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Education Composition and Growth: A Pooled Mean Group Analysis of OECD Countries

Summary: This paper uses the pooled mean group (PMG) estimator and a dataset restricted to OECD countries to examine the relationship between different levels of education, i.e. between education composition and growth. The PMG estimator allows a greater degree of parameter heterogeneity than the usual estimator procedures used in empirical growth studies by imposing common long run relationships across countries while allowing for heterogeneity in the short run responses and intercepts. Results point to a significant long-term relationship not only between higher education and growth but also between lower schooling levels and growth. This indicates that public spending on education in OECD countries should be spread across the different levels of education in a balanced way.

Key words: Levels of education, Economic growth, Dynamic heterogeneous panels.

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This paper sets out to examine the importance of level-specific educational investments i.e. education composition to economic growth. Thorough reviews regarding the empirical assessment of the importance of education to economic growth can be found in, for instance, Robert Topel (1999), Mark Bils and Peter J. Klenow (2000), and Barbara Sianesi and John van Reenen (2003). These generally show an overall positive association, although some criticisms have been made of these conclusions. We aim to enhance our knowledge of this relationship by investigating whether improvements in the different levels of education reflect differently on economic growth. The motivation for this kind of analysis comes from the need to identify the most efficient allocation of scarce public resources between the different levels of schooling. This necessity is clearly identified in the Organisation for Economic Co-operation and Development - OECD's 1998 report "Human Capital Investment. An International Comparison.", where it is stated that "The widespread acknowledgment of the benefits of education and other forms of learning should not lead governments and others to invest indiscriminately in human capital. In deploying finite resources, they need to know which forms of investment produce the best value for money." (p. 53).

The view that the link between education and economic growth is not the same across levels of education has its empirical roots in labour economics literature

regarding rates of return to education. George Psacharopoulos and Harry Patrinos (2004) provide a comprehensive review of these results. If, however, education has economic externalities in the form of expanding the technological frontier of a country, the overall economic benefits of education are better assessed through the study of the relationship between the different education or schooling levels and economic growth. Empirical growth studies that try to assess the relative importance of level-specific educational investments for growth include Norman Gemmell (1996), Robert J. Barro (2001), George Agiomirgianakis, Dimitrios Asteriou, and Vassilis Monastiriotis (2002), Panagiotis E. Petrakis and Dimitris Stamatakis (2002), Hans-Jurgen Engelbrecht (2003), Chris Papageorgiou (2003), Jérôme Vandenbussche, Philippe Aghion, and Costas Meghir (2006), and João Pereira and Miguel St. Aubyn (2009). These studies use a variety of human capital measures, country samples and different estimation procedures but the general picture that emerges is that the effect on growth depends on the schooling level under consideration with the importance of each education sub-category for growth varying with the countries' level of development. To investigate the relationship between schooling levels and growth an annual dataset comprising twenty-three¹ OECD countries over the period 1961-2000 was compiled. Empirical studies of the link between education and growth at different levels of schooling focusing on OECD or developed countries (see e.g. Gemmell 1996; Petrakis and Stamatakis 2002; Papageorgiou 2003; Engelbrecht 2003; Vandenbussche, Aghion, and Meghir 2006) have only been able to find a significant connection between higher education and growth. This is attributed to the higher level of technological development in these countries.

This paper contributes to the debate by exploring the validity of such findings using the pooled mean group (PMG) estimator developed by Mohammad H. Pesaran, Yongcheol Shin, and Ron P. Smith (1999). This particular estimator allows us to deal with an important problem that confronts empirical growth studies: that of parameter heterogeneity. This calls for extreme care in the interpretation of parameter averages (see e.g. Stephen Durlauf, Paul A. Johnson, and Jonathan Temple 2005). The PMG estimator provides a way of at least partly overcoming this problem by assuming homogeneity of the long run coefficients while allowing the short run coefficients and error variances to differ across countries. We focus on OECD countries because this group of economies most likely show similar behaviour. Including developing countries would make it impossible to uncover a specific influence for the different levels of education because of their widely differing levels of development. It would not be correct either to draw inferences regarding behaviour in OECD countries based on such results. Agiomirgianakis, Asteriou, and Monastiriotis (2002), Andrea Bassanini and Stefano Scarpetta (2002), and Jens Arnold, Bassanini, and Scarpetta (2007) are examples of studies that apply the PMG methodology to assess the importance of education to growth, finding robust results clearly pointing to a positive relationship.

Our results also indicate that when using estimation procedures similar to those of earlier empirical growth studies there is no evidence that either primary or

¹ Appendix A identifies all the countries included in the sample.

secondary education are significantly associated with growth in OECD countries. However, when we introduce a higher degree of parameter heterogeneity, as the PMG estimator allows, we do find a significant, positive relationship between lower education levels and growth. These results suggest that a balanced distribution of public investment across the different levels of education should be implemented by governments.

