

Gender Equality (f)or Economic Growth?

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on Economic Growth in OECD Countries

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ABSTRACT

This paper assesses the extent to which the increase in women's human capital, as measured by educational attainment, has contributed to economic growth in OECD countries over the past five decades. Using longitudinal cross-country data covering 30 countries from 1960 to 2008 on education (the Barro-Lee dataset) and growth (update of OECD data), our results point out a positive and significant impact of the increase in women's educational attainment relative to men on output per capita growth – as measured by GDP per capita. Our results are robust to the distinction between sub-periods and indicate that the effect of the equalisation of years of completed education on economic growth has been higher in the most recent periods. Results also hold when countries with an above-average increase in years of completed education are removed from the sample.

RÉSUMÉ

Ce papier évalue dans quelle mesure la croissance du capital humain détenu par les femmes, tel qu'il est mesuré par les années d'éducation, a contribué à la croissance économique des pays de l'OCDE au cours des cinq dernières décennies. Mobilisant des données couvrant 30 pays de 1960 à 2008 sur l'éducation (Base de données de Barro et Lee) et la production économique (données OCDE actualisées), nos résultats pointent un effet significatif et positif de l'augmentation du nombre relative d'années d'éducation des femmes par rapport aux hommes sur la croissance du PIB par tête. Nos résultats sont robustes à la distinction de sous-périodes, et semblent indiquer que l'effet de l'égalisation du nombre d'années d'étude sur la croissance économique a été plus important dans les années plus récentes. Les résultats sont aussi confirmés lorsque les pays ayant connu une croissance des années d'éducation significativement supérieures à la moyenne sont retirés de l'échantillon.

1. Introduction

Investment in human capital improves the economic and social opportunities of young individuals, thereby helping to reduce poverty and foster technical progress. In addition to the direct effects of education on economic participation, education – especially female education – also affects other societal outcomes such as child mortality, fertility, individual health outcomes, and the investment in the education and health of future generations. Investing in women's human capital is therefore key to economic growth and social cohesion, especially in developing countries where the gender gap in education is still large.

A large body of theoretical and empirical analysis exists on the link between investment in human capital and economic growth (see the literature surveyed in Thévenon *et al.*, 2012). The empirical evidence on the relationship between economic growth and gender equality in the distribution of education is instead not as conclusive. This paper tests this relationship in an internationally comparative perspective using a human capital growth model augmented to include the effect of gender inequality in educational attainment, using pooled longitudinal cross-country data covering 30 OECD countries over the 1960-2008 period.

The structure of the paper is the following. The next section provides a review of the relevant literature; section 3 describes the growth model; section 4 discusses the econometric approach for the growth model; data used in the growth model is reviewed in section 5; section 6 provides the results of the growth model before summarizing our main points in the conclusion.

2. Education and growth since the 1960s

It has now become a widespread argument in developing but also in more economically advanced countries that the economic gains from educating girls are greater than those from educating boys (Schultz, 2002). There are various reasons to believe that female and male education affect output levels and growth in different ways. Female education, just as male education, promotes growth by expanding the skilled working-age population and by improving the productivity of the female labour force (Mammen and Paxson, 2000) both directly – by raising output levels – and indirectly – through the increase in physical capital investment and technological change that follow from higher output levels (Barro and Sala-i-Martin, 1995). A balanced distribution of education among men and women is also likely to foster economic growth if male and female human capital are production factors with diminishing returns and are imperfectly substitutable (Knowles *et al.*, 2002). Moreover, higher education levels among women are argued to produce additional social gains by reducing fertility and infant mortality, increasing life-expectancy, and increasing the quantity and quality of investments in children education (Schultz, 1988). This *spillover* effect is expected to occur even if not all educated women enter the labour market and female labour force participation rates remain lower than those of males.

Over the past decades women have benefited from growing access to post-elementary and post-secondary education, especially in EU and other OECD countries (OECD, 2012b). Yet, there is so far not much empirical evidence and consensus on the extent to which the reduction in the gender gap in education has contributed to foster economic growth.

For instance, Barro and Lee (1994) and Barro and Sala-i-Martin (1995) – based on a panel of 138 countries – report the puzzling finding that years of schooling have an effect on economic growth that is positive for men but negative for women. They suggest that this is related to the fact that a high spread between male

and female secondary attainment is a measure of «backwardness» in the returns of education, and that higher female education attainment implies that countries have reached a stage of economic development from which no rewards can be expected from an additional increase in years of education (Barro and Lee, 1994). Their interpretation of the coefficients of male and female human capital is not straightforward, however, since they may reflect not only sex-specific educational attainment, but also the the gender gap in attainment (see below). Persistent gender differences in the chosen field of education and gender differences in the labour market returns to educational investment might be another explanation. Furthermore, the relationship between education, female employment and growth changes over time and through the process of economic development. The extent to which the aggregate economic output responds to female education and to the increase in female labour market participation is found to be U-shaped, *i.e.* positive only after countries have reached the “industrial” stage of development and women have gained access to more productive sectors of economic activity (Goldin 1995; Esteve-Volart, 2000; Mammen and Paxson, 2000; Lincove, 2008 and Tam, 2011). The increase in girls’ access to post-elementary education is certainly a driver of this process, as suggested by the evidence of a convex relationship between the decrease in gender discrimination in primary schooling and growth reported by Esteve-Volart (2000).

The Barro and Lee (1994) results have been challenged by several other empirical analyses. Dollar and Gatti (1999) show, for example, that the negative association between the years of schooling completed by girls and per capita output growth disappears once country-specific factors are controlled for. Even more challenging are those results of Caselli *et al.* (1996) and Forbes (2000) who come to the opposite conclusion. Using a Generalized Methods of Moment estimation to control for the potential endogeneity of human capital variables and measurement errors, female schooling is found to be associated with a statistically significant positive coefficient while a negative coefficient is obtained for male schooling. The authors interpret the positive sign for female education as the net effect of a positive impact due to the influence of education on fertility out weighting the negative “direct” effect of human capital found by Barro and Lee. The negative coefficient for males, by contrast, reflects only the human capital effect. Yet, the negative human capital effect remains ‘puzzling’ as it runs against most theoretical arguments (Topel, 1999; and, Krueger and Lindhal, 2001).

Esteve-Volard (2000), Klasen (2002) and Knowles *et al.* (2002) emphasize the problems affecting the empirical identification of effect of male and female education due to the multicollinearity caused by their close correlation. Their suggested empirical strategy involves using a direct measure of the gender gap in educational attainment in the growth equation instead of separate measures of male and female education. Klasen (2002) finds that gender inequality in the initial levels and in the expansion of education significantly reduces economic growth; the results – which are robust to different specifications of the education variables and control for possible simultaneous determination of human capital stocks and economic growths – are consistent with those of Knowles *et al.* (2002) and the older results by Hill and King (1995). Knowles *et al.* (2002) further suggest that female education has a statistically significant positive effect on the productivity per worker, while the role of male education is less clear and depends on the inclusion of benchmark (base-period) values of human capital and other control variables.

In addition to differences in the chosen data and econometric strategy, many other reasons can explain the conflicting results in the literature. Differences in stage of economic development can be one factor: less developed countries may experience higher growth rates in spite of lower educational attainment due to convergence mechanisms. Thus, the Barro and Lee (1994) results may be essentially driven by the

inclusion of the four East Asian Tigers and countries in Sub-Sahara Africa (Stokey, 1994; and, Lorgelly and Owen, 1999). This drawback motivates our choice to limit our sample to OECD countries, which are relatively homogeneous in terms of stage of development and education trends.

