Human capital in a global and knowledge-based economy

APPENDIX

1. Human capital, productivity, and earnings: a survey of the microeconometric literature

a. Estimating the individual return to schooling: methodological issues

Schooling costs time, effort and money, but at the same time it augments an individual's earnings capacity. Hence, expenditure on schooling can be considered an investment and the extra income earned due to the completion of an additional year of schooling as part of its return.

Most recent studies of education and wage determination are embedded in Mincer's framework discussed in Box 1 in the main text and a large amount of research has estimated the Mincerian earnings equation for different countries and time periods. The common estimation method is ordinary least squares (OLS). Estimating returns to schooling by OLS is an easy task but has the drawback that the estimated return may be biased, i.e. the OLS return may not reflect the "true" reward the labour market places on an additional year of schooling. In what follows we first discuss why OLS estimates on returns to education may be biased and then discuss possible remedies.

When estimating returns to education by OLS, the econometrician encounters basically three problems. The first problem is that schooling attainment is not randomly determined, but is rather the result of an optimising decision influenced by individual characteristics such as ability, taste for schooling or access to funds. For example, if more able individuals spend more time in school and receive higher earnings, then differences in earnings of individuals who have different levels of education overestimate the true causal effect of schooling on earnings or put differently, the returns to education will be upward biased. The most straightforward approach to tackle the issue of unobserved ability is to include measures that proxy for unobserved ability in the earnings equation, such as IQ or other test scores (Griliches (1977)). But schooling itself determines this kind of ability measure leading to a downward bias in the estimated returns to schooling. Another method, which attempts to directly control for unobservable factors, adds information on family background variables such as education or earnings of the parents to the Mincerian earnings equation. Controlling for family background in the OLS estimates may reduce the upward bias in the OLS estimates, but may be unable to eliminate it completely unless the family background variables absorb all unobservable components. The second problem is that returns to education may be heterogeneous, i.e. may vary across individuals, which will usually lead to biased estimates. The third problem are measurement errors in the schooling variable, which is likely to bias the OLS estimator of the returns to schooling downward.

One approach taken to resolve these econometric problems is the use of twin studies. These studies exploit the fact that members of the same family such as siblings or twins are more alike then randomly selected individuals. Twins (or siblings) are less likely to face differences in home environment or financial support and for identical twins even genetic variation in ability may be ruled out. Differences in schooling and earnings of siblings and twins can thus be used to estimate returns to schooling. But even this setting only yields unbiased estimates of returns to schooling if twins face no heterogeneity in any factor that may be correlated with the schooling variable and if the distributions of abilities among twins equals the distribution of abilities in the population as a whole. If these requirements are not satisfied the standard cross-sectional, then the OLS estimator may have a smaller upward bias then the within-twin estimator. Another problem of twin studies is that is that they are likely to exacerbate measurement errors in the schooling variable.

The instrumental variable (IV) approach is a further strategy to solve the bias in the OLS estimates on returns to education. Generally, an IV approach requires an exogenous instrument that is the existence of an observable variable that affects the variable of interest, i.e. years of schooling, but is not correlated with the earnings residual. The IV approach proceeds in two stages: First, the researcher obtains an estimate of the effect of the instrument variables on schooling and then on earnings. Dividing the effect of the instrumental variable on earnings by the effect of the instrumental variables on schooling will yield an unbiased estimator of the returns to schooling if the exogenous instrument only affects earnings through schooling.

Researchers obtained instrumental variables from natural experiments or family background. The identification of natural experiments provides the researcher with variables that are likely to influence the level of schooling but are independent of unobserved individual characteristics. Natural experiments exploit natural variation in the data such as institutional changes, which affect for example, the minimum school leaving age or tuition costs for higher education or factors. Alternatively variables such as geographic proximity to colleges or quarter of birth have been used (Card (1995), Angrist and Krueger (1991)). The instrumental variable estimator has the advantage that if the instrumental variable is not correlated with the measurement error, then measurement errors do not introduce a bias in the IV estimator. However, if returns to education are heterogeneous the existence of an exogenous instrument is not sufficient to guarantee that the IV approach yields unbiased estimates of the average return to schooling. Consider for example a natural experiment that affects mostly a certain subgroup of the sample. Card (1999) shows that in this case the IV estimate may reflect the return of this subgroup and not the average return in the population.

b. Review of the estimates

We will now present a selective review of the evidence on the return to education with a special focus on Europe. We start by providing evidence on instrumental variables based on natural experiments. These findings are presented in Table A1.1 (which is enclosed at the end of the Appendix, along with Table A1.3). Next, we summarize estimates obtained from twin studies and then turn to returns to education whose estimation use information on family background.

Using college proximity as instrumental variable, Card (1999) obtains estimates of returns to schooling for the US, which are nearly twice as large as the corresponding OLS estimates on returns to schooling. He further finds that college proximity affects children of less-educated parents more and therefore interacts college proximity with family background as instrumental variable, adding college proximity as a direct control variable to the earnings equation. Returns to education using this IV procedure are 0.097, which compares to OLS estimates of 0.073. Similarly, Connelly and Uusitalo (1997) find for Finland that IV estimates of the returns to schooling based on college-proximity instrumental variables are 20 to 30% higher than the corresponding OLS estimates.

In Britain the minimum school leaving age was raised from 14 to 15 years in 1947 and from 15 to 16 years in 1973. Harmon and Walker (1995) use these institutional changes as instruments and obtain IV estimates, which are about 2.5 times higher than the corresponding OLS estimates. However, as Card (1999) points out, their IV estimates may be upward biased due to the fact that the effect of the 1947 law change cannot be separated from the changes in educational attainment induced by World War II. Changes in compulsory education are also used as an instrumental variable by Viera (1999) to estimate the return to schooling in Portugal. Again, IV estimates exceed standard OLS estimates. For Italy, Brunello and Miniaci (1999) exploit the fact that in 1969 the possibility to enrol in a college was no longer determined by the curriculum chosen in secondary school. Using family background variables on the parental education level and actual occupation as additional instruments, they obtain IV estimates that exceed OLS estimates by nearly 20%.

Angrist and Krueger (1991) propose individual's quarter of birth as an instrument. They find that individuals who are born earlier in the year reach the minimum school-leaving age at a lower grade than people born later in the year. Hence, individuals who are born earlier in the year and want to drop out legally leave school with less education. Levin and Plug (1999) however find that Dutch individuals born later in the year have significantly lower schooling. They explain this with the fact that within classes older students are likely to receive higher marks, which encourages further schooling. Hence, the net-effect of quarter of birth on schooling attainment is ex ante not clear. An alternative explanation is that in the Netherlands, students are obliged to finish the schooling year they have started, even if they reach the minimum leaving age in the course of this year. Levin and Plug find that returns to education increase by about 10% relative to standard OLS estimates when season of birth is used as an instrument.

Recently, there has been a large increase in estimates of the return to schooling based on twin studies because of the availability of new, relatively large data sets. The data set used by Ashenfelter and Rouse (1998), for example, consists of 340 pairs of identical (monozygotic) twins. It provides the response of each twin about their own and their sibling's schooling level. Hence, the difference in schooling between twins according to one twin can be used as an instrument for the response on the difference in schooling for the other one. As can be seen in Table A1.2, Ashenfelter and Rouse find that within-twins estimates of the returns to education are about 30% lower than the corresponding OLS estimates. Once they control for measurement errors, within-twins estimates increase by about 25%, remaining still below the cross-sectional OLS estimates.

Table A1.2: Estimates based on twin-studies

Author	*Controls*	*Data*	*Specification*	*Cross-section OLS*	*Differences OLS*	*Differences IV*
Isacsson (1999) Sweden	Gender, marital status, quadratic in age and residence in a large city.	Swedish Twin Registry. Administrative and survey measures of schooling	identical twins	0.046 (0.001)	0.022 (0.002)	0.027/0.060 (0.003/0.007)
			fraternal twins	0.047 (0.002)	0.039 (0.002)	0.044/0.060 (0.002/0.003)
			Subsample identical twins	0.049 (0.002)	0.023 (0.004)	0.027 (0.008)
			fraternal twins	0.051 (0.002)	0.040 (0.003)	0.054 (0.006)
Ashen-felter Rouse (1998) US	Gender, race, quadratic in age. Additional controls tenure, marital and union status.	Princeton Twin Survey, 1991-1993. Identical male and female twins.	Without additional controls.	0.110 (0.010)	0.070 (0.019)	0.088 (0.025)
			With additional controls.	0.113 0.010)	0.078 (0.018)	0.100 (0.023)

Isacsson (1999) uses a data set, taken from the population of twins born in Sweden between 1926 and 1958, which consists of 2492 pairs of identical (monozygotic) and 3368 pairs of fraternal (dizygotic) twins. Furthermore information on two measures of schooling (administrative and self-reported level of education) is available for a sub-sample of the data and allows him to correct for measurement error. Isacsson finds that for the sub-sample with both education measures, the within-twins estimate is about 50% (20%) lower than the OLS estimate for identical twins for (fraternal twins). For fraternal twins, the measurement corrected within-twins estimate exceeds the OLS estimate by 35%.

This review suggests that estimates of returns to education obtained from instrumental variable approach or twin studies usually exceed OLS estimates and confirms the results of Ashenfelter, Harmon, Oosterbeek (1999). Analysing estimations from 1974 to 1995 in the US and seven non-US countries (Finland, Honduras, Indonesia, Ireland, Netherlands, Portugal and the United Kingdom), they find that IV estimates and twin studies estimates differ from OLS estimates by 3.1 and 1.6 percentage points.

c. The return to schooling over time and across countries

Many developed countries witnessed major changes in their wage distribution during the last decades. In the US, returns to education decreased during the 1970s and rose sharply during the 1980s. Average returns to education in Europe followed a similar pattern. When looking at European countries one by one, different trends in returns to education can be observed. In this section, we will provide evidence on the evolution of changes in returns to education during the last decades for the US, Europe and selected European countries. We will try to identify the driving forces behind these changes in returns to education and explain why these patterns of change were so different across European countries. Finally, we provide evidence on returns to education for various European countries and discuss the country-specific determinants of these returns.

Returns to education are usually estimated from cross-sectional data and consequently correspond to the wage differential among different skill groups. In a competitive labour market, wages are determined by supply and demand. The supply of skilled workers is determined by the educational attainment of the work force. Demand for skilled workers may change, for example, due to technological change or trade.

It is a well-documented fact that returns to education in the US decreased during the 1970s and increased during the 1980s generating a U-shaped time pattern of educational wage differentials. There seems to be some consensus that these changes in the returns to education may be interpreted as outcomes of shifts in the supply and demand for human capital. The basic idea is that increases in the supply of skilled workers dominated during the 1970s, while demand growth was the driving force in the 1980s. Katz and Murphy (1992), for example, argue that the deceleration in the growth of highly educated labour supply in the United States during the 1980s relative to the 1970s may explain the rise in returns to education during this decade. Katz and Murphy's hypothesis is based on the assumption that the relative demand for skilled workers increased. The prime candidate for explaining the increase in the demand for skilled workers is skill-biased technological change. New technologies have been introduced during the last decades, such as computers or robots, and organizational changes took place within firms that often replaced labour intensive tasks and increased the demand for skilled labour.

For Europe, Harmon, Walker and Westergaard (2001) find that estimates of returns to education where higher in the 1960s as compared to the 1970s. During the 1980s returns dropped even further, but started to rise again in the 1990s. This describes a U-shaped pattern, similar for what we have observed for the US. Performing a meta-analysis of the data, Denny, Harmon and Lydon (2001) confirm these results. A meta-analysis is basically a regression that takes as dependent variables the estimates of different studies that focus on the same topic and similar methodology. The explanatory variables of this regression describe the characteristics of the estimation such as equation specification, sample size and years of estimation. A meta-analysis, thus, controls for the effect of study-specific features on the estimated returns to education. Comparing the change in the US returns to education with Europe, Denny, Harmon and Lydon (2001) show that rates of returns to education in Europe exceeded US returns in the early 1960's. In the course of the 1960s and 1970s, returns to education in Europe and in the US fell at a similar rate. US returns to education reached their minimum at the end of the 1970s, while European returns continued to decline until the mid 1980s. The subsequent increase in returns to education was much more pronounced in the United States. By 1997 returns to schooling in the US were about 3-percentage point higher than in Europe.

Although the pattern of change in the returns to education was strikingly similar in Europe as a whole and the United States, behaviour across European countries differed widely. Returns to schooling in Austria, Switzerland and Sweden decreased, but increased in Denmark, Portugal and Finland. Other countries faced no trend at all or different behaviour of male and female returns to education (Harmon, Walker and Westergaard (2001)). In what follows, we provide some evidence on the evolution of returns to education for selected European countries and try to identify the underlying forces. We start with Great Britain, whose behaviour largely mirrors that of the United States and then turn to Spain and Portugal, which also faced increases in the returns to education during the 1980s. Next, we discuss why wage inequality in France, Germany and Italy remained rather stable. We conclude with Austria, whose returns to education seem to have declined.

Great Britain largely shares the pattern of change in wage inequality with the United States. Similar to the United States, supply of university-educated workers in Great Britain grew rapidly during the 1970s and differentials across skill groups narrowed. During the 1980s, wage inequality and university wage premia increased substantially. Katz, Loveman and Blanchflower (1993) explain with the deceleration in the pace of growth of the relative supply of high-educated workers during the 1980s and the decline in employment in mining, manufacturing, construction and utilities which affected particularly male, manual, low-skilled workers.

Spain underwent profound changes in the 1980s. It joined the European Union and consolidated its democratic institutions. The share of workers in heavy and manufacturing industries declined substantially, whereas the share of employment in commerce, finance and service industries rose. During the 1980s the average educational attainment of the workforce increased remarkably in Spain. In 1981, 74,2% of the employed population had a primary education or less and only 7,6% had completed higher education. By 1991, the population with primary education or less had fallen to 48,5%. The percentage of employed with secondary school degree had more than doubled and 12,5% held a higher education degree. Generation of employment lagged behind the rapid increase of the Spanish labour force. Unemployment was and is still high, affecting particularly younger and less educated people, as well as women. Vila and Mora (1998) find that from 1981 to 1991 Spanish skill wage differentials increased. Returns to lower secondary and primary education decreased, whereas returns to higher education either increased or remained stable.

Portugal faced a severe economic crisis in the first half of the 1980s and major economic changes during the second half. According to Hartog, Pereira and Vieira (2001) returns to education in Portugal remained largely unchanged between 1982 and 1986. Between 1986 and 1992 they increased substantially for both men and women. In contrast to Spain, the increase in returns to education in Portugal was not driven by a reallocation of employment towards skill-intensive sectors. Employment in Portugal rather shifted after 1986 towards sectors that traditionally employ low-educated workers, such as restaurants and hotels, construction, textiles and services. So why did returns to education in Portugal increase? According to Hartog, Pereira and Vieira an increase in demand for high-skilled workers within industries may well explain this finding. Portugal joined the European Union in 1986. It hence embarked in a process of modernising its productive structure, particularly through the introduction of new production technologies. This was made possible thanks to structural funds from the EU and specific financial aids. Furthermore, the liberalisation of trade with more developed countries, may have enhanced the importation of technologies that require skilled labour.

In France, wage inequality did not increase substantially during the 1980s. It declined until 1984 and increased slightly from 1984 to 1987 (Katz, Loveman and Blanchflower (1993)). The fact that significant relative demand shifts did not result in increases in wage differentials through the mid-1980s may be due to French labour market institutions, in particularly the negotiated and the legislated minimum wage. Collective bargaining in France takes mainly place at the industry level and these industry-level arrangements determine minimum wages for each job category. These negotiated minima apply to all firms of all sizes throughout the industry and are binding in case they exceed the legislated minimum wage. The legislated minimum wage applies to all sectors.

Similar to France, former West Germany did not face any increase in wage inequality during the 1980s (Abraham and Houseman (1985), Winkelmann (1994)). One possible explanation is that in contrast to the United States, the growth of the high-educated work force did not decelerate in Germany. Furthermore, as Abraham and Houseman (1985) point out, the high quality of the German apprenticeship system may have prevented returns of education from increasing. They argue that the high level of education for relatively low-skilled workers facilitates the substitutability of workers with different levels of education and experience. Moreover, German solidaristic wage policies, pursued by German trade unions and sought to narrow the gap between highly paid and less highly paid workers, may have tended to depress earnings differentials.

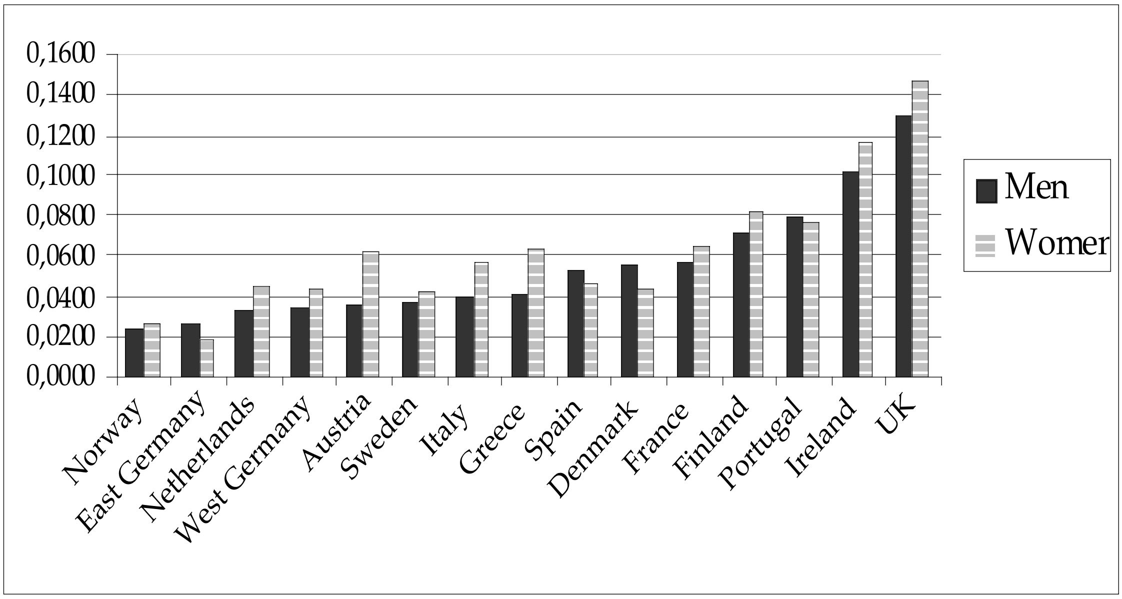
Italy experienced a compression of wage differentials during the 1970s, which according to Ichino and Erikson (1992) came to a halt around 1982-83. The break in the evolution of the wage differentials in 1982-83 coincides with the slow-down in inflation, industrial restructuring, the introduction of an escalator clause in Italian union contracts and the loss of support for unions and their egalitarian pay policies.

In Austria, returns to an additional year of schoolings dropped remarkably from 1981 to 1997. The rapid increase in the labour supply of workers with secondary and tertiary education may explain this fall in Austrian returns to education. Fersterer and Winter-Ebmer (1999) provide evidence for this hypothesis. They find that workers, who belonged to education, age and gender groups with the highest increase in supply, faced the lowest growth of wages. In particular, the drop in returns to education was largest for university students, while returns to a vocational school degree or apprenticeship training remained fairly constant.

Returns to education do not vary only over time, but also across countries. Using a common specification across European countries, Harmon, Walker and Westergaard (2001) find that the Scandinavian countries (Norway, Sweden and Denmark) have the lowest returns to an additional year of schooling. Returns are highest in Ireland and the UK, followed by Germany, Portugal and Switzerland. A meta-analysis reveals an average return to an additional year of schooling of around 6.5% in Europe. It confirms that Scandinavian countries have the lowest returns to schooling, followed by Italy, Greece and the Netherlands, while returns to schooling in UK and Ireland are indeed higher on average.

Similar evidence is provided by, Denny, Harmon and Lydon (2001). They estimate returns to education by OLS using the International Social Survey Program Data 1995. This data set is designed to be consistent across countries. As can be seen in Figure A1.1, they find a large difference in returns to education for men across countries ranging from 2.29% in Norway to 17.66% in North Ireland. Austria, Germany, Netherlands and Norway have relatively low returns to schooling for men, while male returns to schooling are highest in Portugal, Ireland and Great Britain. Returns to education for women exceed male returns in the majority of countries. Female returns to education are lowest in the Netherlands (1.81%), Norway and New Zealand and highest in Great Britain, the Republic of Ireland and North Ireland (16.81%).

Figure A1.1: Returns to schooling in selected EU countries.



- *Source:* Trostel, Walker and Wooley (2001) and Denny, Harmon and Lydon (2001).

d. The role of schooling for male-female wage differentials

In most industrialised countries the gender wage decreased dramatically during the last decade. For a long time a large part of the gender wage had been attributed to differences in schooling between men and women. But differences in average years of schooling among male and female full-time workers have largely disappeared, contributing to a significant decrease in the gender wage, (Blau and Kahn (1997) for the US and Harkness (1996) for the UK). Not only schooling of women, but also female labour force participation and consequently women's accumulated labour force experience has increased. These changes in experience seem to have been even more important in closing the gender wage gap than the increase in years of education (see for example Blau and Kahn (1997) for the US). Relative changes in schooling and work experience altogether, seem to have narrowed the gender wage gap in the US by one-third to one-half between the mid 1970s and late 1980s (O'Neill and Polacheck (1993)). But despite dramatic reductions in the male-female wage gap during the last decades, differences in the earnings of men and women continue to persist.