The remainder of the paper is organized as follows: Section 1 presents the empirical growth model under estimation, describes the data used, and briefly discusses the econometric approach that allows us to account for a greater degree of heterogeneity among the OECD countries that make up our sample. Section 2 presents and discusses the empirical results. Finally, in Section 3, we draw conclusions arising from our research.

1. Model Specification, Data and Econometric Approach

The specification used to test for the existence of a long run relationship between levels of education and growth can be derived from a human capital-augmented growth model like that of Gregory N. Mankiw, David Romer, and David Weil (1992) where each schooling level is entered as a separate input into production and takes the following form:

$$\ln y_{it} = \theta_{0i} + \theta_{1i} \ln s_{Kit} + \theta_{2i} \ln(n_{it} + g + d) + \theta_{3i} \ln(H_P)_{it} + \theta_{4i} \ln(H_S)_{it} + \theta_{5i} \ln(H_T)_{it} + u_{it} \quad (1)$$

where y is real output per worker, θ_{0i} a country-specific intercept, s_K the fraction of output invested in the accumulation of physical capital, H_P the stock of human capital resulting from primary education, H_S the stock of human capital resulting from secondary education, H_T the stock of human capital resulting from tertiary education, n the labour force growth rate, g the growth rate of exogenous technological progress, d the depreciation rate, u_{it} an error term, and i represents the country under analysis while t refers to the time period. For a detailed derivation of the structural specification using a single human capital input see Agiomirgianakis, Asteriou, and Monastiriotis (2002), and Bassanini and Scarpetta (2002).

This specification assumes a common rate of growth in technology across the OECD countries, in line with other studies that emphasize the role of human capital on growth, as employed by Agiomirgianakis, Asteriou, and Monastiriotis (2002), Bassanini and Scarpetta (2002), Arnold, Bassanini, and Scarpetta (2007), and Romain Bouis, Romain Duval, and Fabrice Murtin (2011). Robert M. Solow (2007), in an article that considers the main achievements and issues that remain to be clarified by growth theory, reminds us that this fairly standard assumption is in any case quite a remarkable one since it implies a belief that new technology diffuses around the world almost automatically. Solow also suggests that “Once growth theory abandons the implausible limitation to uniform TFP growth rates, it is natural to wonder about the actual pattern of national growth rates, and about the likely determinants of this pattern.” (p. 11). This leads to an interesting avenue for further research on the role

of human capital but which cannot be included in the scope of this paper. Furthermore, Kieran McQuinn and Karl Whelan (2007) argue, and provide supporting empirical evidence suggesting that this assumption of a common rate of growth in technology overemphasizes the role of factor accumulation in explaining cross-country income differences.

The data on output, labour force, and investment shares were taken from the AMECO database. For reasons of comparability we converted national figures at 1995 prices into 1995 purchasing parity (PPP) values. Real output per worker is measured as GDP per worker in 1995 PPPs. The fraction of output invested in the accumulation of physical capital is proxied by the ratio of gross fixed capital formation to GDP² and the effective labour force growth rate by the annual labour force growth rate plus 0.05. This reflects the common value assumed for the sum of the depreciation rate and the growth rate of technology (see e.g. Mankiw, Romer, and Weil 1992).

The data on human capital refers to the average years of schooling of the population aged 25 and over and was taken from the revised version of the Barro and Jong-Wha Lee human capital data set contained in Barro and Lee (2001). This human capital dataset is the most widely used in empirical growth studies and contains detailed information regarding the average years of schooling at the different schooling levels: primary, secondary, and tertiary. Following Ludger Woessmann (2002, 2003) we consider the average years of schooling to be the best available measure of the stock of human capital of the labour force³. This is also the human capital measure used in most of the empirical growth studies that relate directly to ours (see e.g. Gemmell 1996; Bassanini and Scarpetta 2002; Engelbrecht 2003; Vandenbussche, Aghion, and Meghir 2006; and Pereira and St. Aubyn 2009). The education data is provided at five-year intervals so we filled the gaps using linear interpolation to get annual data.

This paper explores the link between education and growth in the long run for a panel of twenty-three OECD countries ($N=23$) with annual data for the period 1961-2000 ($T=40$). Since both N and T are large, it is possible to choose from several alternative estimation procedures that imply different degrees of parameter heterogeneity.

Nazrul Islam (1995) showed that the estimation of growth regressions using cross-section data for large samples of countries with growth averaged over a long period of time leads to biased estimates. This is due to the presence of omitted variable bias as the initial technological level is unobserved and is thus included in the error term. Therefore, a fundamental explanatory variable in growth regressions, the initial income, is correlated with the error term. To overcome this problem Islam (1995) suggested using panel data and static fixed effect estimators that impose homogeneity of all parameters except the country-specific intercepts thus allowing control of unobserved, country-specific effects. Moreover, we cannot apply pooled OLS

² In this group of countries the investment share does not differ significantly from the ratio of gross fixed capital formation to GDP.