This paper attempts to reassess the influence of gender differences in educational attainment on the long-run steady-state of economic output. There are different reasons for doing so. First, we want to update the conclusions and allow for longer term assessment by taking into account recent trends in education. Our sample covers the 30 OECD countries from 1960 to 2008 and we use the education data published in the revised and cleaned Barro-Lee dataset (Barro and Lee, 2010). We also use the updated version of the data GDP per capita published by the OECD. Our goal is also to address some of the econometric issues that were overlooked by many of the previous assessments. Firstly, we tackle the assumption that the effect of the determinants of growth (including physical and human capital) is homogenous across countries. This assumption can be weakened by considering that countries will converge towards the same set of economic steady-states in the long-run, but at a different pace. This assumption should not be overly strong as we restrict our analysis to OECD countries, which have access to common technologies and share intensive intra-industry trade and foreign direct investment (Arnold *et al.*, 2011, p. 6) and are therefore quite homogeneous in terms of stage of development, specialization, technological and institutional settings (Pesaran and Smith 1995; Durlauf *et al.*, 2005; Pedroni, 2007; and, Eberhardt and Teal, 2011). This assumption is supported by Kourtellos (2011) and Di Vaio and Enflo (2009) that show how two growth regimes have emerged post World War II: a group of countries characterized by higher growth rates and convergence of per capita income (mainly OECD countries) and another group characterized by divergent and lower growth rates.

Against this backdrop, we model the steady-state level of output per capita as a function of the propensity to accumulate physical and human capital, the population growth rate, the level and growth rates of technological and economic efficiency, and the (constant) rate of depreciation of capital, as set by Mankiw *et al.* (1992) in the human-capital-augmented Solow growth model. The model is then tested with estimation procedures based on more or less flexible assumptions regarding the convergence process towards steady-states. We control for the incidence of country-specific and time-constant (but unobservable) factors shaping economic efficiency by means of a fixed-effect panel approach. The inclusion of country-specific time trends also allows us to capture changes in technologies or social institutions that affect economic efficiency, even though they are not explicitly modelled (Pedroni, 2007). We thus follow quite closely the perspective adopted by Bassanini and Scarpetta (2002) and Arnold *et al.*, (2010), but we add the gender dimension to the impact of educational attainment.

The next two sections present respectively the theoretical framework and the empirical setting. Basic statistics and data properties are then discussed, before presenting the regression results.

3. Theoretical background and econometric approach

The human-capital-augmented Solow growth approach first presented by Mankiw *et al.* (1992) provides an adequate framework to model the influence of education on growth. Following Arnold *et al.* (2011), we consider a human-capital-augmented Solow model with a standard Cobb-Douglas production function and account for short-run components as annual data is used to estimate the model (Pesaran and Smith, 1995; Bassanini and Scarpetta, 2002; and, Arnold *et al.* 2011):

$$\begin{aligned} \Delta \ln y(t) = & a_0 - \phi [\ln y(t-1) + \theta_1 \ln s_k(t) + \theta_2 \ln h(t) - \theta_3 \ln n(t)] + \gamma t + b_1 \Delta \ln s_k(t) \\ & + b_2 \Delta \ln h(t) + b_3 \Delta \ln n(t) + \varepsilon(t) \end{aligned} \quad [1]$$

where $y(t) = (Y(t)) / (A(t)L(t))$ and $k(t) = (K(t)) / (A(t)L(t))$ are the output and physical capital quantities per effective unit of labor; $h(t) = (H(t)) / (A(t)L(t))$ stands for the average human capital which sums the contribution of formal education completed by men and women; s_k is the investment rate in physical capital; and n is the growth rate of labor. Arnold *et al.* (2011) prove that this reduced form of output growth is also compatible with a two-sector AK model à la “Uzawa-Lucas” model of economic growth.

In this context, gender differences in education can be captured by including a log-ratio ($\ln R$) measuring the difference in years spent on average in education by women compared to men:

$$\begin{aligned} \Delta \ln Y(t) = & a_0 - \phi [\ln Y(t-1) + a_1 \ln S_k(t) + a_2 \ln H(t) + a_3 \ln R^{f/m}(t) - a_4 \ln N(t)] \\ & + \gamma t + b_1 \Delta \ln s_k(t) + b_2 \Delta \ln H(t) + b_3 \Delta \ln n(t) + \varepsilon(t) \end{aligned} \quad [2]$$

$R^{f/m}$ is defined as the female-to male ratio in the average number of years spent in the education system by men and women aged 25 to 64 ratio and is used to capture gender gaps avoiding the multicollinearity problems encountered when the average years of education of men and women are included separately. The growth equation described in [2] can be re-written within an error-correction model where growth rates are expressed as a function of the long-run evolution of the steady-state and of short-run variations, as appropriate for an empirical estimation based on pooled cross-country annual data (Lee *et al.*, 1997; Bond *et al.* 2004; Bassanini and Scarpetta, 2002; and, Arnold *et al.*, 2011). An alternative common technique to reduce the influence of short-run variation is to take averages of the data, typically over 5 years (Islam, 1995; Caselli *et al.*, 1999; and, Bond *et al.* 2001) but this would not fully take advantage of the complete set of information provided by the annual data.

$$\begin{aligned} \Delta \ln Y_{it} = & -\phi_i (\ln Y_{it-1} + \theta_1 \ln k_{it} + \theta_{2S} \ln H_{it} + \theta_{2R} R^{f/m}_{it} + \theta_{3m} \Delta \ln N^m_{it} + \theta_{3f} \Delta \ln N^f_{it} - \alpha_i \cdot t - \theta_i) \\ & + b_{1i} \Delta k_{it} + b_{2Si} \Delta \ln H_{it} + b_{2Ri} \Delta \ln R^{f/m}_{it} + b_{3mi} \Delta^2 \ln N^m_{it} + b_{3fi} \Delta^2 \ln N^f_{it} + \varepsilon_{it} \end{aligned} \quad [3]$$

where subscripts indicate country (i) and time (t) and ϕ_i is the country-specific speed of adjustments ($a_{ki}/\phi_i = \theta_k$). We thus assume that the steady-state of the growth rate of per capita output depends on country-specific factors that may shift the long-run path of economic development and/or produces short-run differences in the convergence towards the steady state of the economy.

This approach has the desirable feature of estimating a dynamic specification with country-idiosyncratic speed of convergence and deviations from the steady-state, while still allowing for country-specific production levels and growth rates. Another advantage is to overcome the fact that the output steady-states are unobservable, and observed changes in output per capita may well depend on shifts in the steady-state output per capita arising from other factors than technology. These circumstances make it necessary to have an empirical setting that clearly separate the long-run evolution of the steady states levels of output from their transitory variations.

The estimation of equation [3] also requires a choice on the allowed degree of heterogeneity across countries. To account for the fact that population and productivity growth patterns may vary considerably across countries (Lee *et al.*, 2007; Bassanini and Scarpetta, 2002b; and, Eberhardt and Teal, 2011) the model will be estimated with a Pooled Mean Group (PMG) estimator, which assumes that countries converge towards the same steady-state but the speed of adjustment can differ across countries (Lee *et al.*, 1997; and, Pesaran *et al.*, 1999). The PMG estimator is more restrictive than a Mean Group estimator – which does not impose long-run coefficients to be the same across countries – but it still allows short-run variations in the pace of adjustment.¹ This approach is consistent with the fact that the production functions of OECD countries are progressively becoming more homogenous on account of access to common technologies, intensive intra-industry trade and large foreign intra-OECD direct investment (Arnold *et al.*, 2011). Another advantage of the PMG estimator is that it is not affected by the ‘downward bias’ in the estimated coefficients that characterizes Dynamic Fixed-Effects estimators and Mean Group estimators when the lagged dependent variable is endogenous to the fixed effect in the error term (Nickell, 1981).