Differences in the earnings of men and woman can arise for a variety of reason. Differences in schooling, labour force participation rates, work experience, hours worked, job tenure and turnover rates are only the most obvious. Even among equally qualified men and women a substantial gender wage differential remains. Differences in social roles, parental preferences concerning the level of education or job, financial attractiveness of home versus market work, occupational preferences, tastes for jobs or labour market discrimination have been proposed in order to explain these facts. In what follows, we will discuss the different factors that are likely to determine the gender wage gap with a special focus on the role of schooling and "careers".

Differences in schooling between full-time working men and women have largely disappeared in many industrialised countries. Today it is not the amount of schooling, but rather differences in what men and women study, as well as differences in aptitudes and achievement scores across subjects through which schooling affects gender wage gap. The Programme for International Student Assessment (PISA) 2000 of the OECD (OECD 2001a) finds that while males are likely to under-perform in reading, women seem to have a measurable disadvantage in mathematics. Similarly, Brown and Corcoran (1997) conclude for the US that twelfth grade boys score higher on math achievement tests and lower on reading and vocabulary tests. There exists some evidence for the US that these differences in aptitudes may translate into earnings differentials. Altonji (1995) and Brown and Corcoran (1997), for example, find that differences in high school courses play only a modest role in the gender gap among high school students. But differences in the type of college major (e.g.

engineering, physical science, business or law) account for a substantial share of the differential among male and female wages.

A large part of the gender wage gap is generally attributed to the fact that women accumulate a lower amount of experience than men do. Women often interrupt their careers, work fewer hours and have a higher propensity to work part-time than men. As a consequence they may spend a smaller proportion of their working-age time actually working. But it is not only the total amount of experience, which matters. Differences in the timing of and the returns to experience account for a sizeable fraction of the gender wage gap. Light and Ureta (1995) show that about 12 percent of the raw differential in male/female earnings is due to differences in the timing (i.e. differences in the frequency duration and placement of non-work spells) of work experience, while 30% of the gap is due to differences in returns to experience. They further find that career interruptions play a smaller role for women than for men and that women recover more quickly from interrupting their career. This suggests that women may tend to work in occupations that allow them to restore their skills faster, while men may have career interruptions for reasons that are more negatively related to productivity.

Fewer working hours and fewer years in the labour market lead according to the standard human capital theory to less investment in general human capital. Furthermore, women have traditionally higher turnover than men. Expected separation from the current job may discourage investment in employer specific human capital from the part of the women and the employer. A woman may decide not to invest in employer specific human capital, as she knows that she is going to interrupt her career soon. The same holds true for employers. Under imperfect information, the employer may discriminate against the women in terms of training opportunities, if he assumes that the women is likely to interrupt her career in the near future, for example because she is in child-bearing age.

But women do not only have higher turnover rates. It also seems that the reasons underlying the decision to quit a job vary systematically among men and women. Findings by Sicherman (1996) suggest that women take short-run considerations into account when changing jobs, while men place more importance on long run that is career considerations. Sicherman (1996), for example, finds for the US that 12% of women and 4% of men left their job due to a change in residence, which is consistent with the idea that women put more weight on having a job close to their home. Not only proximity to home, but also working hours seem to be a job attribute which is relatively more important to women. Empirical evidence suggests that it is rather hours than wages that play a role in the job choice of women and that job mobility of women is strongly linked to changes in hours (Altonji and Paxson (1991))

Differences in job characteristics such as occupation, industry, unionisation and job-related amenities further contribute sizeably to the male female wage gap. Blau and Kahn (1997) show that the unexplained part of male/female wage regressions reduces from 22% to 13% in 1988 once industry, occupation and collective bargaining variables were included.

Going beyond occupations, there has been a growing amount of research on the impact of part-time and temporary work on wages. Women are heavily over-represented in part-time and temporary jobs. These jobs pay typically less than full-time, permanent jobs. Whether a greater relative fraction of women actually prefers part-time or temporary jobs or whether this behaviour is due to labour market constraints is far from clear. There exists some evidence that a part of the negative effect of part-time work on female wages may be due to selection. Blank (1990) for example finds that controlling for women's selection significantly reduces the negative effect of part-time on women's wages. This is also consistent with the findings of Harkness (1996), who shows that while the wage gap has been closing for full-time working women over the last decades, the relative earnings position of women working part-time has changed little. But while the qualification gap among male and female full-time workers disappeared completely for younger workers, part-time working women continue to be less qualified than full-time working men and women.

Men and women do not earn the same. Differences in the subjects they study and the occupation they choose, as well as differences in work experience play an important role in explaining the gender wage gap. Lifetime work expectations and career considerations are likely to affect female wages considerably. But even among equally qualified men and women the gender differentials persist. Discrimination has been often proposed as a candidate for explaining the unexplained part of the gender wage differential. An alternative explanation is that women have different preferences, which translates into different career choices. The question of how much of the gender wage differential is due to differential choices by women and how much can be ascribed to discriminatory barriers in the labour market is difficult to address. It is even not clear in the first place, if the two theories are not likely to interact rather than to be separated, as past labour market discrimination may have induced women to develop a certain set of preferences, which reflects itself in present choices. The distinction between choice and constraint as determinants of the gender wage gap, thus, remain difficult and controversial.

e. Technological change

An econometrician, who wants to estimate the effect of technological change on human capital and employment, encounters several econometric problems. The first problem is how to measure technology. Often researchers try to address this problem by repeating their estimation with different measures of technologies. This allows them to check that their results are not susceptible to a specific technology measure. The simplest approach of measuring technology is the use of time trends. Unfortunately, time trends are likely to capture much more then solely technological change, such as changes in demand conditions or prices. As a consequence, evidence on the effects of technological change based on this measure should be treated with care.

A prime candidate for the measurement of technology is R&D expenditures. This information is available at the firm level for many countries and different periods. Furthermore, it has the advantage that it is rather comparable across countries and time. The major disadvantage of using R&D is that the data usually refers to the industry in which the innovation originates, not where it is actually used. Moreover, in most European countries, firms are not obliged to disclose the amount of R&D expenditure in their company accounts. While R&D measures the input to knowledge, patents measure its output. Using patents as a measure of technology involves the complication that many of them are of very low value. Taking this into account by weighting the different patents according to their value is not straightforward. Furthermore, the likelihood of an innovation being patented has differed historically across industries. The advantage of patent data is that it measures the direct use of innovations. During the 1980s, the use of computers (and computer-based) resources at the workplace has grown enormously. As a consequence, computers maybe the most concrete example of technological change during 1980s and a good proxy for the rate of technological change at the work place. The advantage of using a technology measure based on investment in computers lies in the fact that it captures the use and not the production of an innovation.

The past decades witnessed major changes in technologies, such as the rapid spread of computers at our workplaces, the expansion of computer-assisted production techniques and robots and the more intensive use of the Internet. How do these changes affect the relative demand for skilled workers? There exist basically two hypotheses, which try to explain the relation between the relative demand for skilled workers and technological change. The first hypothesis relates technological change and demand of skilled workers. If highly educated workers have a relative advantage in adjusting to and implementing of new technologies, then the spread of these new technologies is likely to increase the demand for skilled workers relative to unskilled workers. This means that in a period of technological change, the productivity of highly educated workers increases relative to less educated workers, due to the fact the highly educated workers are more able to adjust to a changing environment. Hence, times of rapid technological change should also be associated with an increased demand for skill and may lead to higher returns to education. Furthermore, industries characterized by high rates of innovations should have a higher demand for highly educated workers. Bartel and Lichtenberg (1987) shed some light on this issue. They find that the relative demand for educated workers declines as the age of the plant and particularly of equipment increases. This is especially the case in R&D intensive industries.

The second hypothesis claims the technological change is skill biased. New technologies mainly replace labour intensive tasks and are likely to complement skilled workers. Thus, the transition from an old to a new technology results in permanent changes in the equilibrium share of skilled labour, holding output and relative prices constant. If the demand for skilled workers outstrips its supply then returns to education increase. There exists clear evidence at the industry level that almost all industries started to employ educated workers during the 1970s and 1980s and that industries that were more computerized increased their demand for college-educated workers at a faster rate (e.g. Berman, Bound and Griliches (1994), Autor, Katz and Krueger (1996), and Machin and Van Reenen (1998)). Machin and Van Reenen (1998) use data of the manufacturing sector in Denmark, France, Germany, Japan, Sweden and the UK. Using R&D intensity as a measure of technology, they provide evidence for skill bias across all these countries. Berman, Bound and Griliches (1994) and Autor, Katz and Krueger (1996) model changes in workforce skill as a function of changes in industry capital intensity and industry-level investment in computer equipment. Their findings reveal a strong positive correlation between the level of computer investment and changes in the skill of workers in the industry.

A positive correlation between the level of computer investment and demand for skilled workers does not necessarily mean that computer investment causes an increase in the demand for skills, since industries that are highly computerized may demand more skilled workers for other reasons as well. Using plant-level data, Doms, Dunne and Troske (1997) also come to the conclusion that a higher proportion of college-educated workers are employed in technologically advanced plants, when using cross-sectional data. The longitudinal analysis however reveals that plants that adopt new factory automation technologies have a higher proportion of skilled workers before and after the adoption of the new technologies. They conclude that the correlation between skill upgrading and the adoption of new technologies is largely due to the fact that plants with a high wage workforce are more likely to adopt new technologies. The authors however emphasise that the type of technology they use is directly used in the production of manufactured goods, whereas computer investment is a main tool for white-collar workers. When they use computer investment as an alternative measure of technology, they find a positive correlation with the growth of skilled worker even in the longitudinal data. This leads them to conclude that the effect of new technologies on the structure of the work force depends critically on which type of technology is adopted.

Industry level studies may be subject to serious aggregation bias. Hence, the fact that even firm-level studies seem to support the existence of skill-biased technological chance considerably strengthens the evidence. Using a data set on British plant, Haskel and Heden (1997) find evidence, that computers positively affect the growth of skill intensity. Dueguet and Greenan (1997) use an innovation survey for a panel of French manufacturing firms from 1986 to 1991 and come to a similar conclusion. They argue that skill bias arises mainly from the introduction of new products. For Spain, Aguirrebriria and Alonso-Borrega (1997) find that the introduction of technological capital, defined as "successful innovations generated externally to the firm" has a strong negative effect on blue-collar workers. However, they find no robust effect of R&D.

While there exists evidence, that technological change affects the relative demand for skilled workers, only a few studies try to understand the mechanisms through which technological change operates. Unobserved factors play a large role in the analysis of technological change and demand for skills. Some conjecture (Dunne, Haltiwanger and Troske, 1996 and Machin and van Reenen (1998)) that organizational change might well be one of them. In most industrialized countries, there has been a trend towards less hierarchy and more flexible organizational forms. More autonomy and responsibility is given to the workers and often they are performing a wider range of tasks. Caroli and van Reenen (1999) go even further and claim that lack of necessary organizational structures that facilitated the introduction of new technologies, may to some extend explain the so-called "productivity paradox." This states that huge investments in computers often fail to result in significant increases in productivity. Caroli and van Reenen (1999) use a panel of British and French establishments in order to investigate whether organizational changes such as the decentralization of authority, delayering of managerial functions and increased multi-tasking affects the skill composition of firms. They find that organizational changes tend to reduce the demand for unskilled workers and lead to greater increases in productivity in establishments with larger initial skill endowments. They conclude that the widespread introduction of new organizational forms may be important in explaining the declining demand for less skilled workers.

In contrast to the studies on technology and demand for skilled workers, most of the studies that try to analyse the relation between technological change and wages use individual data. The rise in wage inequality in the US and the widespread notion that technological change may be the driving force behind it triggered a large amount of studies on this subject.

Mincer (1993), using data from 1963 to 1987, shows that relative earnings of college graduates increased with R&D intensity. Similarly, Allen (1998) provides evidence that changes in innovative activities as measured by R&D intensity and the usage of high-tech capital play an important role in explaining changes in the wage structure. He finds that increases in the return to schooling between 1979 and 1989 were most pronounced in industries with a greater R&D intensity and more high-tech capital. Krueger (1993) argues that computers change the structure of wages and shows that workers that use computers are paid more. The effect of computer use on wages is greater for educated workers.

A positive relation between returns to education and computer use (or other technologies) does not necessarily mean that it is technology that drives wages of skilled workers up. There exists a large body evidence demonstrating that workers with the highest ability and hence the highest wages are given the best technologies to use. This means that it is rather selection and not the increase in productivity that explains the computer wage premium. DiNardo and Pischke (1997), for example, find a positive correlation between wages and computer use in German data, which is similar to the finding of Krueger (1993). However, they show that the correlation between wages and pencil use is equally robust, which is a point in favour of the selection hypothesis. Likewise, Entorf and Kramarz (1997) emphasise that in France the cross-sectional association between wages and computer use disappears once they control for unobserved individual characteristics. One should thus be careful in interpreting the computer-wage correlation as the causal effect of technical change on wages. Several studies that use firm level data find a strong positive relation between technology and inter-industry wages. It is not clear, however, whether this effect arises because of sorting. Bartel and Sicherman (1999) address this issue by using individual-level data in order to explain differences in inter-industry wages. They conclude that sorting is the dominant explanation for higher wages in industries that are subject to faster technological change. Similarly, Doms, Dunne and Troske (1997) find that the positive effect on wages disappears once they account for individual fixed effects. Chennels and Van Reenen (1997) show that the effect of technology on wages disappears once they use industry level measures of technological opportunity as an instrument for the adoption of new micro electronic technologies at the plant level.

To understand the effect of technological change on employment, assume that a firm decides to implement a computer-assisted production process. The implementation of this new process allows the firm to save on labour, which means that it can produce the same amount of output as before with a lower level of employment. This initial drop in employment is accompanied by a cost reduction, which may reflect itself in a decrease in prices. The latter may translate into an increase in the output of the firm, inducing employment to increase. Whether employment is higher before or after the adoption of the new technology depends on a variety of factors. The positive employment effect is smaller if the firm has some degree of market power and passes only part of the cost reduction on in form of lower prices. Economies of scale may magnify the positive employment effect. If consumers react strongly to changes in prices then the positive employment effect is likely to be large. And product innovations as opposed to process innovation have generally a stronger output expansion effect and hence are more likely to affect employment positively. Summarising, the effect of technological change on the level of employment is a priori far from clear.

Cross-industry studies on the relationship between employment and technology have been relatively scarce. Analysing the OECD STAN/ANBERD database on manufacturing, Blechinger, Kleinknecht, Licht and Pfeiffer (1998) show that industries with higher R&D intensity expanded more quickly. Firm level studies provide a wide variety of results from different countries. It appears that product innovation has a positive effect on employment growth (e.g. Entorf and Pohlmeier (1990) for German firms.) Evidence concerning process innovations is rather mixed. Some studies find positive effects (e.g. Blanchflower and Burgess (1998) for the UK and Australia and Blechinger et al (1998) for Dutch firms). Greenan and Guellac (2000) conclude for France that process innovations have a strong positive effect at the firm level, but that this effect disappears at the industry level. They find the opposite for product innovations. Note, however, that firms may introduce new technology when they expect demand conditions to improve, which may lead to an upward bias in the coefficient on the measure of technology. Finally, Entorf, Gollac and Kramarz (1999) show that computer users are protected from job losses in the short run, that is, as long as bad business conditions do not last too long. But also here the question arises whether it is not rather selection, which determines not only wage gains but also job losses.

This review of recent studies on technology, demand for education, wages and employment reveals a strong positive correlation between technology and the relative demand for skilled workers for different time periods and across countries. This finding seems to be robust and suggests that technology is on average skill biased. There also exists some evidence of a positive correlation between wages and large technological innovations. However, measures based on the diffusion of technology, such as computer use provide no evidence on the existence of a causal effect of technology on wages. Similarly, the positive relation between inter-industry wages and technology seems to be largely due to sorting. Evidence on total employment is mixed. Product (process) innovation seems to be positively (negatively) associated with employment. Hence, the only definite conclusion we are able to draw is that the recent technological change was on average skill-biased.

2. Cross-country data on human capital[1]

Most governments gather information on a number of educational indicators through population censuses, labour force surveys and specialized surveys. Various international organizations collect these data and compile comparative statistics that provide easily accessible and (supposedly) homogeneous information for a large number of countries. Perhaps the most comprehensive regular source of international educational statistics is UNESCO's *Statistical Yearbook*. This publication provides reasonably complete yearly time series on school enrollment rates by level of education for most countries in the world and contains some data on the educational attainment of the adult population, government expenditures on education, teacher/pupil ratios and other variables of interest. Other UNESCO publications contain additional information on educational stocks and flows and some convenient compilations. Other useful sources include the UN's *Demographic Yearbook*, which also reports educational attainment levels by age group and the IMF's *Government Finance Statistics*, which provides data on public expenditures on education. Finally, the OECD also compiles educational statistics both for its member states (e.g. OECD (2000)) and occasionally for larger groups of countries.

a. Data on schooling

The UNESCO enrollment series have been used in a large number of empirical studies of the link between education and productivity. In many cases this choice reflects the easy availability and broad coverage of these data rather than their theoretical suitability for the purpose of the study. Enrollment rates can probably be considered an acceptable, although imperfect, proxy for the flow of educational investment. On the other hand, these variables are not necessarily good indicators of the existing stock of human capital since average educational attainment (which is often the more interesting variable from a theoretical point of view) responds to investment flows only gradually and with a very considerable lag.

In an attempt to remedy these shortcomings, a number of researchers have constructed data sets that attempt to measure directly the educational stock embodied in the population or labour force of large samples of countries. One of the earliest attempts in this direction is due to Psacharopoulos and Arriagada (P&A, 1986) who, drawing on earlier work by Kaneko (1986), report data on the educational composition of the labour force in 99 countries and

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[1] This section is partly based on de la Fuente and Doménech (2000).

provide estimates of the average years of schooling. In most cases, however, P&A provide only one observation per country.

More recently, there have been various attempts to construct more complete data sets on educational attainment that provide broader temporal coverage and can therefore be used in growth accounting and other empirical exercises. The existing data sets on educational attainment have been constructed by combining the available data on attainment levels with the UNESCO enrollment figures to obtain series of average years of schooling and the educational composition of the population or labour force. Enrollment data are transformed into attainment figures through a perpetual inventory method or some short-cut procedure that attempts to approximate it.

Most of the studies in the macroeconomic literature we review in this report rely on one of the following data bases:

• **Kyriacou (1991)** provides estimates of the average years of schooling of the labour force *(YS)* for a sample of 111 countries. His data cover the period 1965-1985 at five-year intervals. He uses UNESCO data and P&A's attainment figures to estimate an equation linking *YS* to lagged enrollment rates. This equation is then used to construct an estimate of *YS* for other years and countries.

• **Lau, Jamison and Louat (1991) and Lau, Bhalla and Louat (1991)**. These studies use a perpetual inventory method and annual data on enrollment rates to construct estimates of attainment levels for the working-age population. Their perpetual inventory method uses age-specific survival rates constructed for representative countries in each region but does not seem to correct enrollment rates for dropouts or repeaters. "Early" school enrollment rates are estimates constructed through backward extrapolation of post-1960 figures. They do not use or benchmark against available census figures.

• **Barro and Lee (B&L 1993)** construct education indicators combining census data and enrollment rates. To estimate attainment levels in years for which census data are not available, they use a combination of interpolation between available census observations (where possible) and a perpetual inventory method that can be used to estimate changes from nearby (either forward or backward) benchmark observations. Their version of the perpetual inventory method makes use of data on gross enrollments[2] and the age composition of the

[2] The gross enrollment rate is defined as the ratio between the total number of students enrolled in a given educational level and the size of the population which, according to its age, "should" be enrolled in the course. The net enrollment rate is defined in an analogous manner but counting only those students who belong to the relevant age group. Hence, older students (typically repeaters) are excluded in this second case.

population (to estimate survival rates). The data set contains observations for 129 countries and covers the period 1960-85 at five-year intervals. Besides the average years of education of the population over 25, Barro and Lee report information on the fraction of the (male and female) population that has reached and completed each educational level. In a more recent paper (B&L, 1996), the same authors present an update of their previous work. The revised database, which is constructed following the same procedure as the previous one (except for the use of net rather than gross enrollment rates), extends the attainment series up to 1990, provides data for the population over 15 years of age and incorporates some new information on quality indicators such as the pupil/teacher ratio, public educational expenditures per student and the length of the school year. Some further extensions, refinements and updates of this data base have been made available by the authors in recent years and are discussed in Barro and Lee (2000) and Lee and Barro (2001).

• **Nehru, Swanson and Dubey (NSD 1995)** follow roughly the same procedure as Lau, Jamison and Louat (1991) but introduce several improvements. The first one is that Nehru et al collect a fair amount of enrollment data prior to 1960 and do not therefore need to rely as much on the backward extrapolation of enrollment rates. Secondly, they make some adjustment for grade repetition and drop-outs using the limited information available on these variables.

We can divide these studies into two groups according to whether they make use of both census attainment data and enrollment series or only the latter. The first set of papers (Kyriacou and Barro and Lee) relies on census figures where available and then uses enrollment data to fill in the missing values. Kyriacou uses a simple regression of educational stocks on lagged flows to estimate the unavailable levels of schooling. This procedure is valid only when the relationship between these two variables is stable over time and across countries, which seems unlikely although it may not be a bad rough approximation, particularly within groups of countries with similar population age structures. In principle, Barro and Lee's procedure should be superior to Kyriacou's because it makes use of more information and does not rely on such strong implicit assumptions. In addition, these authors also choose their method for filling in missing observations on the basis of an accuracy test based on a sample of 30 countries for which relatively complete census data are available.