³ Nevertheless, this measure of the stock of human capital has been criticized for the fact that it does not consider differences in the quality of the education system between countries.

or static fixed effects estimators to growth regressions because of the presence of the lagged dependent variable, initial income. This acts as a regressor so that the coefficient on the lagged dependent variable is biased and will only be consistent for large values of T (see Stephen Nickell 1981). To deal with this issue, dynamic fixed effect (DFE) estimators were proposed and have been widely used in empirical growth studies (see e.g. Stephen R. Bond, Anke Hoeffler, and Jonathan Temple 2001).

In terms of parameter heterogeneity, the cross-section regressions impose homogeneity all over, while the fixed effects estimators only allow for intercept heterogeneity. Pesaran, Shin, and Smith (1999) show that in a dynamic model, although the parameter estimates based on the former are consistent, in case of coefficient heterogeneity pooled estimators are not consistent. This can lead to serious biases and the authors point out that dynamic fixed effects estimators “(...) can produce inconsistent, and potentially very misleading estimates of the average values of the parameters in dynamic panel data models unless the slope coefficients are in fact identical. (...) But tests on most panels of this sort, indicate that these parameters differ significantly across groups.” (p. 622).

Consider an ARDL (1,1,1,1,1,1) for observed output per worker, as in equation (1)⁴:

$$\begin{aligned} \ln y_{it} = & \gamma_i + \lambda_i \ln y_{it-1} + \delta_{10i} \ln s_{Kit} + \delta_{11i} \ln s_{Kit-1} + \delta_{20i} \ln(n_{it} + g + d) + \\ & + \delta_{21i} \ln(n_{it-1} + g + d) + \delta_{30i} \ln(H_P)_{it} + \delta_{31i} \ln(H_P)_{it-1} + \\ & + \delta_{40i} \ln(H_S)_{it} + \delta_{41i} \ln(H_S)_{it-1} + \\ & + \delta_{50i} \ln(H_T)_{it} + \delta_{51i} \ln(H_T)_{it-1} + \varepsilon_{it} \end{aligned} \quad (2)$$

The corresponding error correction equation is:

$$\begin{aligned} \Delta \ln y_{it} = & \phi_i (\ln y_{it-1} - \theta_{0i} - \theta_{1i} \ln s_{Kit} - \theta_{2i} \ln(n_{it} + g + d) - \theta_{3i} \ln(H_P)_{it} - \theta_{4i} \ln(H_S)_{it} - \theta_{5i} \ln(H_T)_{it}) - \\ & - \delta_{11i} \Delta \ln s_{Kit} - \delta_{21i} \Delta \ln(n_{it} + g + d) - \delta_{31i} \Delta \ln(H_P)_{it} - \delta_{41i} \Delta \ln(H_S)_{it} - \delta_{51i} \Delta \ln(H_T)_{it} + \varepsilon_{it} \end{aligned} \quad (3)$$

where $\phi_i = -(1 - \lambda_i)$ is the adjustment coefficient, $\theta_{0i} = \frac{\gamma_i}{1 - \lambda_i}$, $\theta_{1i} = \frac{\delta_{10i} + \delta_{11i}}{1 - \lambda_i}$,

$\theta_{2i} = \frac{\delta_{20i} + \delta_{21i}}{1 - \lambda_i}$, $\theta_{3i} = \frac{\delta_{30i} + \delta_{31i}}{1 - \lambda_i}$, $\theta_{4i} = \frac{\delta_{40i} + \delta_{41i}}{1 - \lambda_i}$, $\theta_{5i} = \frac{\delta_{50i} + \delta_{51i}}{1 - \lambda_i}$ are the long

run coefficients, and Δ is the first-order difference operator.

Pesaran, Shin, and Smith (1999) propose two estimating procedures, the Mean Group (MG) and the Pooled Mean Group (PMG), that allow for a higher degree of parameter heterogeneity in growth regressions than the other estimators described above. The MG estimator allows for heterogeneity of all coefficients, intercepts and slopes, by estimating a separate equation for each country while the coefficients for

⁴ The following exposition applies to a general ARDL (p, q, ..., q). Implementing the PMG method requires the specification of the appropriate lag order.

the whole panel are computed as unweighted averages of the individual coefficients. The Pooled Mean Group (PMG) estimator considers a lower degree of heterogeneity since it imposes homogeneity in the long run coefficients while still allowing for heterogeneity in the short run coefficients and the error variances. The basic assumptions of the PMG estimator are (see Pesaran, Shin, and Smith 1999): i) the error terms are serially uncorrelated and are distributed independently of the regressors, i.e., the explanatory variables can be treated as exogenous⁵; ii) there is a long run relationship between the dependent and explanatory variables⁶; and iii) the long run parameters are the same across countries. This estimator is also flexible enough to allow for long run coefficient homogeneity over a single subset of regressors and/or countries. It is also possible to test for the suitability of the PMG estimator relative to the MG estimator based on the consistency and efficiency properties of the two estimators, using a likelihood ratio test or a Hausman test.