A few caveats are worth discussing before estimating the model. First, the main requirement to implement the PMG estimator is to have a large N-large T panel. As a consequence it is necessary to use panel data with annual observations, which will prevent us from using reduced panels with data averaged by periods of 5 years used by other studies on economic growth. Secondly, long-run growth accounting equations are relations between variables in levels, which may not be stationary. Regressions are likely to be spurious if the production function relating the variables in the long run is not a cointegrating vector, and the standard statistics used to assess the quality of adjustment (R^2 and standard errors) no longer apply (Philipps and Perron, 1999; and, Kao, 1999). For this reason, we employ an ECM which encompasses the true model: the levels terms are dropped out if there is no cointegration whereas they form a stationary relationship if there is cointegration. We also check that all panel residuals are stationary with the unit root test assuming that panel units are cross-section independent (Im *et al.*, 2003) or not (Pesaran, 2007).

Exogenous changes in the production technology and in the institutions that condition the efficiency of the production function can also introduce cross-country heterogeneity in the pace of convergence towards the steady-state. These differences cannot be observed and are proxied in our model by country-specific trends, which we define through a sequence of 5-years dummies. Time trends, however, have a limited accuracy in accounting for the unobservable changes if technology parameters don’t vary randomly across countries and are correlated with the regressors and/or the errors terms (Durlauf *et al.*, 2005). This misspecification can have serious implications if the observable and/or unobservable variables are non-stationary: spurious results may come from a failure to account for heterogeneity in the technology parameter, which leads to a breakdown of the cointegrating relationship between inputs and output (Kao, 1999; Smith and Fuertes, 2007; and, Eberhardt and Teale, 2011). This can occur, for example, when countries experience a common shock or are exposed to same processes (even if not with the same strength), which creates cross-section

¹ Another way to put it is that the Mean Group estimator consists of an unweighted average of country-specific (long-run) coefficients and yields consistent estimates but is very sensitive to outliers. The PMG can be instead thought of as a weighted average of individual group estimators, with weights proportional to the inverse of their variance; this allows for heterogenous short-run coefficients, but it constraints long-run parameters to be the same across countries. The PMG therefore exploits the efficiency of the pooled estimation while avoiding the inconsistency problem of pooling heterogenous dynamic relationships. The poolability restriction of the long-run parameters is tested using a Hausman type test applied to the difference between the MG and the PMG estimator (see Tables A2 and A3)

dependence between panel units. This can be detected with a test of the cross-section independence of residuals provided there are enough common observations across panels (Pesaran, 2004).²

Last, as in any panel regression using macro-level data, attention should be paid to possible endogeneity between economic growth and the accumulation of human and/or physical capital. Using stock data for human capital, lagged values of the explanatory variables and a rather long panel mitigates the reverse causality problem but it does not settle the issue completely since ΔY_{it-1} and ε_{it} might be correlated. However, as shown in Pesaran (1997) and further discussed in Pesaran and Shin (1999), if the independent variables have a finite order autoregressive representation, augmenting the ARDL specification with an adequate number of lags makes the estimation of the long-run coefficients immune to endogeneity problems. Pesaran further suggests that a two-step strategy whereby the lag orders of the ARDL model is first selected using either the Akaike Information Criterion (AIC) or the Schwartz Bayesian Criterion (SC), and then the long-run coefficients are estimated on the basis of the selected model, performs reasonably well in medium-sized samples. To satisfy this requirement, we checked the robustness of results to different lag structures in the independent variables and results are not different from those presented here. As shown in Annex Table A1, the best model according to the Schwartz Bayesian Information Criterion (BIC) would include one lag of the physical and human capital and the population growth. Only minor parameter differences emerge from this and our baseline specifications. Moreover, the key test statistics are robust to this type of changes in the specification (see the Annex –detailed results available on request).³

4. Data overview

One of the key features of education trends in the past decades is the drastic rise in women's educational attainment and the decline of gender inequalities in education that took place in most regions of the world. In OECD countries, primary school enrolment is nowadays nearly universal and gender equal (OECD, 2012b), while the picture is more mixed at secondary and post-secondary level. In the past decades the increase in post-secondary education graduation rates has been greater among women than among men across OECD countries, and – except for Turkey – boys are nowadays more likely than girls to drop out before completing secondary education. As a result, younger women increasingly have higher educational attainment than young men in the OECD (OECD, 2012b).

Qualitative differences in education persist, however. PISA data show that boys lag behind girls in reading skills at the end of compulsory education to the equivalent of a year of schooling, on average. In many countries boys are ahead in mathematics, but the gender gap is overall small compared with reading.

² An appropriate strategy to account for the incidence of the common unobserved factors is the Common Correlated Estimator proposed by Pesaran (2006), which includes cross-section averages of the dependent and independent variations in the regression equation. However,

³ GMM with lagged instrumental variables can also be used to address the possible endogeneity of education and/or physical capital. Nor the difference or the system GMM, proposed respectively by the Arellano and Bond (1991) and Blundell and Bond (1998), provide here a very convincing approach due to data specificities. On the one hand, by construction, yearly observation for education are obtained by interpolation, which implies that the use of lagged values of education provides invalid and/or weak instruments, which is problematic for the estimation of difference and system GMM models (Bond et al., 2001; Bun and Windmeijer, 2010; Bazzi and Clemens, 2013). On the other hand, the same argument and the superiority of PMG estimators over MG estimates (see next section on results and tables A2 and A3) refrain us from using Mean Group GMM estimators to address the possible endogeneity issues, as used for example in Bond *et al.* (2010) and Arnold et al. (2011).

Differences in the field of study are also quite large: girls are significantly less likely to choose scientific and technological fields of study in tertiary education, and when they do are then less likely to take up a career in these fields.

Results on the relationship between human capital accumulation and growth are sensitive to data quality and to the variables used to measure educational attainment and gender differences. The choice of data and indicators used to measure educational attainments and compare them across gender is crucial in driving results. To illustrate this point, Barro (1999) uses an updated version Barro and Lee (1996) data set on education and no longer finds a negative role for female education.⁴ In this paper we measure education as the average number of years of schooling attained by the adult male population over the average number of years attained by the adult female population. Different datasets can be used to obtain information on the level of education,⁵ but only few of them provide education data disaggregated by gender with sufficiently long time series: the data collected by Lutz *et al.* (2007) and those collected by Barro and Lee and updated recently (Barro and Lee, 2010). In this paper we use the latter because it provides longer time series – from 1960 to 2010 instead of 1970-2000 in the Lutz *et al.* (2007) – and because the 2010 version of the data has been revised to address most of the concerns raised by critics on the former versions, including those of Cohen and Soto (2007) and de la Fuente and Domenech (2006).⁶ The data are reported at 5-year intervals from 1950 to 2010; we use a linear interpolation between each reported value to obtain annual data.⁷ Educational attainment is measured as the number of years of education completed by men and women from 25 to 64 years (to limit the bias due to incomplete education). This indicator is preferred to enrolment rates because it captures the stock of education and because data on enrolment rates are affected by to cross-country differences and changes over time in the classification of educational attainment. The quality of cross-country data on years of education remains, however, not fully satisfactory. Morrison and Murtin (2013), who estimated their own database of growth rates of years of schooling, find inconsistencies with the Barro Lee dataset for a number of countries. Most notably, the years of schooling are rather low in a number of European countries. Unfortunately Morrison and Murtin (2013) don't provide data by gender so their dataset cannot be used for the purposes of this paper.