The second group of papers (Louat et al and Nehru et al) uses only enrollment data to construct time series of educational attainment. The version of the perpetual inventory method used in these studies is a bit more sophisticated than the one in the first version of

Barro and Lee, particularly in the case of Nehru et al.[3] On the other hand, these studies completely ignore census data on attainment levels. To justify this decision, Nehru et al observe that census publications typically do not report the actual years of schooling of individuals (only whether or not they have completed a certain level of education and/or whether they have started it) and often provide information only for the population aged 25 and over. As a result, there will be some arbitrariness in estimates of average years of schooling based on this data and the omission of the younger segments of the population may bias the results, particularly in LDCs, where this age group is typically very large and much more educated than older cohorts. While this is certainly true and may call for some adjustment of the census figures on the basis of other sources, in our opinion it hardly justifies discarding the only direct information available on the variables of interest.

Methodological differences across different studies would be of relatively little concern if they all gave us a consistent and reasonable picture of educational attainment levels across countries and of their evolution over time. Unfortunately, this is not the case. Different sources show very significant discrepancies in terms of the relative positions of many countries and practically all of them display implausible estimates or time profiles for at least some countries. Although the various studies generally coincide when comparisons are made across broad regions (e.g. the OECD vs. LDCs in various geographical areas), the discrepancies are very important when we focus on the group of industrialized countries. Another cause for concern is that existing estimates often display extremely large changes in attainment levels over periods as short as five years (particularly at the secondary and tertiary levels).

To a large extent, these problems have their origin in the deficiencies of the underlying primary data. As Behraman and Rosenzweig (1994) have noted, there are good reasons to worry about the accuracy and consistency of UNESCO's data on both attainment levels and enrollment rates. De la Fuente and Doménech (2000), after reviewing the available data for OECD countries, argue that the problems noted above can be traced back to shortcomings of the primary statistics, which do not seem to be consistent, across countries or over time, in their treatment of vocational and technical training and other courses of study,[4] and reflect at times

[3] Differences across these studies have to do with the correction of enrollment rates for dropouts and repeaters and with the estimation of survival probabilities. Latter versions of Barro and Lee have improved the treatment of these issues.

[4] Steedman (1996) documents the existence of important inconsistencies in the way educational data are collected in different countries and argues that this problem can significantly distort the measurement of educational levels. She notes, for example, that countries differ in the extent to which they report qualifications not issued directly (or at least recognized) by the state and that practices differ as to the classification of courses which may be considered borderline between different ISCED levels. The stringency of the requirements for the granting of various completion degrees also seems to vary significantly across countries.

the number of people who have started a certain level of education and, at others, those who have completed it. They conclude that --despite the fact that the contributions they review represent a significant advance in this area-- the available data on human capital stocks are still of dubious quality.

Concerns about poor data quality and its implications for empirical estimates of the growth effects of human capital have motivated some recent studies that attempt to improve the signal to noise ratio in the schooling series by exploiting additional sources of information and introducing various corrections. De la Fuente and Doménech (D&D 2000) restrict their work to a sample of 21 OECD countries for which they construct new educational attainment series covering the period 1960-90 at quinquennial intervals. They focus on cleaning up the available census and survey data rather than on perfecting the fill-in procedure. After collecting all the information they could find on educational attainment in OECD countries, both from international publications and from national sources, they use a heuristic approach to try to reconstruct a plausible time profile of attainment in each country, eliminating sharp breaks in the series that can only arise from changes in data collection criteria. Their approach involves using judgment to choose among alternative census or survey estimates when several are available and, at times, requires reinterpreting some of the data from international compilations as referring to somewhat broader or narrower schooling categories than the reported one. Missing data points lying between available census observations are filled in by simple linear interpolation. Missing observations prior to the first census observation are estimated, whenever possible, by backward extrapolations that make use of census information on attainment levels by age group. A revised version of this data set (D&D 2001) also incorporates information provided by national statistical offices in response to a request for assistance channeled through the OECD.

Cohen and Soto (2001) follow a roughly similar approach to construct a schooling data set for a much larger sample of 95 countries at 10 year intervals covering the period 1960-2000. They collect census and survey data from UNESCO, the OECD's in-house educational data base, and the websites of national statistical agencies, and exploit to the extent possible the available information on attainment levels by age group to fill in missing cells through forward and backward extrapolations. Remaining gaps in the data are filled using enrollment rates from UNESCO and other sources.

Estimates of reliability ratios for different data sets

Tables A2.1 and A2.2 report estimates of reliability ratios for some of the data sets we have discussed in the previous section. Following the methodology proposed by Krueger and

Lindhal (2001) and briefly reviewed in section 3b.iii of the main report, we estimate the reliability ratio r_k of a given series of average years of schooling (say YS_k) by using YS_k to try to explain alternative estimates of the same variable (YS_j with $j \neq k$). Hence, the figure reported in the tables below for data set k is the average value of the slope coefficient in a series of regressions of the form

$$YS_j = c + r_{kj} YS_k$$

where j denotes the "reference" data set and varies over the last available version of all data sets different from k. The reliability ratio of Barro and Lee's (2000) data set, for instance, is estimated by including these authors' estimate of average years of schooling as explanatory variable in a set of regressions where the reference (dependent) variables are the average years of schooling estimated by Kyriacou (1991), NSD (1995) and Cohen and Soto (2001). Other versions of the Barro and Lee data set, however, are not used as a reference, because the correlation of measurement errors across the same family of estimates is almost certainly very high and this would artificially inflate the estimated reliability ratio.

The exercise we have just described is repeated for several transformations of average years of schooling and for two different samples (OECD and all available countries, including OECD). In particular, we estimate reliability ratios for years of schooling measured in levels (YS_{it}) and in logs (ys_{it}), for average annual changes in both levels and logs measured across successive (quinquennial or decennial) observations (ΔYS_{it} and Δys_{it}), for log years of schooling measured in deviations from their country means (ys_{it} - ys_i) and for average annual log changes computed over the period 1965-85[5] (Δys_i). Notice that Δys_{it} corresponds to annual growth rates and ys_{it} - ys_i is the "within" transformation often used to remove fixed effects. The last row of each table shows average values of the reliability ratio for each type of data transformation (taken across different data sets), and the last column displays the average reliability ratio of each data set (taken over different data transformations). In each table the different data sets are arranged by decreasing average realiability ratios.

A comparison of Tables A2.1 and A2.2 shows that the estimated reliability ratios are lower for the OECD than for the full sample (of up to 110 countries). This is likely to be misleading. The number of available primary sources that can be drawn upon to construct estimates of educational attainment is probably higher in developed than in underdeveloped countries. As a result, the variation across data sets is likely to be smaller in LDCs, and this will tend to raise the estimated reliability ratio. To a large extent, however, the larger ratios obtained for the full sample will simply reflect a higher correlation of errors across data sets (i.e. an upward bias in the estimated reliability ratio). Hence, the results in Table

[5] This is the longest period over which all the available schooling series overlap.

A2.1 are probably a better measure of the amount of measurement error in existing schooling data sets.

Table A2.1: Average reliability ratios, OECD 21 subsample

	YS_{it}	ys_{it}	ΔYS_{it}	Δys_{it}	ys_{it}-ys_i	Δys_i	*average*
D&D (2001)	0.623	0.716	0.376	0.736	0.894	0.898	*0.707*
C&S (2001)	0.619	0.709	0.203	0.595	0.776	0.796	*0.616*
D&D (2000)	0.638	0.727	0.058	0.457	0.873	0.642	*0.566*
B&L (2000)	0.646	0.595	0.027	0.08	0.679	0.502	*0.421*
Kyr. (1991)	0.743	0.831	0.020	0.066	0.446	0.243	*0.391*
NSD (1995)	0.301	0.528	0.059	0.224	0.858	0.277	*0.375*
B&L (1996)	0.558	0.488	0.026	0.052	0.628	0.357	*0.351*
B&L (1993)	0.530	0.428	0.018	0.014	0.403	0.318	*0.285*
average	*0.582*	*0.628*	*0.098*	*0.278*	*0.695*	*0.504*	*0.464*

- *Note:* This subsample is comprised of the 21 OECD countries for which de la Fuente and Doménech have compiled data.

Table A2.2: Average reliability ratios, all available countries

	YS_{it}	ys_{it}	ΔYS_{it}	Δys_{it}	ys_{it}-ys_i	Δys_i	*average*
C&S (2001)	0.788	0.919	0.396	0.848	0.958	0.950	*0.810*
NSD (1995)	0.877	0.920	0.296	0.634	0.834	0.668	*0.705*
Kyr. (1991)	0.981	1.000	0.092	0.436	0.754	0.693	*0.659*
B&L (2000)	0.910	0.781	0.145	0.299	0.823	0.752	*0.618*
B&L (1996)	0.900	0.777	0.117	0.259	0.812	0.709	*0.596*
B&L (1993)	0.897	0.788	0.129	0.256	0.704	0.563	*0.556*
average	*0.892*	*0.864*	*0.196*	*0.455*	*0.814*	*0.723*	*0.657*

Notes:
- The regressions used to estimate the reliability ratios are estimated using all the common observations for each pair of data sets over a sample of 110 countries for which at least two independent estimates are available.
- Data are reported at five-year intervals except by Cohen and Soto who do it at 10-year intervals. To compute reliability ratios for ΔYS_{it} and Δy_{it} in the case of Cohen and Soto, we attribute the observed annualized change or growth rate in H over the entire decade to both of its quinquenni.a.
- While the true reliability ratio must lie between zero and one, a few of the coefficients of the pairwise regressions are either negative or greater than one. To compute the averages reported in the table, I ignore these values, i.e. assign a value of zero to negative estimates and a value of one to estimates greater than this number.
- The version we use of Barro and Lee (1993) is actually taken from Barro and Lee (1994b). We do not know if the two data sets are identical or if there are minor differences between them.

The overall average value of the reliability ratio in the OECD subsample is 0.464. This suggests that the estimated coefficient of schooling in growth equations is likely to suffer from a substantial downward bias, even without taking into account the further loss of signal that arises when additional regressors are included in these equations. The bias will tend to be smaller for estimates obtained using the data in levels or logs, even when fixed effects are included, but is likely to be extremely large in specifications that use data differenced over relatively short periods. The average reliability ratio is only 0.278 for the data in quinquennial log differences, and 0.098 for level differences taken at the same frequency.

Our results also indicate that the importance of measurement error varies significantly across data sets, although their precise ranking depends on the data transformation that is chosen. Two of the datasets most widely used in cross-country empirical work, those by Kyriacou (1991) and Barro and Lee (various years), perform relatively well when the data is used in levels, but contain very little signal when the data is differenced. Recent efforts to increase the signal to noise ratio by de la Fuente and Doménech (2001) and Cohen and Soto (2001) seem to have been at least partially successful, but even in these cases the potential estimation bias remains large.

b. Direct measures of skills and achievement

It is clear that average years of schooling can be at best an imperfect proxy for the average stock of human capital of the population. The level of skill will vary across countries with similar levels of school attainment if there are differences among them in the quality of their educational systems or in the extent to which skills are built up or maintained through other channels, such as various types of post-school training and on-the-job learning.

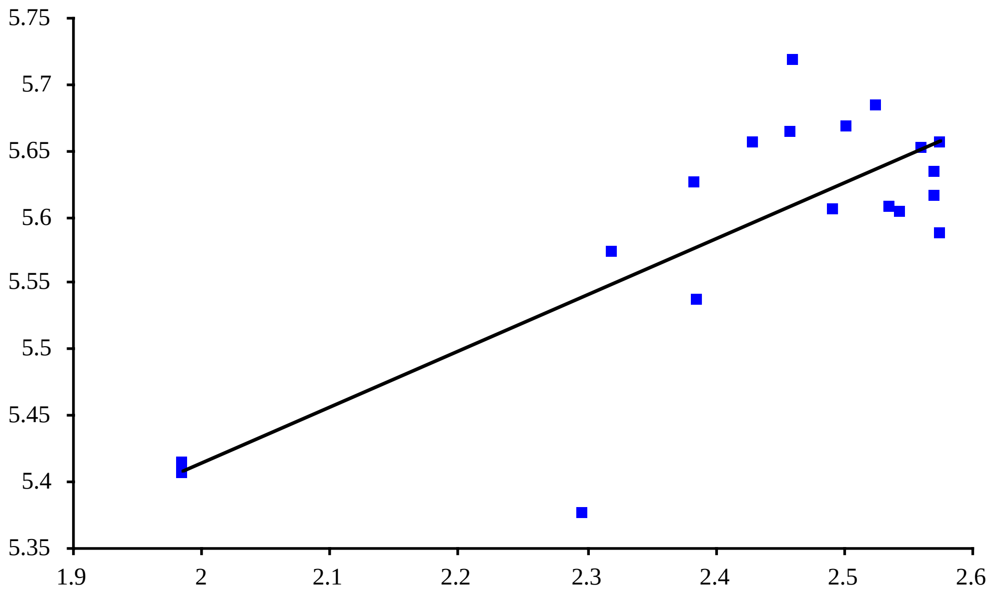
Table A2.3: International achievement and literacy tests

years of data collection	*conducted by*	*subjects*	*no. of countries*	*population tested*
1964	IEA	mathematics	13	13, final sec.
1970-72	IEA	science	19	10, 14, final sec.
		reading	15	10, 14, final sec.
1982-83	IEA	mathematics	20	13, final sec.
1984	IEA	science	24	10, 14, final sec.
1988	IAEP	mathematics	6	13
		science	6	13
1991	IEA	reading	31	9, 14
1990-91	IAEP	mathematics	20	9, 13
		science	20	9, 13
1993-98	IEA (TIMSS)	math	41	9, 13, final sec.
		science	41	9, 13, final sec.
1994-1998	OECD (IALS)	reading and quantitative literacy	23	16-65
2000	OECD (PISA)	reading, mathematical and scientific literacy	32	15

- *Source:* updated from Lee and Barro (2001)
- *Key:* IEA = International Association for the Evaluation of Educational Achievement; IAEP = International Assessment of Educational Progress; PISA = Programme for International Student Assessment; final sec. = final year of secondary schooling.

While the available information is much scarcer than for formal school attainment, student scores in standardized international achivement tests and some recent literacy studies provide some data on the quality of educational outputs and on the skill level of the population that can be a useful complement to the schooling data reviewed above. Table A2.3 summarizes the standardized tests in mathematics, science and reading that have been administered by different international organizations at various times, as well as two recent literacy studies sponsored by the OECD (known as PISA and IALS). This last group of studies is of particular interest because they specifically attempt to measure the extent to which respondents have developed basic skills that will be essential both at work and in everyday life rather than their mastery of a standard curriculum. These skills include the ability to understand and use information and apply simple mathematical techniques and basic scientific knowledge to the solution of practical problems.

Figure A2.1: Literacy skills vs. average schooling



- *Note:* Log IALS average score vs. log average years of schooling in 2000 from Cohen and Soto (2001).

The IALS (International Adult Literacy Survey) study is the only available one that focuses on the entire population of working age rather than on young subjects currently enrolled in school. Hence, it probably provides the best available data for testing the hypothesis, implicit in most of the empirical work we will survey below, that educational attainment can be used as a proxy for the stock of human capital because many of the skills that are relevant in production are probably acquired in school. Figure A2.1 shows the relationship between average national literacy scores (after averaging over the three types of skills measured in the study) and average years of schooling in 2000 (taken from Cohen and Soto (2001)), with both variables measured in logs. When we use average schooling to try to explain literacy scores in a simple regression, the estimated slope coefficient is 0.423, with a *t* ratio of 3.83, and the R^2 of the regression is 0.494. Hence, educational attainment alone explains half of the observed variation in literacy scores, suggesting that average schooling is indeed a useful proxy for skills, but a far from perfect one. If we interpret the coefficient of this regression as a reliability ratio (with schooling as a noisy measure of skill levels), its estimated value reinforces the standard concern in the literature that measurement error will lead to the underestimation of the impact of human capital on productivity. Or, to put it differently, it seems likely that we can learn something useful by examining the correlation between school attainment and growth, but it is almost certain that in order to get a better picture of the importance of human capital we need to find ways to control also for the quality of education and for other ways in which skills can be acquired.

3. Results of macroeconomic studies on human capital and growth

This section reviews the main empirical studies that have attempted to measure the contribution of human capital accumulation to economic growth.[6] We will organize the discussion of the bulk of the literature around groups of studies defined in terms of their econometric specification, distinguishing between papers that estimate production function-based specifications and those based on convergence equations and, within the latter group, betwen those based on ad-hoc specifications and those that have estimated structural equations along the lines of Mankiw, Romer and Weil (1992). We will consider in a separate section some recent studies that have focused on data quality and measurement error. As noted in the text, practically all of these studies use some schooling indicator (either enrollment rates or average years of schooling) as a proxy for human capital. There are also a small number of studies that explore the growth effects of more direct measures of educational or labour force quality based on internationally comparable achievement tests. These will be discussed separately.

For easy reference, Box A3.1 summarizes the notation used in the numerous tables that appear below. Wherever possible, we report the estimated values of structural parameters (i.e. the coefficients of the production and technical progress functions) which are denoted by the same symbols as in Box 2 in the text. Otherwise, the tables show the relevant regressor. Standard explanatory variables include the rates of investment in physical and human capital (s_k and s_h), initial income per capita or per worker *(Q)* and years of schooling *(YS)*. In

[6] This section is based on de la Fuente (2002).

the last two cases, lower-case letters are used to denote logarithms and the symbol Δ to denote annual changes.

Box A3.1: Notation used in the Tables

As noted in the text, most structural analyses of the determinants of economic growth are based on a Cobb-Douglas aggregate production function of the form

$$(1)\quad Y_{it} = A_{it}\, K_{it}^{\alpha_k} H_{it}^{\alpha_h} L_{it}^{\alpha_l}$$

where Y_{it} denotes the aggregate output of country i at time t, L_{it} is the level of employment, K_{it} the stock of physical capital, H_{it} the average stock of human capital per worker, and A_{it} an index of technical efficiency or total factor productivity (TFP). In most applications, H is typically replaced by years of schooling (YS) or by an exponential function of it, $H = exp(\theta YS)$.

The tables shown below respect this notation for the different inputs of the production function and for the relevant output elasticities, α_i (with $i = k, h, l$ or ys). On occasion, the production function also includes as an argument the stock of R&D capital, whose elasticity will be denoted by $\alpha_{R\&D}$. The symbol ρ will denote the coefficient of YS measured in levels in an otherwise standard Cobb-Douglas production function (this is the Mincerian specification discussed in Box 2 in the text), or the result of dividing α_{YS} by average YS in the sample when a standard Cobb-Douglas is estimated with YS in logs. In both cases, this parameter measures the % increase in output that would follow from a unit increase in YS and α_{YS} the elasticity of output with respect to years of schooling.

As in Box 2, lower case letters will be used for factor stocks measured in logarithms, and the symbol Δ will indicate the average annual change in the relevant variable. Hence, YS is years of schooling in levels, ys the same variable in logarithms, ΔYS the average annual increase in years of schooling over the relevant period, and Δys the average annual increase in the logarithm of the same variable, which is approximately equal to the annual percentage change in the original variable measured in levels. Similarly, Δa will stand for the rate of technical progress.

We will use $Q = Y/L$ to stand for output per capita or per worker. The symbol s_i will denote the fraction of GDP invested in type-i capital or, in the case for human capital, some proxy for this variable typically based on school enrollments. The symbol γ_j will be used for the coefficients of the technical progress function, except for the rate of technological diffusion, which will be denoted by λ, as in

$$(2)\quad \Delta a_{it} = \gamma_{io} + \lambda b_{it} + \gamma_{YS}\, YS_{it} + \gamma_{bh}\, YS_{it} b_{it} + \gamma_r\, R\&D_{it}$$

where b stands for the gap with the world technological frontier. The parameter β will be interpreted as the rate of convergence and is typically the coefficient of initial income per capita in a convergence equation.

a. Ad-hoc growth equations

A simple way to explore the connection between human capital and growth is to introduce some indicator of human capital in a convergence equation in which the growth rate of real output over a given period is explained in terms of the initial level of income per capita and other variables motivated by informal theoretical considerations. This approach has been followed with generally encouraging results in a large number of papers in the literature using (mostly cross-section) data for the post-WWII period.

The results of some of the earlier studies in the literature are summarized in Table A3.1. The explanatory variables used in the regressions include the initial level of per capita income (Q_0), different indices of human capital at the beginning of the period (H_0), and the rates of investment (s_k) and population (or labour force) growth (n). Landau (1983 and 1986), Baumol et al (1989) and Barro (1991) find that the coefficient of initial human capital is positive and highly significant. Baumol et al observe that the inclusion of a proxy for education is enough to "set things right" in a convergence equation in which, when the only explanatory variable is initial income, the neoclassical prediction that poorer countries tend to grow faster than rich ones seems to fail.

Barro (1991) estimates two different versions of the convergence equation in a first attempt to identify the channels through which education affects growth. In the first one he does not control for fertility or the investment rate, while in the second equation he includes both of these variables. As can be seen in the table, the human capital indicators lose part of their significance and have smaller coefficients in the second equation. This suggests that an important part of the effect of education on growth is channeled through a reduction in the fertility rate (education increases the opportunity cost of female time) and an increase in the investment coefficient (human and physical capital are complementary inputs). The results of two auxiliary regressions in which fertility and the investment rate are the dependent variables tend to confirm these results, for they show that high school enrollments are associated with high investment shares and low fertility rates.[7]

In the papers cited so far the introduction of human capital variables is justified mainly by their possible impact on the rate of innovation and technology adoption. In principle, the best variable to capture such effects would be some indicator of the average educational attainment of the labour force. However, the lack of comparable data for a sufficient number of countries forces the three authors to use flow variables (enrollment rates) as proxies for the relevant stock variables.[8] Although all of them take the precaution of using lagged enrollment rates, these could be highly correlated with investment in human capital over the sample period. Hence, the results of these studies do not allow us to discriminate clearly between level and rate effects. The work of Kyriacou (1991), however, provides more direct evidence of the importance of the second type of effects. Using a procedure described in

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[7] Barro and Lee (1994a) provide a more detailed analysis of the relationship between education and fertility. Benhabib and Spiegel (1994) also find that education has a positive effect on investment.