2. Empirical Findings

Table 1 shows the results for the three alternative dynamic panel data estimation procedures described above, PMG, MG, and DFE, as well as those for a static fixed effects (SFE) estimator included here for comparison with earlier empirical growth studies. This method ignores the dynamic nature of the growth specification and represents a special case of the error correction model where the coefficient in the error correction term is constrained to be equal to one. Only long run coefficients are given in the table as these are the ones of interest in growth studies. The PMG calculations were obtained by estimating a common ARDL(3,3,1,1,1,1)⁷ for all countries under study.

The results show significant variations depending on the estimation method used, from MG (the least restrictive, but potentially inefficient) to PMG, and to DFE and SFE. The last two only allow the intercepts to differ across countries. Assuming that the long run coefficients are identical across countries, while allowing the short run elasticities to vary (i.e. using the pooled mean group estimator), there is significant support for the hypothesis that the different schooling levels are associated with growth in OECD countries in quantitatively different ways. In the first specification (A), the Hausman test on the long run coefficient of secondary education rejects the homogeneity assumption so the coefficient is left free in specification B, our preferred specification.

The sign of the different estimated coefficients does not change from the MG estimator to the PMG estimator (except for primary and secondary education) but the t-ratios are higher for the PMG estimates. The convergence coefficient is negative and significant as expected, a necessary condition for the existence of a long run rela-

⁵ This can be achieved by introducing a sufficient number of lags in the model.

⁶ See Appendix B for the results of the unit root and panel cointegration tests.

⁷ We used the Akaike, Schwarz and Hanna and Quinn lag selection criteria with a maximum lag order of 3 (in order to maintain sufficient degrees of freedom) to select the most appropriate ARDL model. The results from the three methods point to a lag order of 3 for output per worker and the investment ratio (since business cycle effects on both variables are probably similar), and a lag order of 1 for the effective labour force growth rate and the human capital variables (probably less affected by the business cycle).

tionship between the variables. The coefficient of the investment ratio is positive and significant but implies a rather high physical capital share. The coefficient of the effective labour force growth rate has the expected sign and becomes significant with PMG.

Table 1 Results for the ARDL(3,3,1,1,1,1) Model

PMG				MG	DFE	SFE	
A		B					
Long run coefficients	Hausman test /p-value		Hausman test /p-value				
logs _k	1.095 ^a (8.941)	0.35/0.56	0.996 ^a (8.736)	0.09 /0.77	0.896 ^a (2.496)	0.281 (0.932)	0.029 (0.283)
log(n+g+d)	-0.379 ^a -5.158)	1.05/0.31	-0.426 ^a (-5.418)	0.93/0.33	-1.235 (-1.470)	-0.597 ^a (-1.902)	0.081 ^a (1.78)
logH _P	-0.034 (-1.005)	0.07/0.80	-0.064 ^a (-2.10)	0.11/0.74	0.072 (0.173)	0.0284 (0.399)	0.046 ^a (1.62)
logH _S	-0.088 ^a (-2.951)	6.10/0.01	0.424 ^a (1.969)	free	0.428 ^a (2.026)	-0.0345 (-0.332)	0.06 (1.13)
logH _T	0.211 ^a (7.142)	0.09/0.77	0.190 ^a (6.835)	0.03/0.86	0.159 (0.896)	0.124 ^a (2.06)	0.379 ^a (7.62)
Error correction coefficient							
logy _{t-1}	-0.042 ^a (-5.160)		-0.049 ^a (-5.102)		-0.100 ^a (-2.546)	-0.037 ^a (-3.26)	
No. countries	23		23		23	23	23
No. observations	897		897		897	897	897
Log likelihood	2449		2475				

Notes: All equations include a constant country-specific term. The dynamic fixed effects OLS estimates have been used as initial estimates of the long run parameters for the pooled maximum likelihood estimation. t-ratios in brackets. ^a coefficients significant at least at the 5% level. Short run coefficients not reported for economy of space reasons.

Source: Author's estimations.

As far as human capital is concerned, the results here are consistent with previous work on the topic, but they also go further by improving on the results from earlier empirical growth studies that only find a positive and significant relationship between higher education and growth. The coefficient of human capital acquired through higher education is positive with both PMG and MG although only significant as expected with the former. The coefficient of secondary education is positive and significant in both cases, a result not usually found in other studies⁸. However, the coefficient of primary education, which is positive but not significant with MG, becomes negative and significant with PMG. This is a somewhat surprising result. One would expect to find no link because of the lack of variability of the proxy of human capital acquired through primary education due to the universal coverage of this schooling level across almost all OECD countries. A negative relationship was not expected. Since most primary education is publicly funded this result may indicate that, in the long run, the distortionary effects of taxation needed to finance primary education expenses overcome the positive effects of this level of education on growth. Comparing the PMG results with the most commonly used estimation proce-

⁸ These results imply an output elasticity of higher education of 18% and an output elasticity of secondary education of 40%. Both are quite high but do not contradict some of the previous evidence regarding rates of return to education, at least as far as higher education is concerned.

dures, DFE and SFE, only the results concerning the coefficient of the higher education variable are maintained across estimation procedures, being both positive and significant in all cases.