Table 1 reports the average evolution of years in education for men and women in the 30 OECD countries⁸ from 1960 to 2008, *i.e.* the period preceding the current economic crisis.⁹ The average number of years

⁴ To a large extent this results stems from the revision of the education data that was undertaken to improve its quality and improve the consistency of time series. The influence of data quality is also emphasized by De la Fuente and Domenech (2006) who find overall a positive correlation between the quality of data and the significance of human capital coefficients in growth regressions. Crespo Cuaresma (2005) finds important differences in the distribution and evolution of education in OECD countries across different datasets (namely the data collected respectively by Barro-Lee, Cohen-Soto and De la Fuente-Domenech).

⁵ See in particular de la Fuente and Domenech (2006) for a review of available data, and the data collected by Cohen and Soto (2007), Lutz *et al.* (2007) or Morriison and Murtin (2009).

⁶ More precisely, the data on education are derived from census and survey data obtained from Unesco and Eurostat used to construct estimates of various levels of educational attainment. Missing observations are estimated by extrapolating backwards and forwards from census and survey data. These estimates are corrected for mortality rates that are allowed to differ across different education cohorts. Preliminary testing shows, for OECD countries, that this version provides smoother time profiles for educational attainment in Norway and the United States than the former versions of the dataset (Barro and Lee, 2010).

⁷ Note that from the statistical point of view, this interpolation is likely to smooth the trends in educational attainment, and to some extent limit the remaining erratic movements which might be due to error measurement.

⁸ The countries covered are: Australia, Austria, Belgium, Canada, Chile, the Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Korea, Luxembourg, Mexico, the Netherlands, New Zealand,

spent in education by men and women has steeply increased, and there are small gender differences in the level of education (11.19 years for men and 11.21 for women). Yet, the increase has been steeper for women than for men.

Qualitative differences in educational attainment and trends are not completely captured by the evolution of years in education. Table 1 provides descriptive statistics on the proportion of men and women with completed secondary education.¹⁰ By this indicator, gender differences are quite significant even though the gap is closing on average (Annex Figures A1). Overall, the increase in completion of secondary education has been slightly steeper for women than for men: in the OECD the percentage of women aged 25 to 64 with completed secondary education was 11.7% in 1960 (vs. 16.6% for men), while the proportion in 2010 is at 54.5% slightly higher than for men (Table 1). Qualitative differences in educational attainment across gender remain therefore important despite the decrease in gender gap in terms of quantity of education and differentiating between levels of education may have an impact on results as found by Barro and Lee (1996), who identify a strong effect of secondary and higher schooling on growth. On the other hand, the 1997 revision of the education classification system poses significant consistency issues in the data on the level of education. For this reason and to limit the issue of multicollinearity that would arise by including more education variables, in this paper we will measure education only as the number of years of completed schooling.

Data on GDP, physical capital and working age population are taken from the OECD's Economic Outlook (No 90) data series. GDP per capita and Gross fixed capital formation (physical capital) are expressed in 2005 USD, taking advantage of the OECD's update of calculated time series which changed the base reference year from 2000 to 2005.

Table 4.1 also reports the average values of main variables entering in the growth equation: the dependent variable ($\Delta \ln Y$), measured as the growth in real GDP per head of population aged 15-64 years; the convergence variable ($\ln Y_{t-1}$), measured as the lagged real GDP per head of population aged 15-64 years; the propensity to accumulate physical capital ($\ln S_K$), proxied by the ratio of real private non-residential fixed capital formation to real private GDP; population growth ($\Delta \ln N$), measured in the working age population (15-64 years).

Norway, Poland, Portugal, the Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, the United Kingdom, and the United States. Country level figures are provided in Annex Figure A1.1.

⁹ Data series on GDP per capita and capital formation show indeed important breaks for years 2009 to 2010, as expected from the consequences of the economic crunch.

¹⁰ The same information could have been used for people with tertiary education. However, it is frequently the case that this latter is not completed before the age of 25, therefore introducing bias in the measurement.

Table 1. Basic Statistics

Variables	Year	Sample mean	Standard deviation
GDP per capita (in USD at 2005 PPP)	1960	16295	5711
	1990	35503	10622
	2010	45202	17948
Average years of education – Men	1960	6.49	2.16
	1990	9.65	1.72
	2010	11.19	
Average years of education – Women	1960	5.79	2.31
	1990	9.01	2.02
	2010	11.21	1.56
% of men age 25-64 with completed secondary education and above	1960	16.6	11.6
	1990	40.4	14.8
	2010	53.2	18.1
% of women age 25-64 with completed secondary education and above	1960	11.7	10.7
	1990	36.2	13.9
	2010	54.5	15.4
% per capita annual growth rate of capital		2.13	8.51
% annual growth of male working age population		1.00	0.85
% annual growth of female working age population		0.91	0.94

Sources: Barro and Lee (2010), OECD Economic Outlook (No 90).

Note: annual data obtained through linear interpolation of 5-year intervals data points.

A very high correlation is found (0.99) between male and female years of education (Table 2); although still high, the correlation is far weaker (0.63) between the total average years of education and the female-to-male ratio.

Table 2. Covariance matrix between education variables

	Male average years of education	Female average years of education	Total average years of education	Gender ratio in average years of education
Male average years of education	1.00	-	-	-
Female average years of education	0.99	1.00	-	-
Total average years of education	0.99	0.96	1.00	-
Gender ratio in average years of education	0.73	0.82	0.63	1.00

The issue of multicollinearity between our independent variables is investigated further with the computation of the Besley, Kuh and Welsh (1980)¹¹ statistics reported in Table 3, which clearly suggests that multicollinearity might be serious (condition index greater than 30) because the intercept and the variable measuring the formation of physical capital are collinear. These results confirm that collinearity

¹¹ The implementation of this “test” in a context of modeling economic growth was suggested by L’Angevin and Laib (2005)

between the average years of education and the gender ratio appears to be less of a concern despite relatively high correlation coefficient.

Table 3. Assessment of multicollinearity between the independent variables

	Condition index	Intercept	Years of education	Physical capital	Gender gap in education
1	1.00	0.00	0.00	0.00	0.01
2	2.18	0.00	0.00	0.00	0.37
3	24.51	0.08	0.94	0.04	0.57
4	60.08	0.92	0.05	0.96	0.06

5. Model specifications and results

Before presenting our results, we'll briefly discuss our choice of the estimation procedure. The PMG estimator is preferred for a number of reasons. The PMG estimator, which imposes long-run homogeneity to all slope coefficients but the time trend, yields lower standard errors and therefore more precise parameters compared to MG (see Tables A2 and A3 in the appendix). The measured speed of convergence is also significantly reduced, without changing the sign of the estimated long-run coefficients. Furthermore, a Hausman test comparing the PMG restrictions on long-run convergence against the parameters obtained by the MG estimation does not reject the former at the conventional statistical levels.¹² Log-likelihoods are also lowest with PMG estimations, which suggest that data are better predicted by this procedure. Additionally, the coefficient for education obtained with the PMG is much lower than those estimated by the MG, but significantly positive even when time dummies are included. The coefficient for education also remains highly positive when it is estimated with a GMM procedure to tackle the bias that reverse causality or omitted variables can potentially induce. Last but not least, only the PMG estimates yield stationary residuals.

Two additional features emerging from Table A2 suggest that the results are consistent with an endogenous growth model (Arnold *et al.*, 2011) and only partially corroborate the assumptions of the Solow human-capital-augmented framework. First, the assumption of decreasing return to scale for factors of production cannot be rejected by the specification without time trends, while the estimation including time trends suggests that return to scale are constant (*i.e.* $\hat{\lambda}$ is close to 1 and the test does not reject the null assumption of constant returns to broad physical and human capital). Secondly, the estimated speed of convergence parameter (respectively 0.21 to 0.33) is quite higher than the predicted speed predicted by the Solow model.¹³

These tests give enough confidence on the quality of the preferred PMG estimation and we now turn to model specifications that include the ratio measuring gender inequalities in completed years of education

¹² Results from a dynamic fixed effect model (DFE) are reported in Annex Table A2. The estimators obtained yield a much lower speed of convergence, which is consistent with the expected downward bias in dynamic heterogeneous panels. Moreover, restricting the short term dynamic to be homogenous (as with DFE) affects the sign and significance of the long-run coefficients.

