16	11
4	-0.099	-0.110	0.266	14
5	-1.155	-0.097	0.176	46

Legend: The classification is made by allocating countries to the group with the largest posterior probability. Um_{γ_0}: deviation of the subgroup specific coefficient from the average sample coefficient for the labour augmenting component. $Um_{\gamma_{sh_i}}$: deviation of the subgroup specific coefficient from the average sample coefficient of the human capital parameter. p is the prior probability to belong to that group. Countries have been classified in that group only if $\widehat{p}_{il} = \max(\widehat{p}_{i1}, \ldots, \widehat{p}_{iK})$. Average sample coefficients are reported in Table **??**.

- Group 1: Australia, Austria, Belgium, Canada, Denmark, France, Israel, Italy, Japan, Netherlands, New Zealand, Norway, S.Africa, Sweden, Switzerland, U.K., USA.
- Group 2: Angola, Bolivia, Cameroon, Central Afr. Rep., Dominican Rep., Ecuador, Egypt, Honduras, Ivory Coast, Jamaica, Mauritania, Morocco, Mozambique, Philippines, Senegal, Syria, Thailand, Zimbabwe
- Group 3: Argentina, Botswana, Brazil, Finland, Greece, Ireland, Mauritius, Mexico, Paraguay, Portugal, Singapore, Spain, Trinidad&Tobago, Venezuela
- Group 4: Algeria, Chile, Colombia, Costa Rica, El Salvador, Guatemala, Jordan, Korea, Rep., Malaysia, Nicaragua, Panama, Papua N. Guinea, Peru, Tunisia, Turkey, Uruguay
- Group 5: Bangladesh, Benin, Burkina Faso, Burundi, Congo, Ethiopia, Ghana, India, Indonesia, Kenya, Madagascar, Malawi, Mali, Nepal, Niger, Nigeria, Pakistan, Rwanda, Sri Lanka, Tanzania, Togo, Uganda, Zaire, Zambia

Table 4: Pairwise correlation coefficients between γ_{0i}, γ_{sh_i} and selected country specific indicators

Indicators	γ_{0i}	γ_{sh_i}	obs.
Polrights	-0.7303*	-0.6489*	343
Econfreedom	0.6933*	0.6107*	343
Moneyacces	0.1295*	0.1103*	338
Freedomexc	0.5920*	0.4818*	376
Domcrebank	0.5350*	0.4258*	480
Domcrepri	0.6295*	0.5248*	481
Credlabbus	0.3064*	0.2348*	328
Legstrupro	0.6834*	0.6561*	342
Quality	0.2915*	0.3212*	343
Muslim	-0.3582*	-0.3640*	518
Cri_Prot	0.3148*	0.3046*	468
Cri_Cat	0.2636*	0.1744*	470

Legend: * 95 percent significance ; γ_{0i} measure of the subgroup specific Solow residual in the semi-parametric mixture model; γ_{sh_i} measure of the subgroup specific coefficient of Human Capital in the semi-parametric mixture model;

Table 5: Average values of selected country variables for the five groups

	1	2	3	4	5
Civlib	1.13	4.27	2.82	3.5	5.58
Polrights	1	4.45	2.73	3.7	5.58
Econfreedom	2	0.91	1.64	1.3	0.47
Moneyacces	6.62	6.25	7.15	6.5	6.44
Freedomexc	4.85	5.4	5.49	5.31	5.77
Domcrebank	27.06	44.24	43.23	60.56	41.24
Domcrepri	18.81	35.22	24.52	26.83	29.74
Credlabbus	5.21	5.29	4.85	5.43	5.19
Legstrupro	4.75	5.45	4.74	3.69	4.81
Quality	0.92	0.9	0.88	0.84	0.94
Muslim	19.4	14.3	20.52	28.31	20.09
Cri_Prot	0.15	0.13	0.1	0.07	0.14
Cri_Cat	0.39	0.32	0.43	0.4	0.27
Freq.	16	17	14	15	36

Legend: see tables 6-8.

Table 6: Variables used in the correlation analysis

Human Capital		
	Quality	Hanushek and Kimko's (2000) educational quality index normalized by Wößmann(2003) and divided by domestic schooling years. This index includes a weighted average of schooling years corrected for their estimated returns and further weighted with an average of cognitive test results on country students.
Governance		
	Moneyacces (Access to Sound Money)	Average annual growth of the money supply in the last five years minus average annual growth of real GDP in the last ten years Standard inflation variability in the last five years Recent inflation rate Freedom to own foreign currency bank accounts domestically and abroad
	Freedomexc (Freedom to Exchange with Foreigners)	Taxes on international trade Regulatory trade barriers Actual size of trade sector compared to expected size Difference between official exchange rate and black market rate International capital market controls
	Credlabbus (Regulation of Credit, Labor, and Business)	Credit Market Regulations Labor Market Regulations Business Regulations
	Legstrupro (Legal Structure and Security of Property Rights)	Judicial independence Impartial court Protection of intellectual property Military interference in rule of law and the political process Integrity of the legal system

Governance	Domcrebank	Domestic credit provided by banking sector (%GDP)
	Domcrepri	Domestic credit to private sector (% GDP)
	Statuslibe (Status of Liberties)	Combined Average of the Political Rights and Civil Liberties Ratings Country Status
	Civlib (Civil liberties)	Freedom of Expression and Belief Associational and Organizational Rights1. Rule of Law Personal Autonomy and Individual Rights
	Polrights (Political Rights)	Electoral Process Political Pluralism and Participation Functioning of Government
	Econfreedom (Economic Freedom index)	Size of Government (Expenditures, Taxes, and Enterprises), Access to Sound Money, Freedom to Trade Internationally, item Regulation of Credit, Labor, and Business and Legal Structure and Security of Property Rights
Source: Economic Freedom of the World: 2003 Annual Report (http://www.freetheworld.com)		
Religion	Muslim, Cri_Cat, Cri_Prot	Share of muslim, catholics (cri_cat) and protestant (cri_prot) believers in the given country
Source: Atlante Geografico De Agostini 2003		