The regression was re-estimated for all the possible sub-samples obtained by deleting one country at a time from the original sample. Experimenting with these variations in the regressions, using the pooled mean group estimator, points to the robustness of the results regarding the education-growth link. The estimated coefficients obtained after arranging the estimates in decreasing order across sub-samples are shown in Figures 1-4. In the case of the coefficient of the investment ratio (Figure 1), the sample composition does not make a significant difference in terms of the estimated coefficient. In the case of the coefficient on primary education (Figure 2), its value becomes even more negative when Japan is excluded from the sample and less negative when the US is removed. In other situations it remains stable. In the case of the coefficients of secondary (Figure 3) and higher education (Figure 4), the results are remarkably stable except when the Netherlands is removed. In this case the coefficient of higher education remains significantly different from zero while the coefficient of secondary education becomes negative.

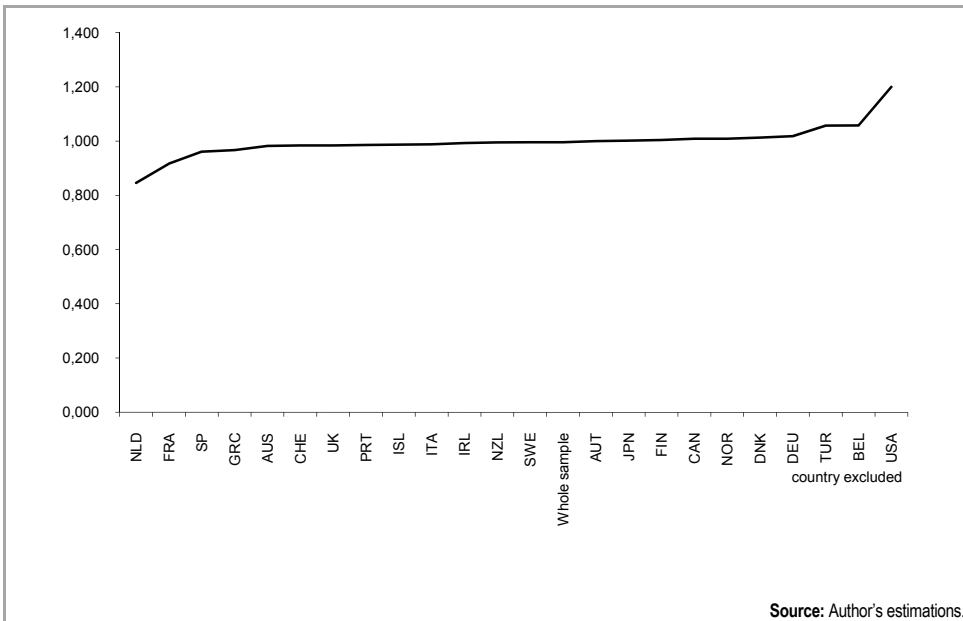


Figure 1 Sensitivity of the Coefficient of Logsk to Sample Coverage

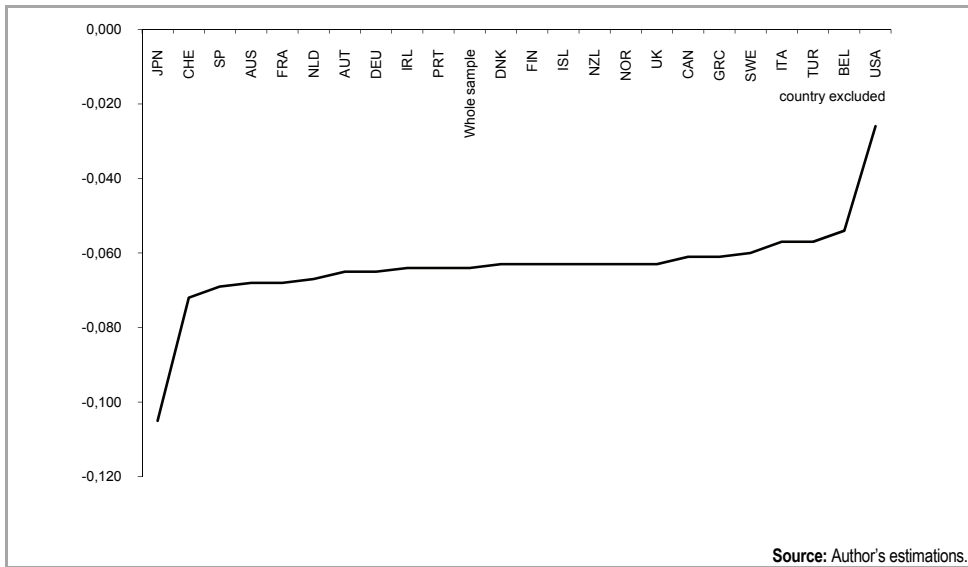


Figure 2 Sensitivity of the Coefficient of LogH_P to Sample Coverage

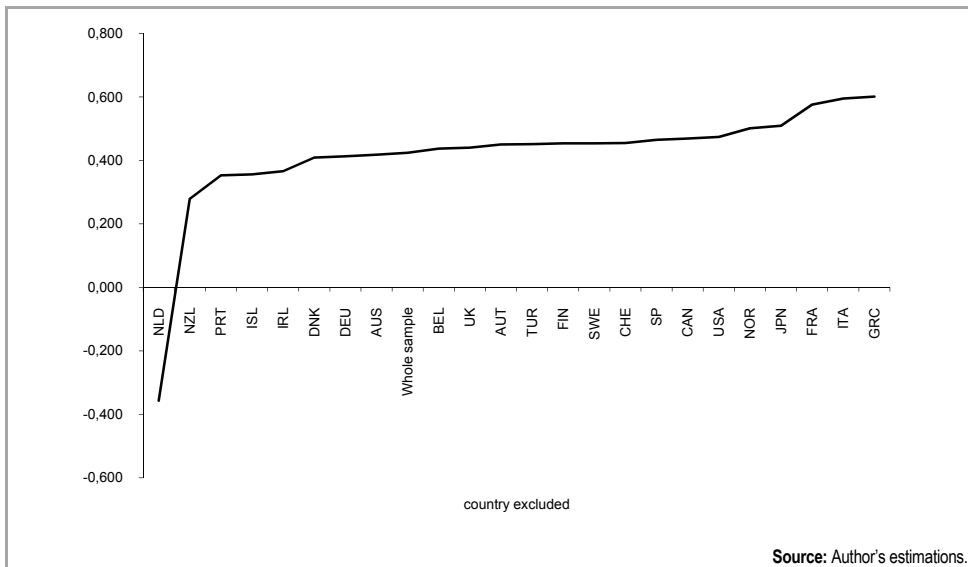
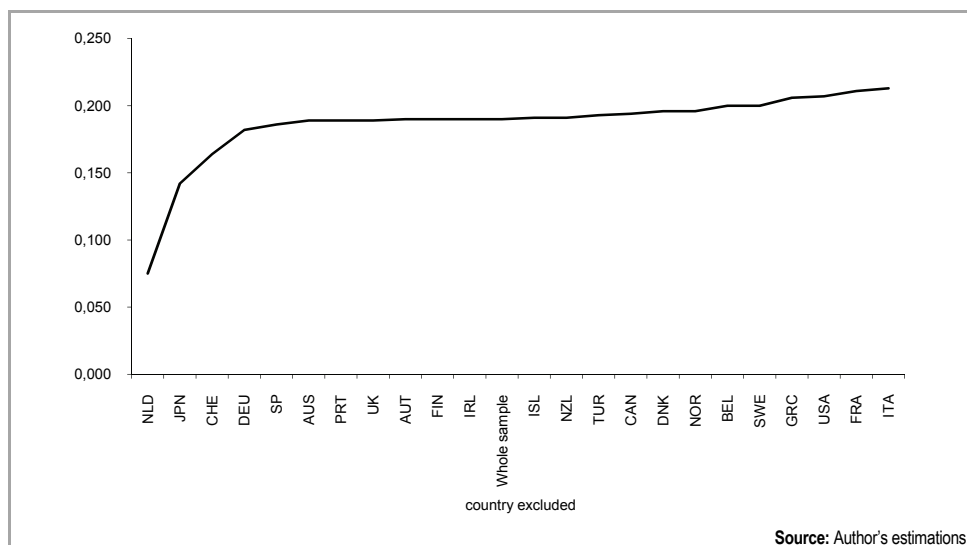


Figure 3 Sensitivity of the Coefficient of LogH_S to Sample Coverage



Source: Author's estimations.

Figure 4 Sensitivity of the Coefficient of LogH_T to Sample Coverage

We also conducted a sensitivity analysis of the PMG results to changes in the lag structure of the dependent and independent variables by re-estimating the regression using the Akaike (AIC) criterion to select the ARDL specification for each country, imposing a maximum lag order of 3 in order to maintain a reasonable number of degrees of freedom. Table 4 shows the results for this specification with the different estimation procedures.