¹³ These values are again close to those estimated by Bassanini and Scarpetta (2002) and Arnold *et al.* (2011) which for $\hat{\theta}$ range between 0.25 and 0.36.

(Table 1). Columns (1) and (2) present the results of the estimation on the overall sample of 30 countries for the years between 1960 and 2008; the baseline specification in column (1) only includes the overall level of education while the specification in column (2) includes the ratio $R^{f/m}$ of the average years of education of women relative to men. As a robustness check, we estimate the specification in column (2) for restricted periods of time: column (3) presents the results estimated on the 1960-1990 period, while column (4) presents the results for the 1984-1990 period. The last column presents the results of an additional robustness check carried out to verify whether results are driven by countries that had very high increases in the overall average years of education during the 1960-2008 period.

The results of the estimations for restricted periods in columns (3) and (4) suggest that important differences over time exist. During the first sub-period (1960-1990) the overall years of education have a larger effect than the gender ratio; in the mid-1980s and onwards, instead, the gender ratio has a much higher coefficient than the overall years of education. Column (5) presents the results of the estimation on a sample that excludes countries where the increase in the overall years of education was more than $\frac{1}{2}$ standard deviation above the average between 1960 and 2008. In the restricted sample both the overall years of education and the gender gap in education have a larger effect on growth (compared to the specification in column 2) suggesting that gender equality matters more in countries with low to moderate average increase in education.

All estimation procedures presented in Table 4 identify a convergence parameter with a negative sign, which is consistent with the assumption that variables converge to a long-run equilibrium. The estimated partial elasticity of output with respect to physical capital ($\hat{\alpha}$) is relatively limited, ranging between 0.23 and 0.32, and consistent with previous findings.¹⁴ The coefficients of the average education of the population and of the gender gap in education are positive and highly significant. The PMG estimation on the overall sample (column 2) suggests a growth-elasticity to the years of education of 0.94; since the average number of years in education has increased on average by 1.2% per annum (from 6.1 years in 1960 to 11.1 in 2008), human capital accumulation is estimated to have induced an increase in growth of 1.1% ($=0.94*1.2$) per annum. As GDP per capital actually grew by 2.1% per annum on average, the model implies that the increase in years of education accounts for about 50% of economic growth, of which just over half was due to the increase of educational attainment among women. The results also suggest that a balanced gender ratio in education ($R^{f/m}=1$) increases output per capita by around 0.8% in comparison to a scenario where women have no access to education. The estimated effect of the average years of education is slightly smaller in this case than those obtained when the gender ratio is not included, but it is balanced by a significant and quite large effect of greater equality in education between women and men.

¹⁴ As a benchmark, Mankiw *et al.* (1992) estimate $\hat{\alpha}=0.48$ and $\hat{\beta}=0.23$ using data from a group of 98 countries over the period 1960 to 1985. For OECD countries, Bassanini and Scarpetta (2002) and Arnold *et al.* (2011) find values between 0.13 and 0.22 for $\hat{\alpha}$, and between 0.52 and 0.82 for $\hat{\beta}$.

Table 4. Growth equation results

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	Pooled Mean Group (PMG)	Pooled Mean Group (PMG)	1960-1990	1984-2008	Without countries with high increase in years of education ⁽³⁾
$\Delta \log Y$					
Convergence coefficient	-0.28***	-0.33***	-0.39***	-0.52***	-0.30***
$\log Y_{t-1}$	(0.04)	(0.06)	(0.07)	(0.07)	(0.06)
Long-run coefficients					
$\log K$	0.28***	0.30***	0.32***	0.23***	0.28***
	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)
$\log H$	1.07***	0.94***	0.98***	0.40***	1.03***
	(0.06)	(0.07)	(0.07)	(0.03)	(0.08)
$\log R^{f/m}$..	0.81***	0.60***	1.42***	1.08**
		(0.16)	(0.18)	(0.16)	(0.19)
$\Delta \log N$	-1.57***
	(0.70)				
$\Delta \log N^m$..	0.82	1.71	0.91	0.69
		(1.05)	(1.52)	(0.76)	(1.24)
$\Delta \log N^f$..	-4.57***	-3.81**	-1.76**	-4.29***
		(1.34)	(1.67)	(0.86)	(1.60)
Diagnostics of residuals:					
Test of cross-section independence abs. ρ (p-value) ¹	0.86 (0.00)	0.88 (0.00)	0.90 (0.00)	0.84 (0.00)	0.93 (0.00)
Stationarity - unit root test ²	0.00	0.00	0.00	0.00	0.00
N. of countries	30	30	23	29	22
N. of obs.	1150	1127	620	617	446
Log likelihood	3184	3184	1882	2148	3185

Notes: Only long-run parameters are presented. Period effects are captured by 5-years time dummies, assumed to have country-specific effect.¹⁵

Standard errors in brackets. ***, **, *: significant at the 1%, 5% and 10% level, respectively.

1) Pesaran (2004) CD test, the null hypothesis assuming that all residuals are cross-section independent. Absolute correlation and p-value of the test are reported; a p-value below 0.05 leads the rejection of cross-section independence.

2) Results of the Pesaran (2007) CIPS tests which assume cross-section dependence between panel units; a p-value below

¹⁵ Eight (out of the potential ten) time dummies defined for sequence of 5 years are included in the set of the short-run regressors. This leaves enough degrees of freedom to run the regression without imposing linear trends, which would also create collinearity with the linear interpolation years of education used to complete the time series. In fact, the regression including linear time trends instead of our set of time dummies prove to bias the estimate of the coefficient of the years of education as part of its influence is in fact absorbed by the time trend.

0.05 does not reject the assumption that all residuals are stationary.

3) Australia, Canada, Czech Republic, Ireland, Israel, New Zealand, Norway, Slovak Republic and the United States are excluded from this specification.

All the reported estimations in Table 4 yield stationary residuals, so that there is no apparent problem of spurious results. On the other hand, the magnitude of estimated effect of human capital appears to be sensitive to the way time trends are modelled: models that include period effects yield significantly higher speed of convergence. The coefficients associated with physical capital also decrease when time controls are included, while those of human capital increase. In other words, the economic returns to education appear to be higher once other unobservable changes in the production function are controlled for, suggesting that economic growth is progressively more dependent on the prolongation of education. These unobserved factors, however, might not be well captured by time trends, as there is evidence of a cross-section dependence between residuals in all models of Table 4: the absolute correlation value is very high (ranging from 0.84 to 0.93), and the CD-Pesaran (2004) test constantly rejects the null assumption of cross-section independence.¹⁶

6. Conclusions

This paper provides an assessment of the extent to which the dynamics of economic growth over the past four decades prior to the ongoing recession are related to the increase in educational attainment of women. Both men and women have experienced an important increase in the number of years spent in education since the early 1960s. In many countries, the increase in the average numbers of years in education spent by women and men aged 25-64 has been roughly equivalent – the average number of years in education increased from 6.5 in 1960 to 11.2 in 2010 for men, and from 5.8 to 11.2 for women.