[8] Landau (1983 and 1986) uses a weighted average of the primary, secondary and university enrollment rates. Baumol et al (1989) reestimate the same equation with each of these variables and find that secondary schooling yields the best results. They argue that this is the preferable variable from a theoretical point of view, since it should be the best proxy for the technological absorption capacity of a broad segment of the population. Barro (1991) includes the primary and secondary enrollment rates as separate explanatory variables.

Section 2a of this Appendix, this author constructs an estimate of the average stock of human capital (the average years of schooling of the labour force, *YS*) which he then includes in convergence regressions with results qualitatively similar to those we have just discussed.[9]

Table A3.1: Human capital in ad-hoc convergence equations

Source:	Q_0	$\boldsymbol{H_0}$	s_k	n		*other variables:*	*sample:*
[1] Landau	-0.0021	0.026			N = 96	GCONS (-). POP (0).	1961-76
(1983)	(6.18)	(7.64)			R^2 = 0.82	CLIM (Y)	96 countries
[2] Landau	-0.311	0.032	0.059	-0.262	N = 151	POP (0). GCONS (-).	1960-80
(1986)	(4.80)	(4.87)	(1.37)	(1.35)	R^2 = 0.714	GINV (0). GED (0). T (0).	65 countries
						INF (-). OIL (+). DP (-)	
[3] Baumol	0.622				N = 103		1960-81
et al (1989)	(1.72)				R^2 = 0.029		103 countries
	-1.47	1.615*			N = 103		
	(2.47)	(5.00)			R^2 = 0.227		
[4] Barro	-0.0075	0.0305*			N = 98	GCONS (-). DISTOR (-).	1960-85
(1991)	(6.25)	(3.86)			R^2 = 0.56	REV (-). ASSAS (-)	98 countries
		0.025**					
		(4.46)					
	-0.0077	0.01*	0.064	-0.004	N = 98	GCONS (-). DISTOR (-).	
	(8.56)	(1.15)	(2.00)	(3.07)	R^2 = 0.62	REV (-). ASSAS (-)	
		0.0118**					
		(2.07)					
[5] Kyriacou	-0.009	0.0062			N = 89		1970-85
(1991)	(2.43)	(4.09)			R^2 = 0.17		89 countries

- *Notes:*
- *t* statistics are shown in parentheses below each coefficient.
- *N* is the number of observations in the sample.
- The dependent variable is the average growth rate of real per capita income during the sample period.
- Definition of H_0: (*) = secondary enrollment rate, (**) = primary enrollment rate. Landau uses a weighted average of three enrollment rates (primary, secondary and university), and Kyriacou an estimate of the average number of years of schooling of the population.
- Other variables: *GCONS* = public consumption/GDP; *POP* = total population; *CLIM* = climate zone dummy; *T* = trend; *GINV* = public investment/PIB; *GED* = public expenditure in education/GDP; *INF* = inflation rate; *OIL* = dummy for oil producers; *DP* = distance to the closest harbour; *DISTOR* = Barro's index of distortions affecting the price of capital goods; *REV* = no. of coups and revolutions; *ASSAS* = number of political assasinations.
- *(+)* and *(-)* indicate a significative coefficient of the corresponding sign; *(Y)* denotes significance, and *(0)* lack of it.
- Landau (1986) uses pooled data with 4-year subintervals; the rest of the regressions use cross-section data by countries.

[9] Actually, the interpretation problem does not disappear completely since Kyriacou's estimate of *YS* is a weighted sum of enrollment rates in the relatively recent past.

Table A3.2 shows some of the results of several more recent studies by Barro and various coauthors using a pooled data set with two or three observations per country based (mostly) on decade-long averages for a large sample of countries. The data comes from various versions of Summers and Heston's Penn World Table and Barro and Lee's (1993) schooling data set and from miscellaneous other sources. The methodology is similar in all the cases: a separate cross-section convergence equation is estimated for each period, imposing the equality of the coefficients across equations and instrumenting some of the regressors with their lagged values in order to mitigate possible endogeneity biases.

The results of the different studies are largely consistent with each other and generally supportive of the view that human capital has a positive effect on growth. The log of life expectancy, which can be expected to be a good proxy for the health component of human capital, appears with a positive and highly significant coefficient in all the equations shown in the table.[10] The pattern of results for the schooling indicators is, as we will see, more complex but is generally consistent with the existence of some sort of positive growth effect and suggests also that an increase in educational attainment helps to speed up convergence, possibly by facilitating the adoption of foreign technologies.

Barro and Lee (B&L, 1994) find that the average number of years of male secondary schooling *(male YS_{sec})* enter the equation with a positive and significant coefficient (equation [1]). This variable, moreover, behaves better than the corresponding flow variable as can be seen in equation [2], where the secondary enrollment rate *($SEC.ENR$)* is not significant. The number of years of university education (YS_{high}), which is added as a regressor in equation [3], is also not significant. Finally, equation [4], which includes both the stock variable and its first difference *(male ΔYS_{sec})*, suggests that male secondary schooling has both level and rate effects. Most of these findings are replicated by Barro and Sala i Martin (B&S 1995) (see equation [5]). In this study, however, the change in the years of male secondary schooling is not significant. On the other hand, B&S find indications that educational expenditure matters and that human capital contributes to fast convergence. This can be seen in equation [5], where public expenditure in education measured as a fraction of GDP *(GED)* and the interaction term between log initial income per capita and average human capital[11] *(H^*q_0)* are significant and display the expected sign. Finally, Barro (1997) confirms the significance of a broader indicator of male schooling (the average years of

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[10] Sachs and Warner (1997) also find that this variable enters significantly in a growth regression. In a more recent paper that uses essentially the same methodology and a slightly longer sample, Barro (2000) finds that health-related variables generally display the expected signs but are often not significant. Except for this, the results of this study are very similar of those of previous ones by the same author.

[11] See the notes to the table for the definition of *H*.

Table A3.2: Results of Barro and various coauthors

	[1]	[2]	[3]	[4]	[5]	[6]
Life expectancy	0.0801	0.0829	0.0806	0.0903	0.076	0.0418
	(5.76)	(5.28)	(5.80)	(6.10)	(5.07)	(3.01)
male YS_{sec}	*0.0138*	*0.0133*	*0.0136*	*0.0199*	*0.0164*	
	(3.29)	(3.09)	(3.16)	(4.15)	(2.83)	
male $YS_{sec+high}$						0.0098
						(3.92)
male YS_{high}			0.000		0.053	
			(0.00)		(1.77)	
H qo*					-0.209	-0.0052
					(2.16)	(3.06)
male ΔYS_{sec}				0.289	0.0066	
				(2.39)	(1.02)	
SEC .ENR male		0.0072				
		(0.62)				
GED					0.205	
					(1.90)	
female YS_{sec}	*-0.0092*	*-0.008*	*-0.0061*	*-0.0162*	*-0.0102*	
	(1.96)	(1.60)	(1.22)	(3.00)	(1.44)	
female YS_{high}			-0.021		-0.071	
			(0.88)		(1.97)	
female ΔYS_{sec}				-0.453	-0.0128	
				(2.35)	(1.54)	
SEC .ENR female		-0.0119				
		(0.73)				
R^2 *(N)*	0.56 (85)	0.56 (85)	0.56 (85)	0.58 (85)	0.64 (87)	0.60 (80)
	0.56 (95)	0.56 (93)	0.57 (95)	0.57 (95)	0.53 (96)	0.52 (87)
						0.47 (84)
Source:	B&L (94)	B&L (94)	B&L (94)	B&L (94)	B&S (95)	B (97)

Notes:
- *t* statistics in parentheses below each coefficient.
- *Additional control variables:* All the equations control for the log of initial GDP per capita (-) and for the following variables (see the notes to the previous table): *GCONS* (-), *REV* (-) and *BMP* (-), where the last variable is the black market premium on foreign exchange and government consumption is measured net of education and defense expenditure. All equations except [6] control for the investment ratio, which is always positive and significant in all cases except for equation [5]. Equations [5] and [6] include also the change in the terms of trade (+). Equation [5] includes as regressors the change in male and female higher schooling, which are not significant. Equation [6] also controls for the log of the fertility rate (-), and index of democracy (+) and its square (-), the inflation rate (-) and dummies for Sub-Saharan Africa, Latin America and East Asia, which have the expected signs but are not significant.
- In equations [1]-[5], two separate regressions are estimated for 1965-75 and 1975-85 (hence the two values of R^2 and sample size reported in the table). In equation [6] the procedure is similar but there is a new observation for 1985-90. The equality of the coefficients across equations is (presumably) imposed. Some regressors are instrumented by their own lagged values.
- The human capital indicator *H* that is used to construct the interaction term with initial GDP per capita (H^*q_0) is different in equation [5] and in equation [6]. In the first case, *H* is the average of five human capital indicators: life expectancy and four schooling variables (male and female average years of secondary and higher schooling), all measured in deviations from sample means. In the second, *H* is the years of male secondary and higher schooling.
- Human capital data from Barro and Lee (1993) and from subsequent revisions of this data set in Barro (1997).
- *Sources:* B&L = Barro and Lee (1994); B&S = Barro and Sala i Martin (1995); B = Barro (1997).

secondary and higher education, *male* $Y_{sec+high}$) and of the interaction effect between schooling and initial income (see equation [6]).

One difficulty with these results is that it is difficult to establish whether the positive schooling coefficients should be interpreted as evidence of level or rate effects.[12] In part, the problem arises because Barro and his coauthors do not use a structural specification that can be used to distinguish sharply between these two effects. A second problematic aspect of Barro et al's studies has to do with their puzzling results about the growth effects of female schooling. The coefficient of female educational variables is often negative and sometimes significant in B&L (1994) (equations [1]-[4]) and in B&S (equation [5]) and not significantly different from zero with the revised schooling data used in Barro (1997).

In a comment to B&L's (1994a) paper, Stokey (1994) provides a possible explanation for these results on the basis of a combination of measurement error and the existence of a handful of influential and atypical observations (in particular, those corresponding to the so-called East-Asian tigers, which are characterized by very high growth rates and display large educational differences across sexes). She suggests dropping the female schooling variable and conjectures that, given its high correlation with male schooling, the coefficient of the latter will fall, casting some doubt on its statistical significance. Lorgelly and Owen (1999) take up Stokey's suggestion and, using the same data, explore the sensitivity of Barro and Lee's results to the omission of the Asian tigers (Hong Kong, Singapure, Taiwan and Korea) and of female schooling. Their results confirm that omitting the East-Asian economies renders both male and female secondary schooling insignificant and that omitting female schooling in the full sample considerably reduces the coefficient of male secondary attainment. When the two schooling variables are combined into a single measure of average years of schooling of the entire population, this variable is only borderline significant. The authors interpret their findings as an indication of the statistical fragility of Barro and Lee's results -- an issue which is also raised in a more general context by Levine and Renelt (1992).

Barro (1997) illustrates and discusses a problem to which we will return repeatedly below. He notes, in particular, that some of his key results (and in particular those pointing to positive growth effects of human capital) tend to break down when the estimation is done in first differences in order to eliminate country-specific effects. This is illustrated in Table A3.3, where the original pooled-data results (using a slightly different specification from

[12] Barro and his coathors tend to interpret the positive coefficient on schooling in terms of the contribution of education to the absorption of technology and the effects of imbalances between the stocks of human and physical capital. For a given initial income, countries with high schooling will tend to grow faster because their stock of physical capital will be low, relative to their stock of human capital, and physical capital can be accumulated more rapidly.

the one shown in Table A3.2) are compared with those obtained with two alternative specifications, a single cross-section in levels with all variables averaged across subperiods, and an equation in first differences. It is interesting to note that, while the results of the cross-section and pooled data specifications are rather close, at least qualitatively, the use of first differences leads to the loss of significance of the educational variables and actually reverses the sign of their coefficients. Barro argues that the first difference specification has several important drawbacks. The main one is that it wastes all the cross-sectional information in the data (which accounts for most of the variation in the regressors) and therefore gives less precise estimates. In addition, he stresses that estimates obtained with first-differenced data are more likely to suffer from measurement error bias and less robust than other estimates to the likely misspecificationi of the timing of the impact of the explanatory variables on growth. While admitting concern about the problem raised by the sensitivity of the results to the specification, he argues that implausible "panel" results such as those given inequation [3] should be heavily discounted.

Table A3.3: Alternative specifications in Barro (1997)

	[1]	*[2]*	*[3]*
Life expectancy	0.0388	0.0172	-0.0820
	(3.13)	(0.93)	(2.15)
$YS_{sec+high}$	0.0123	0.0141	-0.0032
	(5.35)	(4.70)	(0.71)
H^* *qo*	-0.0070	-0.0077	0.0052
	(4.67)	(4.05)	(1.49)
specification:	pooled	cross-section	first differences

- *Note:* - *t* statistics in parentheses below each coefficient. All equations control for the same additional variables as equation [6] in Table A3.2 except for the regional dummies. Equations [1] and [3] are estimated using a SUR technique without instrumenting some of the regressors and equation [2] is estimated by OLS.

b. Results from structural convergence equations

Many recent studies of growth and convergence have made use of the structural convergence equations derived by Mankiw, Romer and Weil (MRW 1992) from a log-linear approximation to an extended Solow model. In this section we will review the results of a number of these studies, starting with MRW's very influential paper. As we will see, the pattern of results on human capital is very similar to the one we found in the previous section. Cross-section and pooled estimates generally yield positive results that are consistent with the existence of sizable level effects. On the other hand, fixed effects and first-difference specifications that rely on the time series variation in the data often produce insignificant or even negative estimates of the coefficient of human capital in the aggregate production function. As will be emphasized in a later section, a possible explanation for these negative findings is related to the weak signal content of differenced schooling data.

Mankiw, Romer and Weil (1992) use cross-section data for the period 1960-85 to estimate a structural convergence equation of the form[13]

$$(A3.1)\quad q_{iT} - q_{io} = \Gamma + (1-e^{-\beta T})\left(\frac{\alpha_k}{1-\alpha_k-\alpha_h}\, ln\, \frac{s_{ki}}{\delta+g+n_i} + \frac{\alpha_h}{1-\alpha_k-\alpha_h}\, ln\, \frac{s_{hi}}{\delta+g+n_i} - q_{io}\right)$$

where q_{iT} is log output per capita (using as denominator the working-age population) in country i at time T, s_k and s_h the the average rates of investment in physical and human capital over the relevant period, δ the rate of depreciation, which is assumed to be the same for both types of capital, g and n the rates of technical progress and (working-age) population growth. The parameters α_k and α_h and are the coefficients of physical and human capital in a Cobb-Douglas aggregate production function, and β the convergence parameter that measures the speed at which the economy approaches the steady state or long-run equilibrium determined by the observed investment rates. MRW assume $g = 0.02$ and $\delta = 0.03$ and use as their proxy for the rate of investment in human capital (s_h) the fraction of the working-age population enrolled in secondary schooling. Implicitly, they also assume a common level of technical efficiency for all countries or, at least, that cross-country differences in TFP can be safely thrown into the error term. Hence, they treat the term Γ in equation (A3.1) as a constant even though the underlying theoretical model suggests that it should vary across countries with differences in initial levels in TFP .

Columns [1] and [2] of Table A3.4 show MRW's results, including the implied values of the coefficients of the production function and the rate of convergence, for two different samples: one formed by 75 countries, and a second one comprising the 22 OECD countries with population above 1 million. The estimated production function coefficients are in general significant and have the expected sign. Their values, moreover, seem quite reasonable when judged from the a priori expectation that they should reflect the shares of the different factors in national income. According to the estimated model, capital's share in national income would be around 40%. Of the remainder, which is labour's share, almost half would be the return on human capital, whose estimated elasticity (α_h) is 0.23.

MRW's paper was extremely influential beacuse its appealing results seemed to indicate that a simple extension of the standard neoclassical model provided a satisfactory description of the process of growth and of the evolution of the regional (or national) income distribution.[14] The only change required, relative to the more traditional models, was the

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[13] See Box 2 in Section 3b.ii of the main report.
[14] See also Mankiw (1995).

broadening of the relevant concept of capital in order to include the accumulated investment in education.

Because of its popularity, MRW's paper provided the starting point for a large number of empirical studies that attempted to extend the original model in various directions, to test the robustness of its results or to improve the quality of the estimation through the use of better data or more adequate econometric techniques. Columns [3] to [8] of Table A3.4 summarize the results of a group of such studies that, making use of cross-section or pooled data, largely corroborated MRW's results and established their robustness to reasonable extensions of the underlying model. Lichtenberg (1992) and Nonneman and Vanhoudt (N&V 1996) consider a further augmentation of the Solow model in which R&D capital is treated in

Table A3.4: Cross-section and pooled data specifications of the MRW model

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*	*[7]*	*[8]*
β	0.0186	0.0206	0.024	0.017	0.021	0.029	0.033	0.034
	(9.79)	(10.30)		(17.99)	(4.20)			(5.25)
α_k	0.44	0.38		0.474	0.354	0.35		0.301
	(6.29)	(2.92)		(10.09)	(4.12)			(5.07)
α_h	*0.23*	*0.23*		*0.236*	*0.259*	*0.148*		*0.204*
	(3.83)	(2.09)		(4.21)	(3.65)			(3.74)
$\alpha_{R\&D}$					0.066	0.084		0.060
					[(2.54)			(2.22)
$ln\ s_k$	0.506	0.396	0.550			0.413	0.491	
	(5.33)	(2.61)	(2.90)			(2.65)	(3.61)	
$ln\ s_h$	*0.266*	*0.236*	*0.621*			*0.175*	*0.558*	
	(3.33)	(1.67)	(3.37)			(1.55)	(3.60)	
$ln\ s_{R\&D}$						0.098	0.099	
						(1.78)	(2.25)	
specification	cr.sect.	cr.sect.	cr sect	cr.sect.	cr.sect.	cr.sect.	cr.sect.	pooled
	1960-85	1960-85	1960-85	1960-85	1960-85	1960-85	1960-85	1965-95
sample	75 ctries.	22 OECD	22 OECD	53 ctries.	53 ctries.	22 OECD	22 OECD	19 OECD
source	MRW	MRW	V&C	Licht.	Licht.	N&V	V&C	dF

Notes:

- t statistics in parentheses below each coefficient. For ease of comparison, I have calculated some of them using the originally reported standard errors. These calculations may not be entirely accurate due to rounding error.
- Source: MRW = Mankiw, Romer and Weil (1992); V&C = Vasudeva and Chien (1997); Licht. = Lichtenberg (1992); N&V = Nonneman and Vanhoudt (1996); dF = de la Fuente (1998).
- Some authors estimate the coefficients of the production function directly, others infer them from the coefficients of the $ln\ s_i$ terms and others report only the latter.
- dF controls also for the share of government spending in GDP and changes in the unemployment and labour force participation rates and includes a dummy for technological laggards and the interaction of this variable with a trend. In this paper, the convergence equation is estimated using pooled data with averages over five-year periods and the proxy for s_h is total secondary enrollment as a fraction of the labour force, averaged over the current and previous five-year subperiods.

the same way as physical and human capital. De la Fuente (1998) further controls for government spending, labour-market indicators and technological diffusion, and considers a broader measure of human capital investment that takes into account university as well as secondary schooling. In the same line Vasudeva and Chien (1997) replicate MRW's and N&V's estimates using as a proxy for educational investment a weighted average of the primary, secondary and university enrollment rates (with weights of 0.2, 0.3 and 0.5 respectively). As can be seen in the Table, the results are generally quite satisfactory. Human capital only fails to be significant at conventional levels in N&V (column [6]) and (if we consider the coefficient of $ln\ s_h$ rather than the corresponding parameter of the production function) in MRW's OECD subsample (column [2]). Using essentially the same data and the exact same sample, N&V show, however, that results improve considerably when a broader measure of educational investment is used.

Table A3.5: Various specifications of MRW model

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*	*[7]*	*[8]*
β	0.0186	0.014	0.0206	0.015	0.0142	0.014	0.047	
	(9.79)		(10.30)		(7.45)			
α_k	0.44		0.38		0.48		0.468	
	(6.29)		(2.92)		(6.86)		(5.57)	
α_h	*0.23*		*0.23*		*0.23*		*-0.121*	
	(3.83)		(2.09)		(4.60)		(1.53)	
$ln\ s_k$	0.506	0.66	0.396	0.13	0.500	0.59		
	(5.33)	(5.50)	(2.61)	(0.65)	(9.62)	(6.56)		
$ln\ s_h$	*0.266*	*0.00*	*0.236*	*0.13*	*0.238*	*-0.01*		*0.00*
	(3.33)	(0.00)	(1.67)	(0.76)	(3.97)	(0.17)		(0.08)
specification	cr.sect. 1960-85	cr sect 1960-85	cr.sect. 1960-85	cr.sect. 1960-85	cr.sect. 1960-85	cr.sect. 1960-85	diff.	cr.sect. 1960-85
sample	75 ctries.	69 ctries.	22 OECD	21 OECD	98 ctries.	92 ctries	98 ctries.	58 LDCs
source	MRW	Temple (1998a)	MRW	Temple (1998a)	MRW	Temple (1998a)	H&M (1998)	Temple (1998b)

Notes:
- t statistics in parentheses below each coefficient (calculated using the originally reported standard errors).
- Equations [2] and [6] from Temple (1998a) include dummies for Africa (-, -), Latin America (0, 0), East Asia (0, +) and the industrial countries (0, 0).
- The countries considered atypical by Temple (1998a) and excluded from the original samples of MRW are Japan in the OECD sample (equation [4]); Argentina, Cameroon, Chile, Hong-Kong, India and Zambia in the intermediate sample (equation [2]) and Chad, Chile, Hong-Kong, Mauritania, Somalia and Zambia in the broader sample (equation [6]).
- Equation [8] controls for investment in equipment (+) and structures (+) and includes dummies for Latin America (0), Africa (-) and East Asia (+). The schooling variable is also non-significant in other samples, especially when regional dummies are included.