Table 2 Results for the ARDL Specification Chosen through the AIC Criterion

	PMG		MG	SFE
	A	B		
Long run coefficients	<i>Hausman test</i> <i>/p-value</i>	<i>Hausman test</i> <i>/p-value</i>		
$\log s_k$	0.398 ^a (6.076)	0.510 ^a (6.942)	-0.288 (-0.336)	0.029 (0.283)
$\log(n+g+d)$	0.155 ^a (2.360)	0.207 ^a (3.149)	6.444 (0.961)	0.0814 ^a (1.78)
$\log H_p$	0.145 ^a (4.558)	0.115 ^a (4.137)	2.543 (1.05)	0.0462 ^a (1.62)
$\log H_s$	-0.198 ^a (-4.974)	-0.063 (-0.288)	0.554 (1.440)	0.06 (1.13)
$\log H_r$	0.621 ^a (20.277)	0.695 ^a (22.05)	-0.657 (-0.691)	0.379 ^a (7.62)
Error correction coefficient				
$\log y_{t-1}$	-0.058 (-1.316)	-0.057 (-1.270)	-0.155 ^a (-2.50)	
No. countries	23	23	23	23
No. observations	897	897	897	897
Log likelihood	2378	2424		

Notes: All equations include a constant country-specific term. The dynamic fixed effect OLS estimates have been used as initial estimates of the long run parameters for the pooled maximum likelihood estimation. t-ratios in brackets. ^a coefficients significant at least at the 5% level. Short run coefficients have not been given due to lack of space

Source: Author's estimations.

The results show some sensitivity to the choice of the lag order. For our preferred specification and estimation procedure, i.e., the PMG results for specification B, we still get positive and significant coefficients of the investment ratio for higher education. However, the coefficient of the effective labour force growth rate is now both positive and significant, contrary to our predictions. The results concerning the coefficient of primary education were improved: now the coefficient is positive and significant as expected. The results concerning secondary education are, however, worse since the respective coefficient is now negative although not significant. The most problematic result concerns the error correction coefficient, which is negative but not significant.

3. Conclusions

This paper presents empirical estimates of the long run relationship between education and economic growth in a sample of twenty-three OECD countries focusing on the importance of heterogeneity for empirical results. We introduce heterogeneity in the relationship between education and growth through the consideration of the separate influence of each schooling level, primary, secondary and tertiary, and by applying a recently developed econometric technique: the pooled mean group estimator. This estimator allows the consideration of a higher degree of heterogeneity among countries than is common in other empirical growth studies. The PMG estimator considers country-specific effects as usual but also allows the short run growth responses of each country to vary, imposing only common long run relationships.

The results support the findings of previous studies that there is a significant positive link between higher education and growth in OECD countries. Furthermore, a significant positive relationship between lower schooling levels was detected, especially between secondary education, and growth. Considering the fact that in most countries the lower levels of schooling are mainly publicly funded, the major implication of this finding is that national governments should balance budgetary efforts across schooling levels. Despite the growth benefits from higher education, which are most likely due to the externalities associated with its role in promoting innovation (see e.g. Paul Romer 1990), OECD countries should not focus primarily or even increase their emphasis on costly higher education. This is because lower levels of education also seem to have their own independent impact, an indication that technology absorption remains an important growth promoting activity in this group of countries. The former (especially upper-secondary education) are necessary for the diffusion and effective implementation of new ideas (see e.g. Richard R. Nelson and Edmund S. Phelps 1966; Moses Abramovitz 1986; Jess Benhabib and Mark Spiegel 2005). It is generally understood that OECD countries need to train a scientific and technological elite but, at the same time, they need to maintain general access to the educational system at the primary and secondary levels, and not neglect either the quality or the efficiency of the respective educational systems (see e.g. OECD 1998; Kjetil Storesletten and Fabrizio Zilibotti 2000).

Attention should also be paid to the possibility that the different growth effects of human capital according to the level of education considered might in fact reflect a non-linear relationship between human capital and economic growth. It follows that

parameter heterogeneity in growth regressions can be the result of nonlinearities in the growth process (see e.g. Winford H. Masanjala and Papageorgiou 2004). For instance, Pantelis Kalaitzidakis et al. (2001), using data for 93 countries over three decades, find evidence that the relationship between mean years of total schooling and per capita GDP is nonlinear in the sense that only the countries with medium levels of human capital register growth benefits from human capital accumulation. The authors then disaggregate the influence according to levels of education, primary and post-primary. They find that increases in the average time spent in primary education have a positive impact on growth. However, this is only the case between 0.6 and 3.6 years of primary education. Post primary education seems to have no impact on growth in developing countries, whereas male post-primary education in developed countries is shown to have a positive impact. Kalaitzidakis et al. (2001) then augment a standard growth regression with a cubic polynomial in the different human capital measures. Their results using standard estimation methodologies show that this nonlinear specification fits the data better than the linear specification. No doubt our findings could be further enriched by extending the analysis in this way but this nonlinear specification is quite demanding on the data⁹ and alternative nonlinear specifications could also apply.

The results leave some further questions open for future research. Since the implied elasticities for physical and human capital are rather high, future research should consider endogenous growth specifications of the growth equation. There is also the need to include in the analysis the quality dimension of human capital investments. Finally, a more systematic analysis of the causality relationship between education and growth is in order, involving for instance the estimation with the PMG estimator of additional equations where the different levels of education are the dependent variable. Negative and statistically significant coefficients of the error correction mechanism in these equations as well as in the growth equation would provide evidence of bi-directional causality.