There are several reasons to argue that greater gender equality in education boosts economic growth. Assuming that boys and girls have a similar distribution of innate abilities and that children with more abilities are more likely to receive better quality and/or longer education, gender inequality in education implies that boys with lower abilities than girls are more likely to be enrolled in education. As a result, the average level of human capital in the economy would be lower than in a context of equal opportunities in education for boys and girls, which in turn might slow down economic growth. By the same reasoning, gender inequality in education reduces the impact of male education on economic growth and raises the impact of female education (Dollar and Gatti, 1999 and, Knowles *et al.*, 2002). It may also hinder economic growth by reducing returns on investments. Finally, greater gender balance in human capital also leads to higher growth if male and female human capital are imperfect substitutes and if the marginal returns to education are declining (Knowles *et al.*, 2002).

Our analysis is based on new OECD data series on GDP per capita and the updated version of the Barro and Lee (2010) data set on educational attainment of men and women. Our results support the assumption

¹⁶ Also, this assumption is rejected when cross-section averages of the dependent and independent variables are added as regressors in the CCE-MG specification (table A1). This specification, which is expected to sort out the issue raised by unobserved but correlated factors, fails to completely remove cross-section dependence although correlation is reduced – which is not surprising because the years of education have followed very similar trends in most countries and thus adding cross-section averages to the set of regressors generates multicollinearity of variables more than it helps wiping out the incidence of unobserved correlated factors.

that the increase in the number of years spent in formal education by the working age population in OECD countries has shifted up the steady-state of economic growth. Convergence to this steady state takes place at different pace, however, depending on population growth and investments in physical capital, technological and institutional change. The empirical specification of the model we retained, based on a Pooled Mean Group Estimation, suggests that one additional year of schooling in the population is estimated to raise output per capita by around 10% per annum, which is close to the estimate obtained by Arnold *et al.* (2011) or Eberhardt and Teale (2010) but lower than the upper bound suggested by Canton (2007). Overall, our estimation implies that the increase in years of education accounts for slightly more than 60% of output per capita growth, of which 34 percentage points result from the increase in years of education of women. These estimates, however, do not provide a fully satisfactory control of the unobserved but correlated country characteristics that potentially alter the influence of human capital on growth.

Although we found evidence that a more equal access to prolonged education raises growth rates, we are not able to explicitly identify the causes. One possibility, as suggested by Knowles *et al.* (2002), is that male and female human capital are both characterized by decreasing returns but are complementary, so that for certain values of the average of human capital stock, years of education of women are more rewarding than men. Another possibility is that women now perform better and are less likely to lack basic skill than boys and are thus more valuable in the labour market, despite persistent discrimination and professional segregation (OECD, 2012b; Hanushek and Woessmann, 2010). Other positive externalities of female education on the quality of life and productivity might also be at play, above and beyond their greater integration in the labour market. There is therefore much scope for further research on the subject.

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ANNEX

Figure A 1.1. Evolution of years in education – men and women aged 25-64

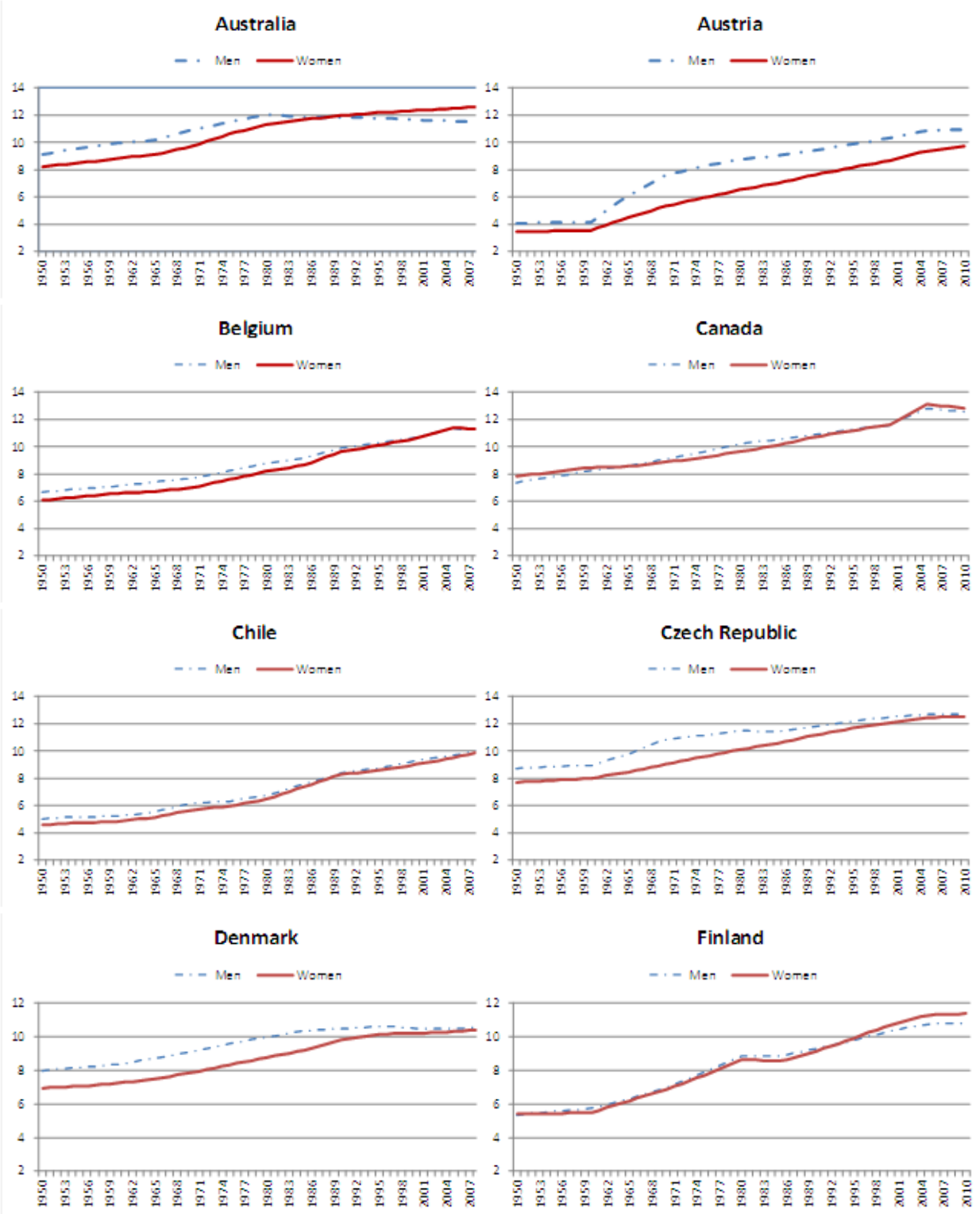


Figure A 1.1. Evolution of years in education – men and women aged 25-64 (cont.)