On the other hand, a second set of studies stemming from MRW's paper have shown that these authors' results are not robust along a number of dimensions. Temple (1998a) shows that

MRW's results are largely driven by a few influential observations. To identify outliers, Temple first estimates the model by a robust estimation technique (least trimmed squares, due to Rousseeuw (1984)) that fits the model to the half of the sample that provides the best fit, uses the results to identify as outliers those countries with the greatest residuals, and then reestimates the model by OLS after excluding outliers. His results for the three samples considered by MRW are shown in Table A3.5 (equations [2], [4] and [6]) next to MRW's original results (equations [1], [3] and [5]) that are reproduced here for convenience. In all cases, he finds that the exclusion of a few outliers (listed in the notes to the table) renders the coefficient of human capital insignificant. The same author (Temple 1998b) also finds that schooling is not significant in a variety of samples in an extension of MRW's model in which investment in physical capital is disaggregated into its equipment and structures components following De Long and Summers (1991) (see equation [8] in Table 7).

Hamilton and Monteagudo (1998) find that MRW's schooling indicator also loses its significance when their model is used to try to explain changes in growth performance across decades. They essentially reestimate MRW's model in first differences (calculated as the difference between average values for 1960-70 and 1975-85) with the results shown in equation [7] in Table A3.5: while the coefficient of investment in physical capital is very similar to the original estimate (equation [5]), the point estimate of the schooling variable is actually negative.

Table A3.6: Jones (1996)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*
$ln\ s_k$	0.425	0.437	0.394	0.506	0.377	0.353
	(2.85)	(2.60)	(4.15)	(3.95)	(2.73)	(3.72)
ys	*1.032*	*0.500*	*-0.050*			
	(5.61)	(3.65)	(0.39)			
YS				*0.191*	*0.189*	*0.159*
				(6.16)	(6.10)	(2.48)
implied α_k	0.298	0.304	0.282	0.336	0.274	0.261
	(4.08)	(3.75)	(5.76)	(6.00)	(3.75)	(5.02)
implied α_{YS}	0.724	0.348	-0.036			
implied ρ				0.127	0.137	0.118
R^2	0.668	0.522	0.141	0.678	0.571	0.205
specification	levels	levels	differences	levels	levels	differences
year	1960	1990	1990-60	1960	1990	1990-60

Notes:
- Summers-Heston data for 78 countries. Years of schooling are from Barro and Lee (1993).
- *t* statistics in parentheses below each coefficient.
- The rates of investment rate (s_k) and population growth (n) are averages over relatively short periods around the year whose output level is taken.

A study by Jones (1996) reaches rather more optimistic conclusions regarding the contribution of schooling to productivity using a mincerian specification. Starting from a different theoretical model (that emphasizes the role of ideas and technological diffusion), this author derives a steady-state equation that is identical to the one implied by MRW's model when the stock of human capital H is an exponential function of the average years of schooling, YS. Assuming that countries have reached their steady states, Jones derives an expression that relates (the log) of per capita income, q_{it}, to the rate of investment in physical capital (s_{kit}), average years of schooling (YS) and log TFP (a). When we interpret it as coming from MRW's model, this equation can be written as follows

$$(A3.2) \quad q_{it} = c_0 + a_{it} + \frac{\alpha_k}{1-\alpha_k} \ln \frac{s_{kit}}{\delta+g+n_{it}} + \frac{\rho}{1-\alpha_k} \; YS_{it}$$

Jones estimates this equation and its standard (non-mincerian) MRW counterpart (with $\frac{\alpha_h}{1-\alpha_k}$ ys_{it} replacing the last term in (A3.2)) using data in levels for 1960 and 1990 (without controling for possible differences in TFP levels, a_{it}), and with the variables measured in differences across these two years. As can be seen in Table A3.6, the results vary dramatically depending on the specification chosen for the schooling variable. When years of schooling enter the equation in logs (equations [1]-[3]), the results are similar to those obtained by Hamilton and Monteagudo (1998): the coefficient of the human capital variable (which is positive and significant in the cross-section) becomes negative in the differenced specification. When YS is entered in levels, by contrast, the human capital coefficient is always positive and significant, and the estimated value of the returns to schooling parameter (ρ) is slightly above 10%, which is above the average of the available microeconometric estimates when these are properly adjusted.[15]

Panel data specifications

The doubts about the growth effects of educational investment that were first motivated by the apparent statistical fragility of some earlier results have been reinforced in recent years by a set of papers which have approached the empirical analysis of convergence from a panel data perspective. Knight, Loayza and Villanueva (KLV, 1993), Islam (1995) and Caselli, Esquivel and Lefort (CEL, 1996) reestimate the MRW model introducing various fixed effects specifications to pick up possible cross-country differences in levels of TFP. In addition, CEL use an instrumental variables technique to allow for the likely endogeneity of some of the regressors. The results of all three papers indicate that panel estimates of the

[15] Psacharopoulos (1994) reports an average microeconometric estimate of the return to schooling of 10.1% for a large sample of countries. The adjustment required to make this figure comparable to macroeconometric estimates brings it down to 6.7%.

MRW model which rely heavily on the time series variation of the data generally yield insignificant or negative coefficients for human capital.

Table A3.7: Panel estimates of the MRW model

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*	*[7]*	*[8]*
β			0.0069	0.0375	0.0162	0.0913	0.0107	0.0679
			(2.76)	(4.03)	(2.95)	(5.71)	(3.96)	(3.30)
α_k			0.8013	0.5224	0.6016	0.2074	0.496	0.491
			(15.01)	(8.14)	(5.93)	(1.97)	(6.44)	(4.31)
α_h			*0.0544*	*-0.199*	*0.0174*	*-0.045*	*0.18*	*-0.259*
			(0.53)	(1.81)	(0.10)	(0.31)	(3.33)	(2.09)
$ln\ s_k$	0.105	0.023						
	(10.16)	(1.61)						
$ln\ s_h$	*-0.111*	*-0.065*						
	(13.26)	(5.09)						
specification:	fixed effects	fixed effects	pooled ols	fixed effects	pooled ols	fixed effects	pooled ols	fixed eff. &IV
sample:	75 LDCs	96 ctries.	79 ctries.	79 ctries.	22 OECD	22 OECD	97 ctries.	97 ctries.
source:	KLV	KLV	Islam	Islam	Islam	Islam	CEL	CEL

- *Note:* Panel data from Summers and Heston's PWT for 1960-85 with 5-year subperiods. *t* statistics in parentheses below each coefficient (in the case of CEL and Islam, they are calculated using the originally reported standard errors).

This finding is illustrated in Table A3.7, which summarizes some of the key results of these studies. Islam uses a variant of the MRW model in which the growth rate of output per worker appears as a function of the log of the stock of human capital, which is proxied by current average years of schooling from Barro and Lee (1993), rather than as a function of school enrollments. CEL, on the other hand, deviate in this respect from MRW only in that they use the secondary enrollment ratio as a proxy for the investment rate in human capital, and KLV use the same schooling variable as MRW. In spite of these differences in the choice of regressors, and additional differences in the way the fixed effects model is implemented, the results are broadly similar. The estimated coefficient of human capital in the production function is positive and sometimes significant in either cross-section or pooled data specifications, but becomes negative and often significant when fixed country effects are added to the equation. KLV also report that the coefficient of schooling is positive and highly significant when only its average value for each country is used in the regression.

It is interesting to note that the reaction of the authors to their findings regarding human capital is quite different. KLV argue that, because of the long time lags involved, it makes little sense to use quinquennial enrollment rates as a proxy for the relevant investment in human capital, and advocate disregarding the time-series variation in this variable in the estimation (which, as we noted above, yields positive schooling coefficients). Islam (1995) tries to rescue human capital as a determinant of the level of technological development (which is presumably what is being captured by the country dummies) by observing that the fixed effects are highly correlated with standard measures of educational achievement. The argument, however, merely sidesteps the problem: we know that human capital variables work well with cross-section data, but if they really had an effect on the level of technical efficiency they should be significant when entered into the panel equation. Finally, Caselli et al (1996) seem quite willing to take their negative findings at face value.

c. Production function estimates and related specifications

A third group of papers has examined the growth effects of human capital through the estimation of aggregate production functions and related specifications. As far as we know, the earliest studies in this branch of the literature are due to Kyriacou (1991) and Benhabib and Spiegel (B&S 1992, 1994), who estimate a Cobb-Doulgas production function using a single cross-section of growth rates computed over a long period and Kyriacou's (1991) schooling data set. Pritchett (1999) undertakes a similar exercise after constructing a "mincerian" stock of human capital using microeconometric estimate of the returns to schooling parameter and data from both Barro and Lee (1993) and Nehru et al (1995). Finally, Temple (1999 and 2001) uses B&S's and Pritchett's data to examine the robustness of their results to outliers and to some changes in the specification.

The key results of these studies are summarized in Table A3.8. The coefficient of the human capital variable (α_h or α_{YS}) is either non-significant or negative in the basic specifications used in the earlier three studies (equations [1] to [4]). The authors also show that this result is robust to a number of changes in the specification, such as the inclusion of regional dummies or initial income per capita to control for a technological catch-up effect.

Kyriacou (1991) also tests for threshold effects and various non-linearities with generally negative results. Pritchett (1999) argues that the results do not seem to be due to measurement error in human capital, as they remain essentially unchanged when the estimation is repeated using the Nehru et al schooling data to construct an instrument for his (Barro and Lee based) human capital stock (equation [5]). On a somewhat more positive note, Temple (1999 and 2001) finds that the elimination of outliers does generate a positive and significant human capital coefficient[16] (equation [6], but notice that this requires the elimination of 14 out of 78 countries), and that the mincerian (log-level rather than log-log) specification

[16] Temple (1999) follows essentially the same procedure as a previous paper of the same author we have already commented upon (Temple, 1998a). He uses large residuals from LTS estimates to identify influential observations and deletes them before reestimating the equation by OLS.

produces better results than the Cobb-Douglas when the Barro and Lee (1993) data are used (but not with Kyriacou's data, see equations [7] and [8]). Even in this case, however, the schooling variable becomes only borderline significant when regional dummies are added to the equation (equation [9]).

Table A3.8: Aggregate production functions with human capital

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
α_k	0.449	0.457	0.524	0.501	0.460	0.553	0.432	0.490	0.462
	(5.05)	(5.38)	(12.8)	(15.4)	(10.18)	(13.16)	(5.08)	(8.18)	(5.97)
α_l	0.261	0.209				0.241	0.266		
	(0.90)	(1.01)				(2.15)	(1.38)		
α_h / α_{YS}	-0.152	0.063	-0.049	-0.104	-0.120	0.165			
	(1.68)	(0.80)	(1.07)	(2.07)	(1.42)	(4.00)			
ρ							0.015	0.080	0.062
							(0.52)	(2.56)	(1.76)
notes:					IV				reg dum
N	87	78	91	79	70	64.	78	91	91
period	1970-85	1965-85	1960-87	1960-87	1960-87	1960-85	1960-85	1960-87	1960-87
H data	Kyr.	Kyr.	B&L	NSD	B&L, N	Kyr.	Kyr.	B&L	B&L
Source:	Kyr.	B&S	Prit. 99	Prit 99	Prit 99	T99	T01	T01	T01

Notes:

- *t* statistics in parentheses below each coefficient (some of them are computed using the originally reported standard errors). *N* is the number of observations (countries) in the sample.
- ρ is the coefficient obtained from a Mincerian specification, where the regressor is the change in the years of schooling rather than the change in their logarithm. In the case of Pritchett, a Mincerian estimate of the stock of human capital (based on an exponential function of the years of schooling and an outside estimate of the relevant coefficient) is inserted into a standard Cobb-Douglas production function and α_h is the elasticity of this function. Constant returns to scale are imposed when the coefficient of labour (α_l) is not shown.
- Cross section data and estimation in long differences or average growth rates by OLS except in equation [5] where instrumental variables are used. Equation [9] includes regional dummies (presumably for Africa, Latin America, East Asia and developed countries, although the author does not say it explicitly).
- For Kyriacou and Benhabib and Spiegel, the dependent variable is the log change in total output during the sample period; in Pritchett, it is the growth rate of output per worker. Pritchett uses least-squares logarithmic growth rates of output and factor stocks.
- Capital stocks are obtained by accumulating investment flows.
- *Sources:* Kyr = Kyriacou (1991); B&S = Benhabib and Spiegel (1994); Prit = Pritchett (1999); T99 = Temple (1999) and T01 = Temple (2001).
- *Sources of human capital data:* Kyr = Kyriacou (1991); B&L = Barro and Lee (1993) and N = Nehru, Swanson and Dubey (1995).

Rate effects and interaction with technological diffusion

The results of the production function studies we have just reviewed are largely consistent with the hypothesis that the stock of human capital does not enter the production function as a productive input (i.e. that there are no level effects). Some of these papers, however, do find rather clear indications that the level of education is an important determinant of the rate of technological progress. This positive rate effect, moreover, seems to work at least in part through the role of education in facilitating the absorption of foreign technologies.

Following the work of Nelson and Phelps (1966) and Romer (1989), Kyriacou (1991) argues that the level of education (rather than its first difference) should be included in a growth equation as a determinant of the rate of technological progress. This hypothesis leads to equations [1] and [2] in Table A3.9, where he introduces the log of average years of schooling *(ys)* or the level of the same variable *(YS)* and its square to try to capture rate effects with encouraging results, particularly in the second case.

Table A3.9: Rate effects in aggregate production functions with human capital

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
α_k	0.435 (4.88)	0.417 (6.24)	0.479 (5.10)	0.4723 (6.59)	0.5005 (6.49)	0.5076 (5.38)	0.5517 (4.50)	0.5233 (3.66)
α_l	0.176 (0.58)	0.387 (1.49)	0.391 (2.01)	0.188 (1.15)	0.2045 (1.31)	0.1720 (0.74)	0.5389 (1.39)	0.2901 (0.57)
α_{YS}	0.018 (0.12)	0.0359 (0.34)						
q_o			-0.235 (5.11)					
YS		0.0101 (3.25)		-0.00136 (0.09)	0.0021 (0.14)	0.0439 (1.96)	-0.0003 (0.01)	-0.0736 (1.26)
YS^2		-0.001 (3.07)						
ys	0.0068 (1.79)		0.167 (3.09)					
$YS*(Q_l/Q)$				0.0011 (5.50)	0.0007 (2.33)	0.0003 (0.33)	-0.0001 (0.11)	0.0012 (4.00)
Q_l/Q					0.0014 (1.40)			
notes:						rich	middle	poor
N	87	87	78	78	78	26	26	26
Source:	Kyr.	Kyr.	B&S	B&S	B&S	B&S	B&S	B&S

Notes:
- The human capital variable used in all the equations is the average number of years of schooling from Kyriacou (1991). The sample period is 1970-85 in Kyriacou and 1965-85 in Benhabib and Spiegel.
- Equation [3] include continent dummies for Latin America and Africa; equation [2] includes dummies for oil producers, mixed economies and Latin America, as well as an index of political instability.

Benhabib and Spiegel (B&S 1994) follow a similar route and extend the model to allow for technological diffusion and rate effects from human capital. In equation [3] they add the log of the stock of human capital *(ys)* to capture rate effects and the log of initial income per capita income *(q_o)*, interpreted as a proxy for the inital level of technical efficiency, to control for a technological catch-up effect. Both variables are significant and have the expected signs.

Starting from this last specification, B&S try to characterize more precisely the channels through which human capital contributes to technological progress. For this purpose they estimate a "more structural" model in which they include as regressors, in addition to the average years of schooling, YS, (which should capture human capital's contribution to domestic innovation), the ratio (Q_l/Q) between output per worker in the leading country and that in each member of the sample (as a proxy for technological backwardness), and the product of these two variables to capture interaction effects. The results for the complete sample (equations [4] and [5] in Table A3.9) suggest that human capital's effect on growth works mostly through its contribution to technological diffusion and absorption as signaled by the fact that only the interaction term is significant. The results, however, change with the level of development. When the same equation is reestimated separately for each of three subsamples, the catch-up effect dominates in the poorest countries (equation [8]), while the contribution to domestic innovation is more important in the richer group (equation [6]). Neither of these variables is significant in the case of the middle-income subsample (equation [5]).

Table A3.10: Engelbrecht (1997)

	[1]	[2]	[3]	[4]
$\Delta\ ln\ R^d$	0.072 (5.29)	0.098 (6.83)	0.098 (6.70)	0.105 (7.15)
$\Delta\ G7{*}ln\ R^d$	0.17 (5.54)	0.175 (5.01)	0.163 (4.77)	0.166 (4.86)
$\Delta\ m\ ln\ R^f$	0.198 (3.93)	0.303 (5.75)	0.249 (4.56)	0.249 (4.72)
Δys	0.136 (3.89)			
$ys(*)$		-0.007 (0.42)	0.141 (2.93)	0.128 (2.71)
$ys{*}\ ln\ Q_i/Q_{us}$			0.127 (3.34)	0.107 (2.82)
$ln\ Q_i/Q_{us}$				-0.260 (4.51)

Notes:
- Annual panel data for the period 1970-85. Annual data on years of schooling are constructed by interpolating between quinquenial observations from Barro and Lee (1993). *t* statistics in parentheses below each coefficient.
- G7 = dummy variable, = 1 for the G7 countries; *m* is the share of imports in GDP.
- Equations [2]-[4] include both period and country dummies.
- (*) Notice that the coefficient of *ys* is not really an estimate of γ_h as we have defined it. To recover the latter parameter (which measures the contribution of *YS* to TFP growth (rather than that of its logarithm), we have to divide the relevant coefficient in the table by the value of *YS*).

A more recent study that also finds evidence of rate effects in a more complete model is due to Engelbrecht (1997). This paper investigates the connection between education and technical progress using an extension of the model estimated by Coe and Helpman (1995). These authors examine the relationship between (estimated) total factor productivity (TFP) and domestic and foreign R&D investment. For each country in a sample of 21 developed economies, an estimate of the domestic stock of technological capital (R^d) is constructed by accumulating past R&D expenditures. To allow for cross country spillovers, the level of domestic TFP is also allowed to be a function of the stock of foreign R&D capital (R^f), defined as an average of the domestic stocks of a country's trading partners weighted by the share of each country in total domestic imports.

Drawing on Benhabib and Spiegel (1994), Engelbrecht (1997) extends Coe and Helpman's model to allow the rate of TFP *growth* to be a function of the log of average schooling or its growth rate and of the technological gap with the leading country, proxied by the ratio of each country's real per capita GDP to that of the US (Q_i/Q_{us}). The resulting model is estimated using Coe and Helpman's data together with Barro and Lee's (1993) series on average years of schooling.

The results of the exercise are summarized in Table A3.10. Since the dependent variable is the growth rate of total factor productivity, the explanatory variables enter in log differences when we are looking for a level effect and in logs when it is expected that they will have a direct effect on the rate of technical progress. The coefficients on the domestic and foreign stocks of R&D, and the interaction between home R&D and size (proxied by a dummy variable for the G7 economies) confirm Coe and Helpman's results about the impact of research expenditures, the importance of trade as a vehicle for technological diffusion and the existence of scale effects in innovation. The coefficients of the human capital indicators are consistent with the existence of both level (equation [1]) and rate effects, although the two hypotheses are not tested simultaneously (presumably because a high correlation between ys and its first difference that would generate multicollinearity problems). It is interesting to note, however, that rate effects (a significant positive coefficient for ys) appear only when we introduce a catch-up term and its interaction with the schooling indicator.[17]

[17] It is interesting to note that the sign of the interaction term between human capital and the relative productivity variable used as a proxy for technological backwardness is the opposite one than in Benhabib and Spiegel (B&S, 1994). Notice that the latter variable is constructed in different ways in the two studies, with own productivity in the numerator in one case and in the denominator in the other. Hence, B&S's results (equation [5] in Table A3.9) imply that rate effects from human capital are higher in technologically backward countries, whereas Englebrecht (equation [5] in Table A3.10) finds that they will be larger in more advanced countries. Notice, however, that the samples are different. Engelbrecht's sample is presumably a subset of Benhabib and Spiegel's rich country subsample, where the interaction term is not significant.

d. Data quality and measurement error

A number of recent papers argue that the negative results found in the earlier literature can be attributed to low data quality and the resulting measurement error bias (see the main text for a discussion). Krueger and Lindhal (K&L 2001) argue that Benhabib and Spiegel's (B&S 1994) widely cited failure to find significant level effects can be attributed to the almost complete lack of signal in the schooling variable they use. According to K&L's estimates, the simple reliability ratio for the relevant regressor (which is the aveage growth rate of Kyriacou's years of schooling over the entire sample period) is only 0.195. Since the R^2 of a regression of this variable on the remaining explanatory variables of B&S's model is about the same size, the expected value of the human capital coefficient in the absence of a correction for measurement error is zero regardless of its true value. A similar argument, combined with available estimates of reliability ratios for the data sets used in the prior literature (see section 2a of this Appendix) suggests that many previous estimates of the coefficients of interest may be similarly flawed.