Bearing in mind the limitations of the analysis, the evidence presented here points to the need to develop empirical growth studies that consider the existence of a higher degree of heterogeneity across countries and over time than the more traditional estimation procedures. Norman V. Loayza and Romain Ranciere (2006), Ibrahim A. Elbadawi, Linda Kaltani, and Klaus Schmidt-Hebbel (2008), Asteriou (2009), Kang Yong Tan (2009), Vítor Castro (2010), Houdou Ndambendia and Moussa Njoupouognigni (2010), are examples of empirical growth studies that apply the PMG estimator to investigate the impact of growth determinants such as financial development or foreign aid finding more robust results. Additionally, this methodology should be able to produce further insights into the growth process and respective sources in transition economies, taking a little further the previous studies that include this group of countries, such as Thorsten Beck and Luc Laeven (2006), Michael Spagat (2006), Garbis Iradian (2007), Ranjpour Reza and Karimi Takanlou Zahra (2008), Fabienne Bonetto, Srđan Redžepagić, and Anna Tykhonenko (2009).

⁹ Especially since the authors simultaneously include a fourth degree polynomial in initial income.

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Appendix A Countries in the Sample

Australia	Austria	Belgium	Canada
Denmark	Finland	France	Germany
Greece	Iceland	Ireland	Italy
Japan	Netherlands	New Zealand	Norway
Portugal	Spain	Sweden	Switzerland
Turkey	United Kingdom	United States	

Appendix B Panel Unit Root and Cointegration Tests

As a preliminary step to the estimation of the empirical model it is necessary to evaluate whether the series are stationary, which involves using a panel unit root test. We apply one of the most commonly used test procedures proposed by Kyung So Im, Mohammad H. Pesaran, and Yongcheol Shin (2003), henceforth IPS, which consists of a generalization of time series unit root tests to panel data. The IPS test assumes that all series are non-stationary under the null hypothesis but allows for heterogeneity in the autoregressive coefficient, which is assumed to change freely among the countries. Table B.1 displays the results of the panel unit root tests for each variable. It can be seen that most of the variables may be considered as non-stationary or integrated of order one, $I(1)$, at a significance level of 1%. The only exception is $\ln(n+g+d)$ that seems to be stationary. We thus proceed with our analysis assuming that the series are non-stationary.

Table B.1 IPS Panel Unit Root Test Results

	$\ln y^a$	$\ln s_k$	$\ln(n+g+d)$	$\ln H_p$	$\ln H_s^b$	$\ln H_T$
	<i>H₀: Unit root (assumes individual unit root processes)</i>					
IPS t-bar						
Level	-1.719	-1.596	-4.008***	-1.474	-1.265	-0.714
1 st difference	-4.337***	-4.655***	-8.382***	-2.27***	-2.126***	-2.374***

Notes: An *, **, *** indicates rejection of the null hypothesis at the 10%; 5%; 1% significance level, respectively, based on the appropriate critical values obtained in STATA 11.1. ^a Includes a trend in the level unit root test and cross-sectional means were subtracted. ^b Includes a trend in the level unit root test.

Source: Author's estimations.

The next step involves testing for panel cointegration in order to insure that there is a long run relationship between the variables that define the empirical model, which is a key issue for the consistency of the PMG estimator given that the variables are non-stationary. Peter Pedroni (1999, 2001) proposed a residual-based test that assumes a single cointegrating vector but allows the coefficients of each cointegration relation to differ among countries. Seven different statistics to test panel data cointegration are available. The first four are based on pooling and the other three on the between dimension. The estimates of the various cointegration statistics are presented in Table B.2. The cointegration tests broadly reject the null hypothesis of no-cointegration, with five out of the seven tests pointing to the conclusion that the series share a common long run trend and thus allowing the estimation of our empirical model with the PMG estimator.

Table B.2 Panel Cointegration Tests Results

	<i>H₀: No cointegration</i>
<i>Panel v-Statistic</i>	3.758***
<i>Panel ρ-Statistic</i>	1.983**
<i>Panel PP-Statistic</i>	-0.930
<i>Panel ADF-Statistic</i>	-2.918***
<i>Group ρ-Statistic</i>	3.468***
<i>Group PP-Statistic</i>	0.008
<i>Group ADF-Statistic</i>	-2.617***

Notes: All statistics have been standardized so that all reported values are distributed as $N(0,1)$ under the null hypothesis of no cointegration. These statistics must be in absolute value larger than the critical value to reject the null hypothesis of absence of cointegration for all units in the panel. An *, **, *** indicates rejection of the null hypothesis at the 10%; 5%; 1% significance level, respectively, based on the appropriate critical values. Country-specific intercepts and trends are included. All tests were performed with WinRATS v.8.00.

Source: Author's estimations.