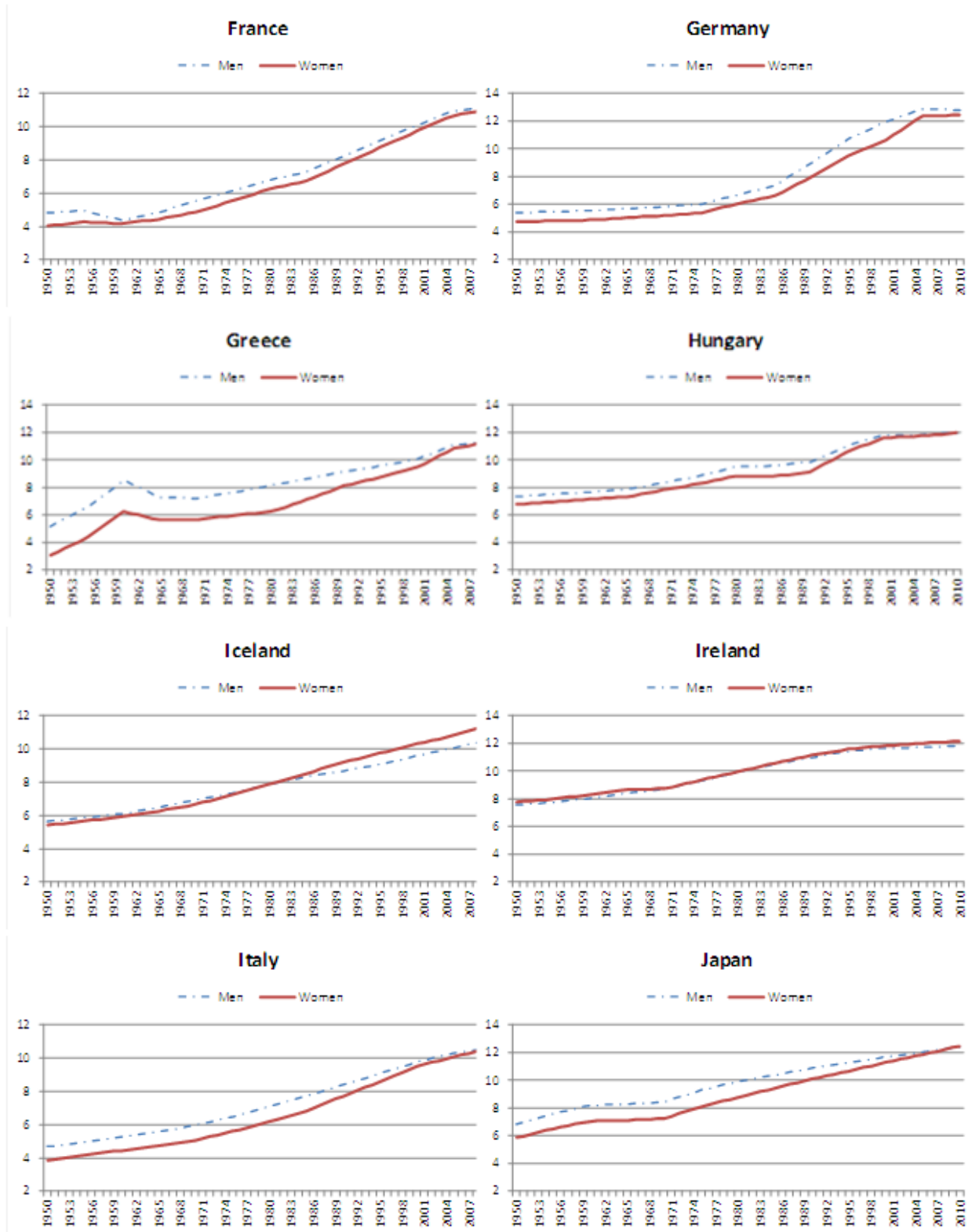


Figure A 1.1. Evolution of years in education – men and women aged 25-64 (cont.)

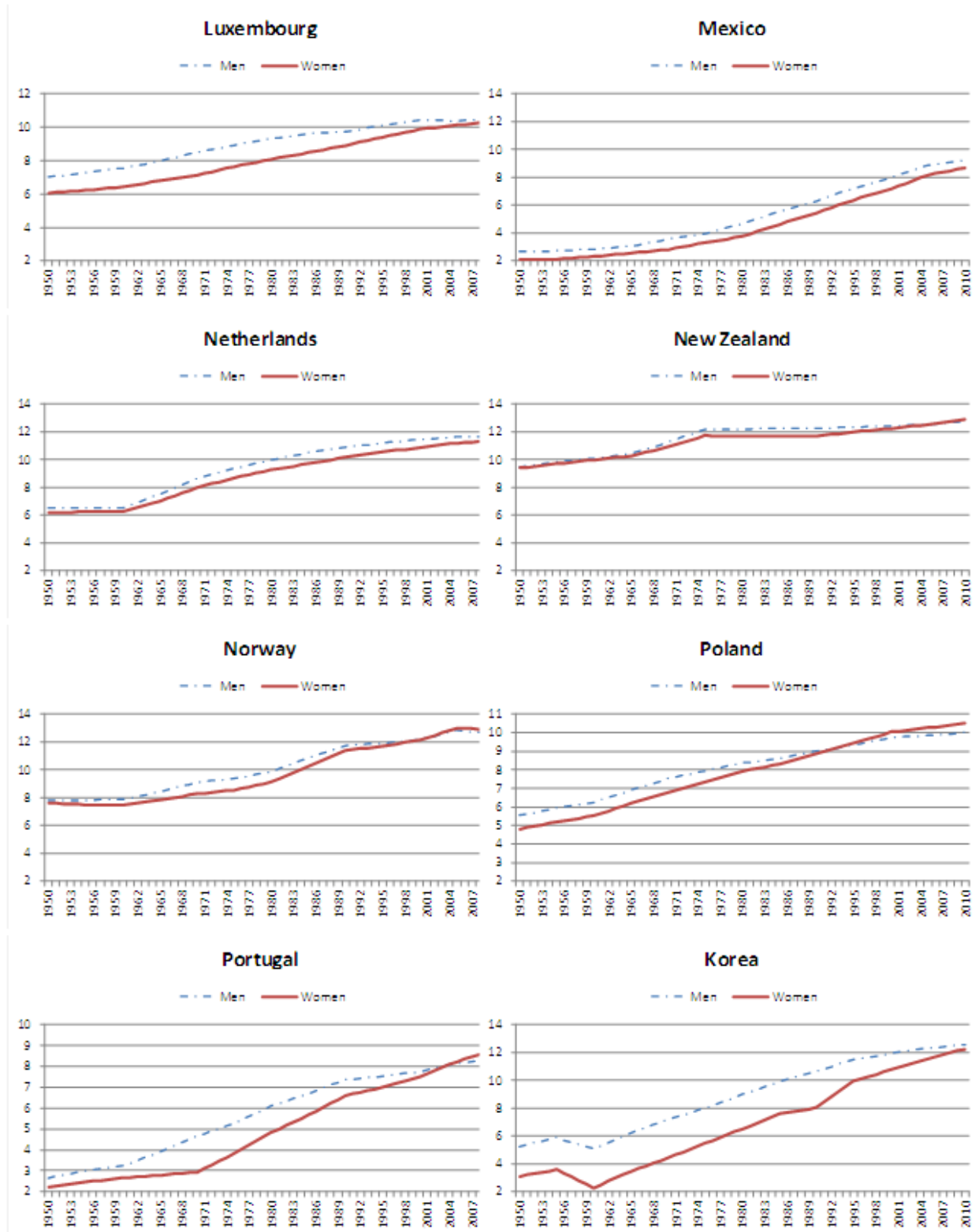


Figure A 1.1. Evolution of years in education – men and women aged 25-64 (cont.)