De la Fuente and Doménech (D&D 2000, 2001a), Cohen and Soto (C&S 2001) and Bassanini and Scarpetta (2001) find clear evidence of sizable and significant level effects using newly constructed data sets which appear to have higher signal to noise ratios than those used in the earlier literature (see Section 2a of this Appendix). De la Fuente and Doménech estimate several production function specifications that allow for level effects using pooled data at quinquennial intervals for a sample of OECD countries. They examine the sensitivity of the results to the quality of the human capital data by reestimating several specifications with three different data sets: their own, and the ones constructed by Barro and Lee (1996) and Nehru et al (1995). Table A3.11 shows the results obtained with their preferred specification, which incorporates a technical progress function allowing for technological diffusion and for permanent TFP differences across countries. The pattern of results that emerges for the different human capital data sets is consistent with the authors' hypothesis about the importance of educational data quality for growth estimates. The human capital variable is significant and displays a reasonable coefficient with their data (D&D 2000, equation [3]), but not with the Nehru et al (NSD) or Barro and Lee (B&L) series (equations [1] and [2]), which actually produce negative human capital coefficients. Moreover, the coefficients of the stocks of physical and human capital estimated with the D&D data are quite plausible, with α_k only slightly above capital's share in national income (which is 0.35 in their sample) and α_{YS} close to one third. Equation [4] is taken from an update of de la Fuente and Doménech (2000) that uses the revised data set described in de la Fuente and

Doménech (2001b). This further revision of the data increases the coefficient of the schooling variable by over one third.

Table A3.11: Results of D&D with different human capital data sets

	[1]	[2]	[3]	[4]
schooling data:	*NSD*	*B&L*	*D&D 2000*	*D&D 2001*
α_k	0.510	0.409	0.373	0.345
	(8.30)	(6.12)	(7.15)	(6.83)
α_{YS}	-0.148	-0.057	0.271	0.394
	(2.62)	(0.88)	(2.53)	(4.57)
λ	0.100	0.063	0.068	0.074
	(6.98)	(8.27)	(6.34)	(7.07)
adj. R^2	0.840	0.811	0.809	0.828

- *Notes:* White's heteroscedasticity-consistent t ratios in parentheses. Only significant country dummies are left in the reported equation. The parameter λ is the coefficient of the technological gap with the US and measures the speed of technological diffusion. Aside from the schooling variable, the data used to estimate equation [4] is slightly different from the data used in the previous equations because it incorporates the latest revision of the OECD's national accounts series. The sample period is 1960-90, except for the NSD data, which only extend to 1985.

In a background study for the OECD growth project (OECD 2001b), Bassanini and Scarpetta (2001) use D&D's (2001) updated schooling series and in-house OECD data to extend the sample period (from 1971) until 1998 and interpolate the schooling series to obtain annual observations for a sample of OECD countries. These authors estimate a convergence equation à la MRW (written in terms of the stock of human capital rather than the investment rate) which includes fixed effects and is embedded into an error-correction model that allows for short-term deviations from the equilibrium path described by the underlying growth model. Their specification permits short-run coefficients and the convergence parameter to differ across countries but imposes (as is usually done in the literature) a common value of the coefficients of the production function. The estimated level effects are highly significant and much larger than those found by D&D (2000). The parameter values obtained in the preferred specification are $\alpha_k = 0.13$ and $\alpha_{YS} = 0.82$, whereas removing Finland from the sample yields $\alpha_k = 0.19$ and $\alpha_{YS} = 0.41$. The authors settle for an intermediate "best-guess" estimate of around 0.60 for α_{YS}, which (since the average years of schooling in the sample is a bit over 10), implies a gross mincerian return to schooling of about 6%.

Cohen and Soto (2001) construct a new data set for a sample of 95 countries which they use to estimate two alternative "mincerian" specifications, finding evidence of sizable level effects. The first one (equations [1] and [2] in Table A3.12) is an MRW-style steady state equation linking income per capita to the rate of investment in physical capital *(s_k)* and school attainment in levels, YS. The (lagged) urbanization rate and continental dummies are used to proxy for differences in TFP levels in this equation. The second specification (equations [3] to [5]) relates the growth rate of income per capita to the average annual change in average years of schooling, ΔYS. The equation includes an LDC dummy and controls for the urbanization rate but not for investment in physical capital. The coefficient of years of schooling in the steady-state equation (which will be an estimate of $\rho/(1-\alpha_k)$) is 0.085 when the equation is estimated by OLS and rises to 0.100 when schooling is instrumented to mitigate any potential endogeneity problems. These estimates imply that the gross return to schooling lies between 5.7% and 6.7%. The estimated coefficient of ΔYS in the growth equation is also consistent with this range of values[18] when Cohen and Soto's own data set is used (see equation [3]), but drop sharply when the estimation is repeated with Barro and Lee's (2000) data set (equation [4]).

Table A3.12: Cohen and Soto (2001)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*
$ln\ s_k$	0.46 (5.7)	0.41 (2.00)			
YS	0.085 (4.0)	0.100 (2.06)			0.00078 (0.76)
ΔYS			0.0845 (2.51)	0.028 (1.45)	0.0864 (2.56)
urban	0.011 (5.3)	0.010 (2.55)	-0.00019 (2.3)	-0.00015 (1.6)	-0.00024 (2.3)
poor			-0.0104 (2.80)	-0.0090 (2.31)	-0.0080 (1.60)
R^2	0.83	0.83	0.20	0.21	0.21
schooling data	C&S	C&S	C&S	B&L 00	C&S
notes	OLS levels	IV levels	OLS growth rates	OLS growth rates	OLS growth rates

The one discouraging feature of the studies reviewed in this section is that they generally do not find clear evidence of rate effects. The coefficient of years of schooling in Cohen and Soto's growth equation, which would be an estimate of γ_h, is positive but extremely small and

[18] The structural interpretation of the coefficient of ΔYS in the growth equations is made difficult by the failure to control for the accumulation of physical capital (K). If there is perfect capital mobility across countries, so that K adjusts instantaneously and optimally to changes in YS, as we have assumed at the individual level to derive the correction factor for microeconometric estimates, the coefficient of ΔYS in equations [3]-[5] will also be an estimate of $\rho/(1-\alpha_k)$) and can therefore be directly compared to the coefficient of YS in the steady state equations.

not significantly different from zero (see equation [5] in Table A3.12). Similarly, the introduction of *ys* in D&D's growth equation yields a positive but small and insignificant coefficient and sharply reduces the precision of the estimate of the level effect (α_{YS}) -- a pattern which suggests that the simultaneous introduction of the level and the growth rate of schooling in the same equation can give rise to serious collinearity problems that make it difficult to untangle level and rate effects. Krueger and Lindhal (K&L 2001), finally, also report some adverse results in this respect. They find that rate effects tend to be positive and significant in standard specifications that constrain the relevant parameter to be equal across countries, but that relaxing this assumption generates insignificant coefficients except for countries with very low levels of education.

e. Educational quality and test scores

All the studies we have reviewed until now use enrollment rates or years of schooling as proxies for investment in human capital or for the stock of this factor. An obvious limitation of these indicators is that they measure only the quantity of schooling. But since workers with the same number of years of schooling may have very different skills across countries depending, among other things, on the quality of national educational systems, one would ideally like to complement the standard schooling indicators with some measure of quality. In this section we review some studies that have tried to do this by using data on educational expenditures and other possible determinants of school quality and/or direct measures of skills such as scores in standardized international achievement tests. Some of these papers also analyze, with conflicting results, the impact of educational expenditures on student achievement.

Dessus (1999) argues that the impact on productivity of an additional year of schooling should vary across countries depending on the quality of the educational system. He uses quinquennial data covering the period 1960-90 for a sample of 83 countries to estimate a variant of the MRW model (written in terms of the stock of human capital) with fixed country effects and a varying parameter specification that makes the coefficient of human capital (α_{YS}) a function of some indicator (QS_i) of the average quality of schooling,

$$\alpha_{YSi} = \alpha_{YS0} + \eta QS_i.$$

While the results of the study are not very sharp, they are generally supportive of the view that human capital elasticities do indeed differ across countries and are responsive to expediture variables. As can be seen in Table A3.13, the share of educational expenditure in GDP *(SEDU)* and the average number of students per teacher in primary school *(PT1)* are significant and have the expected sign when included alone in the varying-parameter specification, but the secondary pupil/teacher ratio *(PT2)* has the wrong sign and is not significant. Dessus also finds that the human capital coefficient increases with average schooling measured at the beginning of the sample period *(YS_0)*. This result may be interpreted as an indication of the importance of intergenerational externalities (children benefit from having educated parents through learning at home and greater motivation) and may generate threshold effects as argued by Azariadis and Drazen (1990).

Table A3.13: Dessus (1999)

	[1]	*[2]*	*[3]*	*[4]*
q_0	-0.444	-0.439	-0.457	-0.459
	(5.45)	(5.26)	(5.31)	(5.53)
$ln\ s_k$	0.214	0.209	0.211	0.220
	(4.62)	(4.43)	(3.84)	(4.49)
h. capital param:				
α_{YS0}	-0.175	0.714	-0.133	-0.351
	(1.57)	(3.05)	(0.45)	(0.05)
YSo	0.080			
	(2.96)			
PT1		-0,018		
		(2.76)		
PT2			0.013	
			(0.86)	
SEDU				0.111
				(2.08)

- *Notes: t* statistics below each coefficient. The varying parameter model is estimated using the specification proposed by Amemya (1978). Average years of schooling are taken from Barro and Lee (1993) and the other educational indicators from UNESCO. Notice that there is no temporal variation in the quality indicators, which are defined as averages over the sample period.

Some studies have examined the correlation between growth performance and standardized achievement measures. A paper by Lee and Lee (1995) obtains some suggestive results with data on average national scores in tests administered by the International Association for the Evaluation of Educational Achievement (IEA) in the early seventies. Using science scores as their proxy for initial human capital, these authors estimate a series of simple cross-section convergence regressions for a sample of 17 (developed and underdeveloped) countries with the results shown in Table A3.14. As usual, the dependent variable is the growth rate of GDP per worker (between 1970 and 1985) and the conditioning variables include the initial level (not log) of GDP per worker *(Q_0)* and one or several schooling indicators. It is interesting to note that the partial correlation between test scores *(SCORE)* and growth is positive and significant even when we control for alternative human capital indicators such as the primary or secondary enrollment rates (*PR.ENROL* or *SEC.ENROL*) or the average years of schooling of the adult population *(YS)* and that all these variables tend to lose their significance when *SCORE* is included as a regressor. Barro (1998, 2000) confirms Lee and Lee's findings on the significance of test scores but finds that, in some but not all specifications, years of schooling continue to be significant when both variables are entered simultaneously in the growth equation.

Table A3.14: Results of Lee and Lee (1995)

	[1]	*[2]*	*[3]*
Q_0	-0.0016 (4.00)	-0.0019 (2.11)	-0.0009 (0.64)
SCORE	0.0018 (4.50)	0.0016 (2.29)	0.0027 (4.50)
PR.ENROL		0.0008 (0.03)	
SEC.ENROL		0.0128 (0.40)	
YS			-0.0042 (1.91)
R^2	0.572	0.507	0.640

- *Note:* t statistics s below each coefficient. *YS* seems to be taken from some version of the Barro and Lee data set, but the authors do not say it explicitly.

A more thorough attempt along similar lines is due to Hanushek and Kimko (H&K 2000). These authors construct an indicator of labour force quality for a sample of 31 countries using their scores in a number of international achievement tests in mathematics and science.[19] This indicator is then included as a regressor in a growth equation with results that are qualitatively similar to those of Lee and Lee (1995). H&K, moreover, conduct extensive robustness checks and provide fairly convincing evidence that the observed correlation between test scores and growth reflects, at least in part, a causal relationship.

To approximate the average quality of the stock of workers (rather than that of current students), H&K combine all the test scores available for each country into a single cross-section indicator that is constructed as a weighted average of the standardized values of such scores. They use two alternative standardization procedures to produce two different (but highly correlated) measures of labour force quality that they denote by *QL1* and *QL2*. In the first case *(QL1)*, the average world score in each year (measured by the percentage of correct

[19] The authors use the results of six such tests that were conducted between 1965 and 1991 (four by IEA and two by IAEP (International Assessment of Educational Progress)). The countries for which direct score data are available are: Australia, Belgium, Brazil, Canada, Chile, China, Finland, France, West Germany, Hong Kong, Hungary, India, Iran, Ireland, Israel, Italy, Japan, Jordan South Korea, Luxembourg, Mozambique, Netherlands, New Zealand, Nigeria, Norway, Philippines, Poland, Portugal, Singapore, Japan, Swaziland, Sweden, Switzerland, Taiwan, Thailand, UK, US and the USSR. Some of these are excluded from the sample used in the growth equations due to lack of other relevant variables.

answers) is normalized to 50. This procedure implicitly asumes that average performance does not vary over time. In the second case, they allow average peformance to drift over time reflecting average US scores in a different but comparable set of national tests. Finally, H&K enlarge their original sample (to around 80 countries) by estimating the values of their quality measures in a number of other countries using an auxiliary equation that is estimated with the original sample. This equation links their quality indicators to the primary enrollment rate, the average years of schooling of the adult population, the share of educational expenditures in GDP, the rate of population growth and regional dummies for Asia, Latin America and Africa.[20]

Table A3.15: Hanushek and Kimko (2000)

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Qo	-0.609	-0.472	-0.460	-0.270	-0.382	-0.370	-0.393	-0.368
	(3.27)	(4.92)	(4.47)	(3.14)	(4.72)	(4.40)	(4.14)	(3.87)
YS	0.548	0.103	0.100	0.085	0.127	0.120	0.070	0.065
	(2.62)	(0.82)	(0.68)	(0.75)	(1.43)	(1.25)	(0.67)	(0.56)
QL1		0.134		0.091	0.108		0.112	
		(5.83)		(3.96)	(5.14)		(5.60)	
QL2			0.104			0.094		0.100
			(6.93)			(5.88)		(6.67)
PT1							0.001	0.006
							(0.04)	(0.25)
PT2							-0.038	-0.038
							(0.86)	(0.84)
SEDU							7.388	3.968
							(0.46)	(0.26)
R^2	0.33	0.73	0.68	0.40	0.41	0.41	0.42	0.42
N	31	31	31	25	78	80	76	78

Notes:
- *t* statistics below each coefficient. *N* is the number of observations (countries).
- The dependent variable is the average annual growth rate of real per capita GDP between 1960 and 1990. Initial income is in levels, not in logs. *YS* is average years of schooling from Barro and Lee (using a 1994 update of their 1993 paper); this variable enters the equation as the average of the quinquennial observations for each country between 1960 and 1985.
- Equation [4] excludes Hong-Kong, Korea, Singapore, Taiwan, Japan and Thailand.
- *PT1* and *PT2* are the pupil to teacher ratios in primary and secondary education and *SEDU* the share of educational expenditure in GDP.

Labour force quality indicators are then entered in cross-section growth regressions that control for initial real income per capita, Q_0, (measured in levels, not in logs) and by Barro and Lee's (1993) measure of average years of schooling of the adult population *(YS)*. As can be

[20] The estimated contribution of these variables to labour force quality is positive and significant in the cases of primary enrollments and average schooling, negative and significant for the rate of population growth, and positive but not significant for educational expenditures.

seen in Table A3.15, the quality variables display the expected positive sign, are highly significant, and tend to drive out other educational indicators, including average years of schooling, which is only significant when the quality variable is omitted. This result holds for both quality indicators in the original and in the enlarged samples (equations [2] and [3] versus [5] and [6]), and are robust to the omission of East-Asian countries (equation [4]), which might conceivably generate a spurious correlation between growth and test scores because of their excellent performance on both accounts. H&K's indicators, moreover, seem to be better measures of schooling quality than pupil to teacher ratios in primary and secondary schooling ($PT1$ and $PT2$) or the share of educational expenditure in GDP ($SEDU$). In fact, these variables are not significant in the growth equation even without controlling for test scores. Finally, the authors also report that their findings are not sensitive, qualitatively or quantitatively, to the inclusion of additional variables such as the share of government consumption in GDP, the investment ratio, a measure of openness to international trade and indices of political instability.

Hanushek and Kimko provide fairly convincing evidence that their results are not, at least in qualitative terms, driven by reverse causation or by omitted variables bias and can therefore be interpreted as evidence of a causal relationship running from the quality of education to growth. They base this conclusion on two separate pieces of evidence. The first one is their finding (discussed further below) that various measures of resource input into the school system do not seem to be positively correlated with test scores. It is conceivable, they argue, that growth may feed back into higher educational quality through increased school funding, thereby generating an upward bias in the coefficient of quality in growth regressions. But since funding seems to have no measureable effect on quality, a crucial link in the chain is broken and it is unlikely that the results are driven by reverse causation.

The second piece of evidence is obtained through the estimation of a mincerian wage equation for a sample of immigrants into the US. H&K find that the quality of schooling in the country of origin enters the equation with a positive and significant coefficient (after controlling in the usual way for years of schooling and experience) but only in the case of those workers who migrated after completing their education abroad, and not for those who completed their schooling in the US. The authors interpret this finding as an indication that their quality variables are not simply proxies for relevant country characteristics that are omitted in the growth equation, or even for cultural or family factors that may persist after migration. They note, however, that the microeconometric estimates obtained with immigrant data seem to imply much smaller productivity effects than their macroeconometric growth equation results, and that this suggests that the latter set of estimates may be picking up something more than direct productivity effects.

Can quality be purchased?

The results we have reviewed in this section suggest that educational quality may be just as important as quantity as a determinant of productivity, if not more. This raises the obvious policy question of what may be done to improve the quality of educational systems. In addition to teaching techniques and curriculum design, a plausible hyptohesis is that quality will tend to rise with educational expenditure, as more resources are likely to translate into more and better teachers and into improved facilities.

The evidence on this issue, which comes mostly from microeconomic studies, is conflictive (see for instance Hanushek (1986) and Card and Krueger (1996)). At the macroeconomic level, we are aware only of two studies that have dealt with the subject and they too reach conflicting results. As noted above, Hanushek and Kimko (H&K 2000) conclude that standard measures of school resources do not have a perceptible effect on the quality of schooling as measured by achievement tests. Lee and Barro (2001), on the other hand, find a positive correlation between test results and some expenditure variables.

The results of both studies are summarized in Table A3.16. Both sets of authors find that the average attainment of the adult population (YS) has a positive impact on school performance. This result, which is consistent with Dessus' (1999) findings, is suggestive of a strong family influence on school outcomes. In the same line, Lee and Barro (2001) also find that income per capita (q), which they interpret as a proxy for parents' income has a strong positive effect on test scores, and H&K report that test scores tend to be lower in countries with greater rates of population growth $(GPOP)$, as suggested by theoretical models emphasizing the tradeoff between the quantity and quality of children.

Turning to measures of school inputs, H&K (2000) find that the primary school pupil to teacher ratio $(PT1)$ and two measures of educational expenditure (total expenditure in education as a fraction of GDP, $SEDU$, and a measure of expenditure per student *exp/pupil*) display the "wrong" sign in the test score regression. Lee and Barro (2001), by contrast, conclude that smaller class sizes tend to be associated with better performance (i.e. obtain a significant negative coefficient for $PT1$) and detect some indications of a positive effect of primary school salaries $(W_{teacher})$ which presumably operates through the quality and motivation of the teaching staff. It is interesting to note that expenditure levels per se *(exp/pupil)* are only weakly positively correlated with performance (see equation [5]), and that this corelation disappears when we control for class size and teacher salaries (equation [6]). This suggests that expenditures that do not affect the quantity and quality of teachers are much less important for peformance than these two items. These results are generally robust to the inclusion of an Asian dummy (which turns out to be positive and highly significant, see equation [7]) and of fixed country effects (equation [8]).

Table A3.16: Hanushek and Kimko (2000) vs. Lee and Barro (2001)
dependent variable = test scores

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*	*[7]*	*[8]*
q				3.19	4.16	3.41	3.39	3.53
				(3.00)	(4.23)	(3.20)	(3.42)	(0.42)
YS	2.04	1.62	1.54	1.33	1.33	1.35	1.17	5.02
	(2.49)	(2.13)	(2.41)	(4.93)	(4.94)	(4.90)	(4.67)	(1.96)
$GPOP$	-4.65	-4.60	-2.64					
	(2.77)	(3.38)	(1.35)					
$PT1$			0.066	-0.15		-0.22	-0.19	-0.76
			(0.41)	(2.44)		(2.54)	(3.20)	(1.70)
$SEDU$		-165.9						
		(1.83)						
$exp/pupil$	-0.69				1.06	-1.34		5.86
	(3.63)				(1.46)	(1.13)		(0.86)
$W_{teacher}$				1.62		2.88	1.92	7.69
				(1.81)		(2.09)	(2.28)	(1.80)
$school\ \ day$				0.01		0.003	-0.02	
				(0.46)		(0.14)	(0.89)	
$Asia$							3.67	
							(3.71)	
N	69	67	70	214	214	214	214	197
source:	H&K	H&K	H&K	L&B	L&B	L&B	L&B	L&B

Notes:
- t statistics below each coefficient. N is the number of observations.
- YS is average years of total schooling of the adult population in H&K and average years of primary schooling in L&B; *exp/pupil* is current public expenditure per student and is measured in levels in H&K and in logs in L&B;$W_{teacher}$ is measured in logarithms and *school day* refers to its length in hours.
- Equation [8] includes country fixed effects.

f. A plausible range of parameter estimates

In this section we will try to extract from the preceding review of the literature a plausible range for the values of the parameters that describe the relationship between human capital and the level and growth rate of productivity. The coefficients of interest are two alternative measures of level effects and one measure of rate effects. The level parameters are the elasticity of output with respect to average schooling, α_{YS}, and the Mincerian "gross return" on schooling, ρ, that measures the percentage increase in output resulting from a one-year increase in average attainment. As the reader will recall (see Box 2 in the main text), ρ can be obtained by dividing α_{YS} by average attainment in years, and viceversa. The rate effects parameter is the coefficient of YS in the technical progress function, γ_{YS}, and measures the contribution of an additional year of schooling to the rate of TFP growth holding other things (and in particular the gap with the world technological frontier) constant.