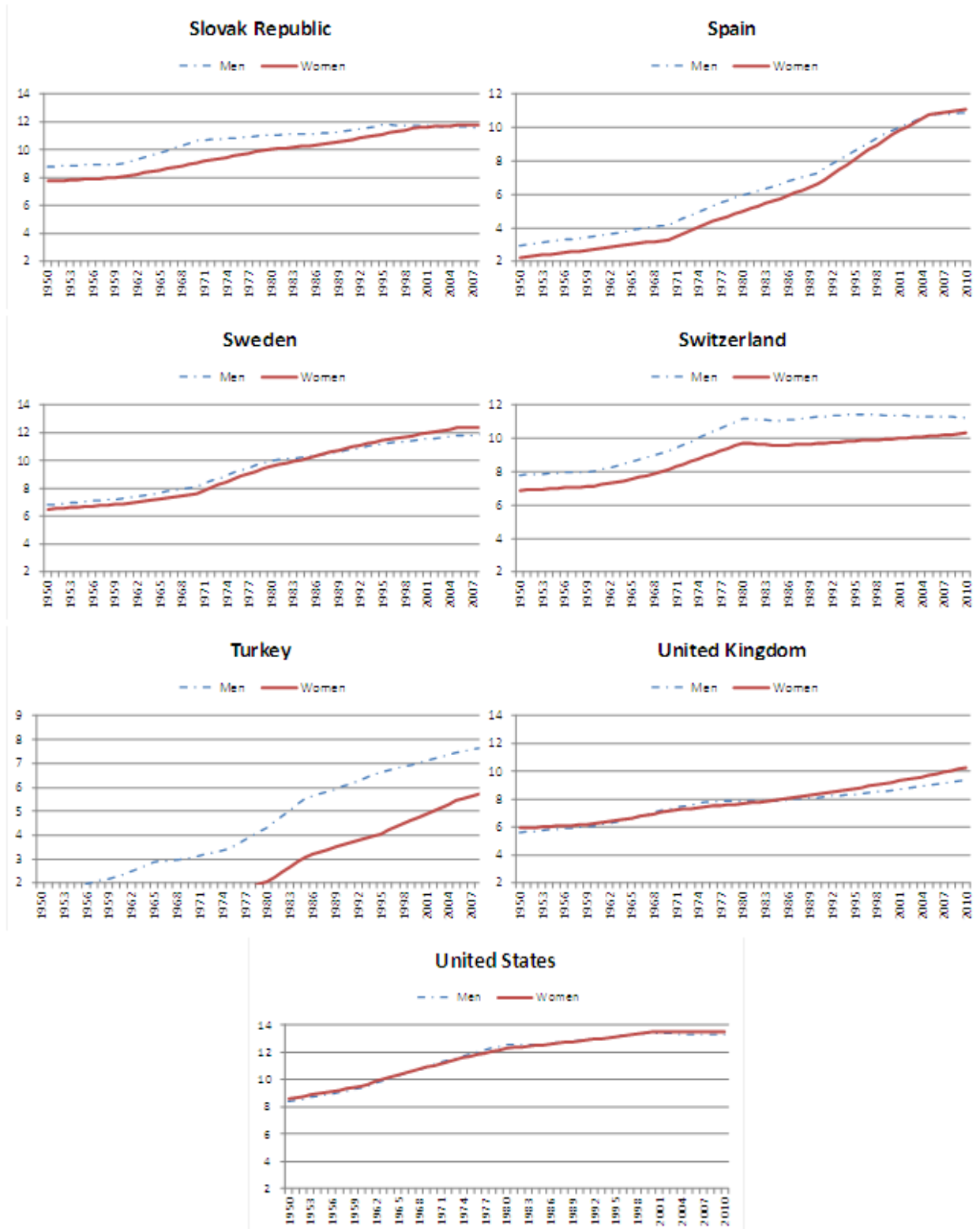


Table A 1. Comparison of specifications based on the Schwartz Bayesian Information Criterion (BIC)

Specification	Number of lags for the short-run dynamics			BIC
	Human cap	Physical cap	Population growth	
a (baseline)	1	1	2	-6656.247
B	1	2	2	-6701.624
C	2	2	2	-6753.304
D	1	1	1	-6737.156
E	1	2	1	-6699.823
F	2	2	1	-6553.937

Table A 2 Growth equation – baseline estimation

Dependent variable: $\Delta \log Y$	Mean Group (MG)			Pooled Mean Group (PMG)	
	(1) No time effects	(2) with 5-year dummies	(3) CCE-MG	(4) No time effects	(5) with period effects
Convergence coefficient $\log Y_{t-1}$	-0.18*** (0.05)	-0.38*** (0.06)	-0.36 (0.22)	-0.09*** (0.02)	-0.28*** (0.04)
Long-run coefficients log K	0.45*** (0.12)	0.41*** (0.09)	0.26*** (0.04)	0.47*** (0.02)	0.28*** (0.01)
log H	1.88** (0.81)	2.28 (1.55)	0.74*** (0.26)	0.32*** (0.04)	1.07*** (0.06)
$\Delta \log N$	-4.82 (21.83)	0.13 (15.9)	-1.52 (3.89)	-5.00*** (1.15)	-1.57*** (0.70)
Hausman test of long-run slope homogeneity (p-value)				0.28	0.45
Estimated returns to scale	1.60	-		0.54	1.06
Test for constant returns to scale (p-value)	0.30	-		0.000	0.24
Diagnostics of residuals:					
Test of cross-section independence abs. ρ (p-value)¹	0.89 (0.00)	0.89 (0.00)	0.83 (0.00)	0.77 (0.00)	0.86 (0.00)
Stationarity - unit root test²	0.06	0.19	0.00	0.00	0.00
Implied α^3	0.31	0.29	0.21	0.32	0.22
Implied β^3	1.30	1.61	0.59	0.22	0.83
Implied λ^3	0.20	0.48	0.45	0.09	0.33
Speed of convergence test (p-value)⁴					
(1)				0.002	0.000
(2)				0.14	0.000
N. of countries	30	30	30	30	30
N. of obs.	1150	1150	1150	1150	1150
N. of instruments	-	-	-	-	-
Log likelihood	3266	3512		3114	3184

Notes: Only long-run parameters are presented. Period effects are assumed to be linear or not and captured by 5-years time dummies. Standard errors in brackets. ***, **, *: significant at the 1%, 5% and 10% level, respectively.

1) Pesaran (2004) CD test, the null hypothesis assuming that all residuals are cross-section independent. Absolute correlation and p-value of the test are reported; a p-value below 0.05 leads the rejection of cross-section independence.

2) Results of the Pesaran (2007) CIPS tests which assume cross-section dependence between panel units; a p-value below 0.05 does not reject the assumption that all residuals are stationary.

3) Implied α is equal to $\hat{\theta}_1 / (1 + \hat{\theta}_1)$; $\beta = \hat{\theta}_2 / (1 + \hat{\theta}_1)$; λ speed of convergence is given by : $-\ln(1 - \hat{\theta})/s$, with s taken at 1.

4) Test for estimated speed of convergence to be compatible with the value predicted by the Solow augmented model. Predicted values are computed on the basis of plausible values for population growth rate, the depreciation rate and the rate of technological progress. (1) assumes a standard value of 2% for depreciation rate; for technological progress, we consider the average estimated time trend (-0.3%) or the average annual shift implied by the five-year dummies (0.1%), depending on the specification. For population growth, we take the average value of our sample (0.9%). (2) assumes a much higher value of capital depreciation (10%) and allow technological progress to growth at higher 3%.

Table A .3. Growth equation estimated with pooled mean estimator

	Baseline – balanced panel sample		Specification c	
Convergence coefficient				
log Y_{t-1}	-0.31*** (0.05)	-0.32*** (0.04)	-0.26*** (0.05)	-0.29*** (0.04)
Long-run coefficients				
log K	0.27*** (0.01)	0.29*** (0.01)	0.31*** (0.01)	0.36*** (0.02)
log H	1.09*** (0.06)	1.02*** (0.07)	0.87*** (0.7)	0.85*** (0.07)
log R^{l/m}	-	0.75*** (0.19)	-	0.83*** (0.17)
Δlog N	-1.78*** (0.67)	-2.29*** (0.65)	-2.26*** (0.74)	-2.81 *** (0.63)
Estimated returns to scale for reproducible factors	1.07	-		
Test for constant returns to scale (p-value)	0.132	-		
Diagnostics of residuals:				
Test of cross- section independence ρ (p-value)³	0.92 (0.00)	0.88 (0.00)		
Pesaran (2007) unit root test	0.00	0.00		

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