The first block of Table A3.17 shows a number of selected coefficient estimates taken from the empirical literature reviewed in previous sections. The first row of the table gives the source of the estimate, the second shows the specific form in which years of schooling enters the equation,[21] the third and fourth rows display the estimated value of the "raw" regression coefficient and the associated t statistic, and the fifth row lists the source of the schooling data. To facilitate the comparability of the coefficients and their interpretation, we have selected only estimates obtained using data on average years of schooling (rather than enrollment rates). We have focused mostly on recent studies that make use of the latest available data sets and use specifications that produce "respectable" signal to noise ratios for the OECD data set. Implicitly, then, we are accepting Krueger and Lindhal's (2001) argument that failure to find significant productivity effects is most likely due to poor data, and not taking into account the negative findings of some of the studies we have reviewed.

The second block of the table shows the values of α_{YS} and ρ implied by the original coefficients when these are interpreted as capturing level effects only. In most cases, the values of these parameters are not given directly by the estimated coefficients displayed in the first block of the table but can be recovered from them using either the explicit structural model that underlies the estimated equation, or a model that generates the same reduced form specification. For instance, Jones (1996) interprets the coefficient of YS in the steady-state equation he estimates as capturing rate effects in a world with technological diffusion. We will do something like this below, but for now we interpret his coefficient within the framework of a Mincerian MRW model (which yields exactly the same steady state specification) as capturing a level effect. In the case of Barro (2000), the estimated convergence equation is not explicitly derived from a structural model, but it can be interpreted as such because the functional form is similar to the one that would be implied by the same Mincerian MRW model when we allow for transitional dynamics.[22] To recover the

[21] The notation is the standard one in this report: YS denotes years of schooling, ys the log of this variable and Δys its annual growth rate, computed as the average annual log change over the relevant period.

[22] Within this model, the coefficient of years of schooling will provide an estimate of $\beta\frac{\rho}{1-\alpha_k}$, where β (the rate of convergence) is the coefficient of log initial income per capita. Barro's equation includes both this variable and its square, but the author reports that the average rate of convergence in the sample is 2.5%. This is the value of β we use in our calculations and is shown in the last block of the table. Barro's equation controls for investment in physical capital, but the investment ratio does not enter the equation in a way that permits us to recover an estimate of α_k. Hence, we assume a value of 1/3 for this parameter.

values of α_{YS} and ρ we typically need an estimate of α_k. When possible, this is taken from the original equation (as in Jones (1996) or in Bassanini and Scarpetta (2001)); otherwise, a value of 0.333 is assumed for this parameter.

Table A3.17: Selected estimates, corrections for measurement error bias and tentative estimates of rate effects

	1. original coefficient estimates:				
source:	*D&D update*	*C&S (2001)*	*Bas&Scarp*	*Barro (2000)*	*Jones (1996)*
regressor:	Δys	YS	ys	YS*	YS
raw coefficient	0.394	0.085	0.95	0.0044	0.159
(t)	(4.57)	(4.00)	(3.96)	(2.44)	(2.48)
data from:	D&D (2001)	C&S (2001)	D&D (2001)	B&L (2000)	B&L (1993)
	2. implied values of the level parameters:				
coefficient interpreted as	α_{YS}	$\frac{\rho}{1-\alpha_k}$	$\frac{\alpha_{YS}}{1-\alpha_k}$	$\beta\frac{\rho}{1-\alpha_k}$	$\frac{\rho}{1-\alpha_k}$
implied ρ	**3.70**%	5.67%	7.76%	11.73%	11.75%
implied α_{YS}	0.394	0.603	0.826	1.248	1.250
	3. level parameters after correcting for measurement error:				
reliab. ratio	0.736	0.788	0.716	0.910	0.897
corrected coeff.	0.535	0.108	1.327	0.005	0.177
implied ρ	**5.03**%	7.19%	10.84%	12.89%	13.10%
implied α_{YS}	0.535	0.765	1.154	1.372	1.394
	4. implied value of γ_h *under the assumption that* α_{YS} *= 0.535/*ρ *= 5.03%*				
corrected coeff. interpreted as:		$\frac{\rho}{1-\alpha_k}+\frac{\gamma_{YS}}{\lambda}$	$\frac{\alpha_{YS}}{1-\alpha_k}+\frac{\gamma_{YS}}{\lambda}$	$\frac{\beta\rho}{1-\alpha_k}+\frac{\beta\gamma_{YS}}{\lambda}$	$\frac{\rho}{1-\alpha_k}+\frac{\gamma_{YS}}{\lambda}$
implied γ_{YS}	**0.00**%	0.24%	0.49%	**0.87**%	0.81%
	5. other parameter values used in the calculations:				
avge. YS	10.64	10.64	10.64	10.64	10.64
λ	0.074	0.074	0.074	0.074	0.074
α_k		0.333	0.130	0.333	0.261
β				0.025	

(*) The regressor is some transformation of the average years of total schooling of the adult population, except in Barro (2000), where it is the average years of secondary and higher schooling of the adult male population.

The calculations we have just sketched will produce an estimate of α_{YS} when the underlying production function is Cobb-Douglas in years of schooling (i.e. when we assume that $H = YS$), and an estimate of $\rho = \theta\alpha_h$ when a Mincerian specification (with $H = Exp(\theta YS)$) is adopted. To compute ρ given α_{YS}, we will divide the latter parameter by 10.64, which is the average years of schooling in 1990 in a sample of OECD countries using D&D's

(2001b) data set.[23] The reverse procedure will be used to compute α_{YS} given the value of ρ. The values of the auxiliary parameters used in these computations are shown in the last block of the table.

The third block of the table shows the effects on parameter estimates of correcting for measurement error using the appropriate reliability ratios taken from Tables A2.1 and A2.2 above. Notice that the correction is only a partial one because it ignores the increase in the attenuation bias that will result from the introduction of additional regressors when these are correlated with schooling (see Section 2b.iii of the main report). The corrected estimates of the raw coefficients are obtained by dividing their original values (in the first block of the table) by the reliability ratios shown in the first row of the third block. The implied values of α_{YS} and ρ are then recovered in the manner explained above, working with the corrected raw coefficients.

The range of parameter values obtained in this manner is very broad. Estimates of the Mincerian return to schooling in OECD countries *(ρ)* range from 3.7% (using D&D's uncorrected estimates) to 13.1% (using Jones' estimates corrected for measurement error). The higher values in this range appear extremely implausible when we interpret them as estimates of direct level effects. After correcting for measurement error, three of the five studies imply values of α_{YS} greater than one, i.e. increasing returns to schooling alone.

We interpret these findings as an indication that, as may be expected from our previous discussion about the difficulty of empirically separating level and rate effects, the coefficient estimates shown in Table A3.17 are picking up both of them. To get some feeling for the likely size of the rate effects, we will take as given the values of the level parameters implied by D&D's estimates (corrected for measurement error) and solve for the value of the rate effects coefficient, γ_{YS}, that is consistent with the raw coefficient of schooling. To do this, we will reinterpret the reported raw coefficients within the framework of an enlarged model with rate effects and technological diffusion. In this context, and under the further assumption that countries are reasonably close to their "technological steady states" relative to the world frontier, the coefficient of the schooling variables will reflect both the standard level effect and an additional term of the form γ_{YS}/λ, where λ is the rate of technological diffusion.[24] The fourth block of the table shows the results of this calculation, which uses the value of λ estimated by de la Fuente and Doménech.

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[23] Hence, the values of ρ given in Table A3.17 refer to this sample and are therefore different from those used in the rate of return calculations in the main report, which correspond to a subset of the OECD sample comprised by 14 member states of the EU.

[24] The details of the required calculations are as follows. Let x be the relevant "raw coefficient" corrected for measurement error and assume for concreteness that we are interpreting this coefficient as

4. Selected educational indicators for the EU and other countries of interest

Tables A4.1-A4.12 collect a number of human capital indicators for the EU and other countries of interest. Each table is divided into four blocks. The first one gives values for EU members (EU14), typically with the exception of Luxembourg for which data are often not available; the second refers to a group of seven advanced OECD economies that serve as a useful reference to gauge the EU's position relative to its most direct competitors; and the third block gives the data available for countries that are currently candidates for accession to the EU. The coverage for the last group varies across tables and is often restricted to countries in this group that are also members of the OECD. The fourth block of each table, finally, displays average values for different subsamples and other statistics of interest.

The values in the tables are always given in relative terms, taking as a reference the average value of each variable taken over the available observations for the group of 21 OECD countries that are listed in the first two blocks. This average value, which is denoted by *avge. OECD21* or *avge. 21* in the tables, is normalized to 100. The original variables can be recovered by multiplying the average value for the reference group (which is listed under *avge. OECD in levels*) by the relative values given in the table.

Tables A4.1-A4.4 contain various measures of the educational attainment of the adult population (i.e. of the quantity of human capital) in selected years between 1960 and 2000. As has been emphasized elsewhere in this report, there are significant discrepancies across sources that introduce a considerable amount of uncertainty in cross-country comparisons. Nonetheless, it seems clear that Southern European countries (Greece, Italy, Spain and especially Portugal) have significantly lower attainment levels than the rest of the Union, and that the Nordic countries and Germany occupy the first positions of the EU ranking by

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$$x = \frac{\rho}{1-\alpha_k} + \frac{\gamma_{YS}}{\lambda}$$

Given the assumed values of λ, ρ and α_k, we can solve for γ_{YS} as

$$\gamma_{YS} = \lambda\left(x - \frac{\rho}{1-\alpha_k}\right) .$$

In the case of Basanini and Scarpetta (2001), an additional step is necessary. Since these authors use years of schooling in logs rather than in levels (ie. $ys = ln\ YS$), the calculation just described will yield an estimate of the change in the rate of technical progress *(g)* induced by a unit increase in log schooling, i.e. of $\frac{\partial g}{\partial lnYS}$ rather than of γ_{YS} which is defined as $\frac{\partial g}{\partial YS}$. To recover the parameter of interest, notice that

$$\frac{\partial g}{\partial YS} = \frac{\partial g}{\partial lnYS}\frac{dlnYS}{dYS} = \frac{\partial g}{\partial lnYS}\frac{1}{YS}$$

so we have to divide the result of the first calculation by average years of schooling to recover γ_{YS}.

Table A4.1: Average years of schooling *(YS)* in 1960 and 1990

source:	D&D01	C&S	B&L00	average	D&D01	C&S	B&L00	average
year:	1960	1960	1960	1960	1990	1990	1990	1990
W. Germany	118.5	117.9	123.6	120.0	121.7	120.9	102.1	114.9
Denmark	129.0	112.5	133.6	125.0	110.2	105.6	114.2	110.0
Sweden	96.2	107.5	114.2	106.0	99.8	110.2	107.9	105.9
UK	102.5	112.9	114.5	110.0	98.9	112.4	98.5	103.2
Finland	91.5	84.9	80.2	85.5	103.1	98.2	106.8	102.7
Austria	107.7	102.6	100.2	103.5	106.3	100.1	92.6	99.7
Netherlands	97.0	103.3	78.7	93.0	102.9	98.1	97.0	99.3
Belgium	92.5	91.6	111.4	98.5	94.7	91.8	95.0	93.8
France	97.3	83.4	86.3	89.0	98.2	94.8	85.2	92.7
Ireland	88.0	89.8	96.3	91.4	88.4	87.2	95.8	90.5
Greece	66.5	73.6	69.3	69.8	74.3	79.7	86.3	80.1
Italy	64.7	72.1	68.1	68.3	75.6	83.3	69.4	76.1
Spain	59.5	71.7	54.3	61.8	66.7	77.2	68.6	70.9
Portugal	52.3	39.0	29.0	40.1	60.2	54.1	48.8	54.4
USA	126.3	126.1	129.3	127.2	119.1	115.5	135.2	123.3
Australia	117.7	121.7	140.8	126.7	121.1	116.8	114.1	117.3
Canada	124.1	112.9	125.0	120.6	119.7	113.1	118.3	117.1
Switzerland	124.8	135.8	109.0	123.2	114.9	118.6	111.8	115.1
N. Zealand	125.1	111.3	142.7	126.4	113.8	100.8	126.0	113.6
Norway	115.8	112.1	91.2	106.4	104.4	112.7	122.3	113.1
Japan	103.1	117.4	102.6	107.7	105.6	109.2	103.9	106.2
Poland			100.62				108.2	
Latvia							107.52	
Czech Rep.							105.83	
Lithuania							104.81	
Bulgaria		90.44	90.77			96.9	104.36	
Romania		89.45	79.57			91.5	104.14	
Estonia							103.35	
Slovakia							102.22	
Hungary		93.78	99.27			99.46	98.16	
Cyprus		68.51	64.04			81.16	94.78	
Slovenia							78.22	
Malta			84.2				76.3	
Turkey		26.51	29.86			57.19	44.52	
Czechosl.			107.34					
E. Germany			131.37				114.62	
avge. 21	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
in years	8.36	8.07	6.70		10.64	10.93	8.87	
avge EU 14	90.22	90.20	89.97	90.13	92.95	93.81	90.60	92.45
avge. cands.		73.74	81.96			85.24	94.80	

Notes:
- *Sources:* D&D01 = de la Fuente and Doménech (2001); C&S = Cohen and Soto (2001); and B&L00 = Barro and Lee (2000).
- The data refer to the population 25 and over in D&D and B&L, and to the population between 15 and 64 in C&S.
- The average values given in the fourth and eigth columns of the Table are simple averages across sources of the normalized values for each year.
- The average for the candidate countries (*avge. cands.*) does not include East Germany.

Table A4.2: Average years of schooling *(YS)* in 2000
and expected years of schooling based on current enrollment rates

	YS (C&S 2001)	*YS (B&L 2000)*	*YS expected (OECD)*	*YS expected (WB)*
Sweden	101.75	119.13	117.00	93.21
Finland	101.40	106.33	105.48	102.86
Denmark	105.92	105.81	102.02	96.43
*Germany***	112.43	102.24	99.14	102.86
UK	113.90	98.05	108.93	106.07
Netherlands	98.45	96.89	98.56	102.86
Ireland	88.29	94.59	92.22	90.00
Austria	99.23	92.28	92.22	93.21
Belgium	94.11	91.55	106.63	109.29
Greece	85.95	89.24	89.91	90.00
France	93.15	87.77	95.10	99.64
Spain	82.48	76.03	99.71	
Italy	89.68	73.40	91.07	
Portugal	63.20	51.49	96.83	93.21
USA	109.65	128.46	99.14	102.86
Norway	108.35	124.37	103.17	99.64
N. Zealand	104.96	120.80	99.14	106.07
Canada	113.47	119.86	95.10	109.29
Australia	113.64	110.84	114.70	109.29
Switzerland	110.52	108.95	93.95	93.21
Japan	109.48	101.93		
Poland		103.82	92.22	83.57
Bulgaria	91.94	102.14		77.14
*Latvia**		100.04		80.36
Romania	86.82	99.73		77.14
Czech Rep.		99.20	87.03	83.57
*Lithuania**		97.52		
Slovakia		96.37		
*Estonia**		96.16		80.36
Hungary	94.37	92.38	92.22	83.57
Cyprus	77.01	91.97		
Malta		79.38		
Slovenia		77.07		
Turkey	54.26	50.33	61.10	64.29
avge. OECD21	100.00	100.00	100.00	100.00
in levels	11.52	9.54	17.35	15.56
avge EU14	*95.00*	*91.77*	*99.63*	*98.30*
avge cands.	*80.88*	*91.24*	*83.14*	*78.75*

Notes:
- Average years of schooling of the adult population *(YS)* in 2000 from Barro and Lee (2000) and Cohen and Soto (2001), and expected years of schooling in the future on the basis of currently observed enrollment rates as calculated by the World Bank (WB) for the *2000/2001 World Development Report* and by the OECD in the 2001 edition of *Education at a Glance*. Both estimates of "school expectancy" are constructed essentially by adding up across successive school grades (excluding pre-primary education) the enrollment rates observed in the late 1990s.
(*) In the case of the Baltic countries the Barro and Lee data refers to 1990.
(**) In Cohen and Soto, Germany is West Germany.

school attainment. In terms of average years of schooling *(YS)* the average EU country was around ten percentage points below the OECD average in 1960. This educational gap had been reduced by only 2.5 percentage points by 1990 and perhaps by one additional point by 2000. (See Tables A4.1 and A4.2). Projections based on current enrollment rates, however, suggest that EU attainment will gradually converge to the OECD average in the future provided current conditions remain unchanged (see the last two columns of Table A4.2).

Table A4.3: Standard deviation of normalized years of schooling in the EU14

	D&D	*C&S*	*B&L*	*average*	*WB*	*OECD*
1960	21.55	20.78	27.98	22.73		
1990	17.00	16.42	17.17	16.16		
2000		12.60	15.97			
future					6.22	7.49
%Δ 1960-90	-21.1%	-21.0%	-38.6%	-28.9%		
%Δ 1990-00		-23.2%	-7.0%			

- Standard deviation of normalized years of schooling and expected future years of schooling from Tables A4.1 and A4.2

The dispersion of relative national attainment levels within the EU has declined significantly during the period we are considering and can be expected to continue to do so in the foreseeable future. This convergence process is illustrated in Table A4.3, which shows the standard deviation of normalized years of schooling according to various sources in selected years and the same dispersion indicator for the OECD and World Bank projections based on current enrollment rates. Between 1960 and 1990, this indicator of educational inequality fell by almost 30% (when we work with an average across the three available sources of data on years of schooling). The decade of the 1990s saw a further reduction of this variable, although the two available sources imply very different convergence rates, and existing projections suggest that, under current conditions, the level of educational inequality within the EU should fall to about half its current level within a generation. Educational convergence is also apparent in the data on upper secondary attainment by age group contained in Table A4.4: the dispersion of the normalized values of this indicator is 50% lower in the 25-34 age group than in the 55-64 age group.

The attainment data available for the countries that are currently candidates for accession to the EU (henceforth the candidate countries) is limited and somewhat hard to assess. Barro and Lee (2000) is the only data set that provides estimates of average years of

Table A4.4: Upper secondary attainment by age group in 1999

	Ages 25-64	*Ages 25-34*	*Ages 35-44*	*Ages 45-54*	*Ages 55-64*
Germany	124.19	112.30	121.62	132.84	147.43
Denmark	121.72	114.98	114.26	128.62	141.19
Sweden	117.10	114.74	116.56	120.24	123.86
Austria	112.89	109.65	111.78	113.15	118.53
Finland	109.32	112.83	117.26	108.80	93.87
France	94.59	100.68	93.48	92.23	84.98
UK	94.32	86.89	91.11	97.40	107.96
Belgium	87.71	96.23	87.52	81.58	72.32
Ireland	78.43	87.93	80.82	66.24	62.37
Greece	76.34	93.81	83.47	69.27	49.51
Italy	64.57	73.10	71.40	59.95	43.05
Spain	53.70	71.87	58.56	40.65	27.26
Portugal	32.48	40.18	30.45	25.10	22.79
Netherlands					
US	132.87	115.67	126.32	144.14	164.28
Norway	129.32	123.81	127.73	128.65	138.47
Switzerland	124.93	117.01	120.27	129.07	145.40
Japan	123.72	122.67	132.21	128.12	120.44
Canada	121.52	114.89	119.15	127.66	125.97
N. Zealand	112.55	104.53	111.13	116.47	121.29
Australia	87.73	86.23	84.89	89.81	89.04
Czech Rep.	131.53	121.98	128.08	137.88	151.73
Hungary	103.01	105.00	109.33	114.36	73.83
Poland	82.54	81.73	84.65	86.48	74.89
Turkey	33.93	34.54	32.71	29.86	23.88
avge. OECD21	100.00	100.00	100.00	100.00	100.00
in levels	65.42	75.86	69.70	61.29	49.41
avge. EU14	*89.80*	*93.48*	*90.64*	*87.39*	*84.24*
avge. cands.	73.16	73.76	75.57	76.90	57.53
SD OECD21	27.10	20.45	25.79	32.66	41.08
SD EU14	*26.95*	*21.05*	*25.62*	*32.48*	*40.37*

Notes:
- *Definition:* percentage of the population that has attained at least upper secondary education by age group.
- *SD* is the standard deviation of attainment and is computed with the normalized data.
- The data refer to 1998 in the cases of Austria, Ireland, Norway and Poland.
- *Source: Education at a Glance*, 2001.

schooling for all the countries in this group. According to these authors, average attainment in the candidate countries (*avge. cands.* in Tables A4.1 and A4.2) was four percentage points above the EU14 level in 1990 and approximately equal to the EU14 average in 2000, with all candidate countries except Malta, Slovenia and Turkey above mean EU14 attainment in this year. These figures, however, may significantly overestimate attainment levels in the candidate countries. Cohen and Soto's estimates of schooling in 2000 are significantly lower than Barro and Lee's for two out of the three formerly socialist countries for which these authors supply data. A recent study for the European Commission (EIC, 2001) also suggests that attainment statistics tend to overstate the human capital stocks of Eastern European countries because a large share of secondary-level qualifications were obtained in vocational schools that typically offered short courses with deficient curricula. The rapid decline in attendace to these schools may be partly responsible for the apparent fall in (relative) enrollment rates that these countries seem to have experienced over the recent period of turmoil caused by the crisis and eventual demise of their communist regimes. This decline is apparent in Table A4.4, where we see that secondary attainment rates in socialist countries decline as we move to younger cohorts, following the opposite pattern than the rest of the sample, and in Table A4.2, where projections based on current enrollment rates suggest that the relative attainment levels of formerly socialist countries are likely to deteriorate rapidly in the future.

Tables A4.5-A4.8 contain various indicators of educational expenditure and school resource input in recent years. The source for Tables A4.5 and A4.6 is the 2001 edition of the OECD's *Education at a Glance*, which provides information for our OECD sample and for four candidate countries (Hungary, Poland, Turkey and the Czech Republic). Some additional information for other candidate countries is provided in Tables A4.7 and A4.8, which are taken respectively from the World Bank and from Barro and Lee's (2000) data set. As in the case of attainment estimates, there are worrisome discrepancies across data sources that make it necessary to interpret international comparisons with great caution.[25] On the whole, however, the OECD data (Tables A4.5 and A4.6) suggest that the EU is only slightly below the OECD average in terms of most indicators of educational expenditure and slightly above this average in terms of direct measures of school input (teachers per pupil and hours of instruction per year). The exceptions to this rule occur at the tertiary level, where both

[25] Many of these indicators are not strictly comparable across sources, but they should in principle capture similar things. The coefficient of correlation among similar OECD and World Bank indicators, computed over common observations and normalized by the OECD21 average, is as follows: public expenditure in education as a percentage of some measure of national income (0.322), expenditure per student as a percentage of income per capita (0.756 at the primary level, 0.457 at the secondary level and 0.452 at the tertiary level) and pupil to teacher ratio at the primary level (0.888).

Table A4.5: Indicators of expenditure in education in 1998 (OECD)

	GEDUps/GDPpc			*GEDUps/GDPpew*			*GEDU/GDP*		
	primary	*second.*	*tertiary*	*primary*	*second.*	*tertiary*	*public*	*private*	*total*
Austria	130.13	130.90	112.71	142.86	144.24	121.80	115.84	62.88	112.34
Sweden	129.22	97.77	142.65	129.97	98.70	141.23	96.30		87.75
Denmark	132.76	106.43	88.07	150.41	121.02	98.22	131.93	59.17	126.62
Portugal	106.93	110.56		107.62	111.68		111.32		101.06
Italy	129.06	110.21	66.94	101.52	87.02	51.83	113.83	59.50	110.14
France	89.82	118.19	80.58	78.22	103.31	69.08	84.30	197.41	97.98
UK	79.93	93.86	108.46	82.45	97.18	110.14	66.65	217.58	84.04
Germany	78.01	102.51	97.55	79.89	105.38	98.34	83.39	66.66	83.16
Finland	107.82	88.74	79.27	103.63	85.61	75.00	93.30	31.86	88.47
Spain	97.08	102.75	69.72	72.87	77.41	51.51	86.90	19.45	81.31
Netherlands	77.81	81.28	102.71	73.31	76.86	95.26	117.10		
Belgium	80.95	94.54	63.63	68.45	80.24	52.96	107.85	13.59	99.79
Greece	83.63	86.77	68.37	69.44	72.32	55.89	86.02	140.79	93.49
Ireland	61.18	65.54	88.46	56.19	60.42	79.98	127.57	30.05	119.53
N. Zealand							89.99	45.71	86.94
Canada			136.32			140.37	83.95	186.02	96.44
Switzerland	119.74	129.30	142.77	145.13	157.30	170.34	106.20	111.44	108.75
US	94.77	91.01	144.64	103.15	99.42	154.97	68.74	192.78	83.33
Norway	111.48	106.21	98.40	126.62	121.08	110.01	131.17	21.14	121.85
Japan	106.54	92.42	96.51	122.80	106.92	109.49	104.24	77.98	103.39
Australia	83.14	91.00	112.24	85.48	93.90	113.59	93.40	265.98	113.61
Hungary	98.69	77.31	124.17				78.87	98.20	82.41
Poland	92.52	66.46	122.74				86.32	96.62	89.04
Czech Rep.	64.31	93.00	101.71				103.56		
Turkey							56.91	89.19	61.43
avge. 21	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00
in levels	19.76	26.44	42.44	8.83	11.77	19.26	5.16	0.61	5.66
avge. EU14	*98.88*	*99.29*	*89.93*	*94.06*	*94.38*	*84.71*	*100.82*	*78.72*	*98.04*
avge. cands.	85.17	78.92	116.21				81.41	94.67	77.63
max/min EU	*2.17*	*2.00*	*2.24*	*2.68*	*2.39*	*2.74*	*1.98*	*16.01*	*1.56*

Definitions:
- *GEDUps/GDPpc* = expenditure per student relative to GDP per capita (expenditure on public and private educational institutions per student, measured as full time equivalent)
- *GEDUps/GDPpew* = expenditure per student relative to GDP per employed worker. It is obtained by multiplying the previous variable by the ratio of employment to the total population using data for 1998 from an updated version of Doménech and Boscá (1996).
- *GEDU/GDP* = direct and indirect expenditure on educational institutions from public and private sources as a fraction of GDP.

Notes:
- Countries are ranked within each group by the average value of all the normalized indicators shown in the Table.
- For expenditure per student as a fraction of GDP per capita or per employed worker, the data refer to public institutions only in Austria, Hungary, Italy, Norway and Portugal; and to public and government-dependent private institutions only in Belgium and Greece.
- For GEDU / GDP, public subsides to households are included in private rather than public expenditure in Austria, Greece, Norway, New Zealand and Poland.
- *Source: Education at a Glance*, 2001.

Table A4.6: Other indicators of school resource input in 1999 (OECD)

	pupil to teacher ratios				hours of instruction per year			
	primary	*second.*	*tertiary*	*average*	*age 12*	*age 13*	*age 14*	*average*
Denmark	63.23	90.67		76.95	93.62	94.43	96.76	94.94
Sweden	79.58	106.27	59.12	81.66	82.54	77.70	77.05	79.10
Austria	86.57	71.80	93.32	83.90	111.68	121.32	129.92	120.97
Belgium (Fl.)	83.34	64.48	113.09	86.97		100.73	99.88	100.30
Spain	92.52	94.28	102.56	96.45	88.49	91.28	90.52	90.10
Italy	67.89	75.15	154.55	99.20	123.15	115.94	114.97	118.02
Netherlands	99.67	129.69	74.81	101.39	118.88	111.92	110.98	113.93
Finland	104.36	99.28		101.82	76.23	89.71	88.95	84.97
Germany	126.02	111.37	76.60	104.66	96.29	96.63	95.82	96.25
France	117.54	93.87	105.17	105.53	93.73	102.72	101.86	99.44
Greece	81.08	78.03	162.25	107.12	115.46	108.70	107.79	110.65
Ireland	129.57	107.23	107.72	114.84	104.23	98.12	97.30	99.88
UK	134.47	107.71	115.22	119.13	108.11	101.77	100.92	103.60
Portugal					103.65	97.58	96.76	99.33
Norway	75.47		83.39	79.43	85.76	89.71	88.95	88.14
Switzerland	96.16	90.12		93.14				
Australia	103.73	92.89		98.31		106.28	105.80	106.04
US	97.34	114.50	87.53	99.79			101.96	101.96
Japan	126.68	113.02	72.04	103.91	97.52	91.81	91.04	93.45
N. Zealand	122.92	117.87	92.62	111.14	100.63	103.66	102.79	102.36
Canada	111.85	141.77		126.81				
Hungary	65.22	77.50	75.73	72.82	86.97	94.63	93.83	91.81
Slovak Rep.	117.20	99.90	64.06	93.72				
Czech Rep.	140.22	107.77	92.92	113.64	88.92	86.81	92.23	89.32
Turkey	179.73	117.62	133.89	143.75	96.29	90.65	89.89	92.28
avge. 21	100.00	100.00	100.00		100.00	100.00	100.00	
in levels	16.70	13.65	16.03		897.25	953.09	961.16	
avge. EU14	*97.37*	*94.60*	*105.86*		*101.24*	*100.61*	*100.68*	
avge. cands.	125.59	100.70	91.65		90.73	90.70	91.99	
max/min EU	*2.13*	*2.01*	*2.74*	*1.55*	*1.62*	*1.56*	*1.69*	*1.53*

Definitions:
- *Pupil to teacher ratio* = ratio of students to teaching staff in public and private institutions, calculations based on full-time equivalents.
hours of instruction per year = total intended instruction time in hours per school year for students aged 12 to 14.

Notes:
- The data for Belgium refers to the Flanders region.
- In the case of hours of instruction per year, the value shown in the Table for the UK is estimated as the (unweighted) average of the values for England and Scotland.
- Countries sorted by the average pupil to teacher ratio
- *Source: Education at a Glance*, 2001.

Table A4.7: Indicators of educational expenditure and school input in 1997 (World Bank)

	public GEDU/GNP	GEDUPps/GNIpc primary	GEDUPps/GNIpc secondary	GEDUPps/GNIpc tertiary	pupil/teacher primary
Denmark	141.99	136.09	148.14	141.23	67.09
Italy	85.89	114.52	122.37	60.36	73.80
Austria	94.66	111.44	106.06	100.51	80.51
Belgium	54.34	45.19	57.54	49.83	80.51
Portugal	101.67	99.12	92.32	69.76	80.51
Sweden	145.49	151.50	146.42	206.15	80.51
Greece	54.34		64.84	63.50	93.93
Netherlands	89.40	76.01	91.03	134.68	93.93
Spain	87.65	86.28	96.61	50.68	100.64
Germany	84.14			107.63	114.06
Finland	131.47	117.09	118.08	129.84	120.77
France	105.18	81.14	115.07	79.73	127.48
UK	92.90	91.41	88.02	115.89	127.48
Ireland	105.18	70.36	94.03	103.36	147.60
Norway	129.72	157.15	76.43	132.97	46.96
Switzerland	94.66	99.12	124.52	129.27	80.51
Canada	120.95				107.35
US	94.66	98.09	102.62	70.33	107.35
Australia	94.66	76.52	72.14	84.57	120.77
N. Zealand	127.96	91.93	102.19	130.13	120.77
Japan	63.11	97.06	81.58	39.58	127.48
Hungary	80.63	94.49	78.15	89.41	80.51
Latvia	110.43		219.41	93.39	87.22
Slovenia	99.92	103.22	32.20	106.78	93.93
Poland	131.47	90.39	73.42	77.45	100.64
Lithuania	94.66		119.37	119.87	107.35
Bulgaria	56.09	157.66		49.54	114.06
Estonia	126.21		194.94	109.34	114.06
Czech Rep.	89.40	84.22	92.32	99.37	120.77
Romania	63.11	104.25	37.79	90.55	134.19
Slovak Rep.	87.65	114.52		87.70	134.19
Turkey	38.56	46.22	39.50	145.50	161.02
avge. OECD21	100.00	100.00	100.00	100.00	100.00
in levels	5.70	19.47	23.29	35.12	14.90
avge. EU14	*98.16*	*98.35*	*103.12*	*100.94*	*99.20*
avge. cands.	88.92	99.37	98.57	97.17	113.45

Notes:
- *public GEDU/GNP* = public educational expenditure as a percentage of GNP in 1997, from the *World Development Report 2000/2001*.
- *GEDUPps/GNIpc* = public expenditure per student as a fraction of gross national income per capita. This variable and the primary school pupil to teacher ratio are taken from the World Bank's *2001 World Development indicators*.

Table A4.8: Indicators of educational expenditure and school input in 1990 (Barro and Lee, 2000)

	pupil/teacher ratios		*GEDUPps/GDPpc*		*hours per year*
	primary	*secondary*	*primary*	*secondary*	*primary*
Sweden	39.02	72.53	241.58	92.65	123.29
Belgium	62.30	56.76	81.38	132.70	
Austria	67.97	66.22	93.58	144.76	98.63
Denmark	72.37	84.36	168.35	114.36	106.85
Italy	72.37	68.59	78.83	106.16	83.84
France	76.78	98.55	61.03	100.37	99.86
Portugal	88.73	81.99	78.83	87.82	100.68
Finland	90.62		103.25	128.84	89.79
Netherlands	108.24	123.78	64.59	109.05	102.74
West Germany	112.02	109.59	61.03	88.30	78.08
Greece	123.34	120.62	47.81	70.93	92.47
UK	124.60	105.64	75.78	125.94	97.60
Spain	138.45	131.66	61.03	66.59	105.31
Ireland	168.03	121.41	62.05	97.95	
Norway	38.39	70.17	180.04	89.27	
Canada	96.28	110.37	114.43	119.19	100.17
Australia	103.84	97.76	100.19	57.90	
New Zealand	113.28	135.60	83.92	69.97	102.74
US	116.42	112.74	76.80	111.47	117.95
Japan	130.27	131.66	82.90	83.00	
Switzerland	156.70		182.59	102.78	
Hungary	78.66	94.61	106.81	127.87	
Bulgaria	96.91	112.74		195.81	88.15
Poland	102.58	143.49			97.60
U.S.S.R.	106.98				
Romania	107.61	221.54			76.85
Czechoslovakia	120.83	78.84			84.76
Cyprus	129.01	93.82	65.61	9.65	86.30
Malta	130.27	95.39			112.19
Yugoslavia	142.85	130.87			69.35
Turkey	191.31	188.42	63.07	54.53	89.90
avge. OECD21	100.00	100.00	100.00	100.00	100.00
in levels	15.89	12.68	19.66	20.72	973.33
avge. EU14	96.06	95.52	91.37	104.74	98.26
avge. cand.	120.70	128.86	78.49	96.97	88.14

- *GEDUPps/GDPpc* = public expenditure per student as a fraction of GDP per capita.

normalized expenditure and the number of teachers per student (the inverse of the pupil to techer ratio shown in Table A4.6) are significantly below the OECD average. The other peculiarity of the EU is that private expenditure in education is generally lower than in the rest of the OECD sample. A comparison of the OECD data with the World Bank's expenditure indicators (which consider only public spending) suggests that the relatively low level of EU spending per student at the tertiary level is due mostly to low private expenditures (i.e. to low tuition fees at universities).

Within the EU, there are very significant differences across countries in terms of the various resource indicators. The ratio between the highest and the lowest value of each indicator within this sample *(max/min EU)*, which is given in the last row of Tables A4.5 and A4.6, is always above 1.5 and often above 2.0. If we measure expenditure per student as a fraction of output per employed worker (which is probably a better reference than GDP per capita as a way to correct expenditure for differences in purchasing power), Austria, Sweden and Denmark have the highest expenditure levels and Ireland the lowest. At the tertiary level, expenditure so normalized is particularly low in Spain, Italy, Belgium and Greece. Pupil to teacher ratios and hours of instruction vary considerably less than expenditures per student, but even here the differences across countries are quite significant.

The available information suggests that the candidate countries as a whole spend less on education than the EU, both as a fraction of national income and on a per student basis. There are, however, large differences across countries in this group and discrepancies across sources that make it difficult to be very precise concerning expenditure patterns in candidate countries. On the whole, it seems clear that expenditure levels are particularly low in Turkey, Romania and Bulgaria, whereas Hungary, the Baltic Republics, Slovenia, Poland and the Czech Republic are not far from EU levels at least in terms of some resource indicators.

Tables A4.9-A4.11 display various indicators of school achievement and labour force quality based on international standardized tests or literacy surveys. Tables A4.9 and A4.10 summarize the results of two recent OECD studies (IALS and PISA) that have already been discussed in Section 2.b of this Appendix. We construct summary measures of national performance in each of these studies by averaging each country's results across the various dimensions of literacy analyzed in these surveys. These two tables display the mean national scores, the scores corresponding to the 5th and 95th percentile of each national distribution, the range of scores (defined as the difference between the previous two values) and the percentage of the adult population (in IALS) or the student population (in PISA) that falls below the literacy level (level 3) that is considered necessary for coping with the demands of work and everyday life in advanced societies. Table A4.11 shows average scores by subject in the PISA study and in another recent international study of achievement in math and science (TIMSS). As usual, all variables are normalized by their average values in our primary sample of 21 OECD countries.

Table A4.9: Average IALS results

	mean score	*5th percentile*	*95th percentile*	*range of scores*	*% in levels 1 or 2*
Sweden	109.88	127.81	107.56	89.63	59.31
Denmark	105.09	127.01	99.02	74.24	76.59
Finland	103.68	115.76	100.45	86.89	86.01
Germany	103.47	124.56	100.27	78.77	90.09
Netherlands	103.23	119.32	98.82	80.66	84.98
*Belgium ***	100.44	95.84	100.37	104.38	95.03
UK	96.35	85.39	100.02	112.98	117.42
Ireland	95.05	88.52	98.51	107.35	123.79
Portugal	81.67	57.96	89.63	117.66	172.65
Norway	106.30	122.59	101.02	81.93	70.16
Canada	100.99	87.03	103.58	118.22	98.24
Australia	99.11	86.91	99.75	111.10	100.77
US	98.46	79.65	103.35	124.32	108.40
*Switzerland**	98.40	88.96	97.47	105.00	104.65
New Zealand	97.89	92.68	100.20	106.86	111.91
Chzech Rep.	103.55	115.61	101.55	89.11	91.01
Hungary	92.97	93.79	95.10	96.26	142.28
Slovenia	85.40	63.93	91.95	116.74	160.88
Poland	83.24	58.51	91.51	120.71	167.43
average OECD21	100.00	100.00	100.00	100.00	100.00
in levels	277.28	169.08	360.12	191.04	43.54
average EU14	*99.87*	*104.69*	*99.40*	*94.73*	*100.65*
average candidates	91.29	82.96	95.03	105.71	140.40
SD all	*7.53*	*22.19*	*4.25*		

Notes:
- The figures shown in the Table are averages of the values corresponding to the three types of literacy assessed in the study (prose, document and quantitative) with weights 0.25, 0.25 and 0.50 respectively.
- For each country we show the overall *mean score* and the scores at the *5th and 95th percentiles* of the national distribution. *Range of scores* is the difference between the 95th and 5th percentile scores, and *% in levels 1 or 2* refers to the fraction of the population which is classified below level 3, which is considered the minimum required for satisfactory performance in everyday situations.
(*) For Switzerland, we report the unweighted average of the values for the German, French and Italian-speaking populations.
(**) Belgian data refer only to the Flanders region.
- *Source:* OECD and Statistics Canada (2000).

Table A4.10: Average PISA results

	mean score	*5th percentile*	*95th percentile*	*range of scores*	*% in levels 1 or 2*
Finland	106.83	115.42	103.03	89.64	57.22
UK	103.92	106.09	104.06	101.87	89.92
Ireland	102.10	105.65	100.32	94.56	79.02
Sweden	101.31	103.74	100.43	96.86	87.19
Austria	101.01	102.85	99.82	96.55	98.09
Belgium	100.12	90.36	100.97	112.43	98.09
France	99.98	101.46	99.56	97.50	100.82
Denmark	98.10	97.56	98.60	99.72	111.72
Spain	96.33	98.74	95.55	92.10	114.44
Germany	95.93	87.65	99.14	111.55	122.62
Italy	94.16	93.89	94.90	95.99	122.62
Greece	91.54	85.96	94.75	104.25	138.96
Portugal	91.39	89.19	93.11	97.34	141.69
Netherlands					
Japan	106.09	112.04	102.11	91.39	76.29
Canada	105.06	111.08	103.03	94.32	73.57
New Zealand	104.71	102.49	105.24	108.22	84.47
Australia	104.41	106.97	103.94	100.67	84.47
Switzerland	99.28	96.76	100.74	105.04	111.72
Norway	99.09	96.83	99.67	102.74	100.82
US	98.64	95.29	101.04	107.26	106.27
Czech Republic	98.30	97.71	98.98	100.36	114.44
Hungary	95.88	95.14	97.65	100.36	130.79
Poland	94.25	90.36	96.66	103.45	130.79
Latvia	90.70	84.71	94.59	105.28	155.31
avge. OECD21	100.00	100.00	100.00	100.00	100.00
in levels	506.88	340.29	655.16	314.88	36.70
avge. EU14	*98.67*	*98.35*	*98.79*	*99.26*	*104.80*
avge. candidates	94.78	91.98	96.97	102.36	132.83
SD all	*4.62*	*8.20*	*3.23*		

Notes:
- See the notes to the previous table.
- The figures shown in the Table are averages of the values corresponding to the three types of literacy assessed in the study (math, science and reading) with weights 0.25, 0.25 and 0.50 respectively.
- The *% in levels 1 or 2* refers to the reading literacy scale, which is the only one for which this information is supplied.
- *Source:* OECD (2001a).

Table A4.11: PISA and TIMSS results by subject

	PISA reading	*PISA math*	*PISA science*	*PISA avge.*	*TIMSS math*	*TIMSS science*	*TIMSS avge.*
Finland	107.57	105.63	106.57	106.83			
UK	103.03	104.25	105.38	103.92	96.68	101.64	99.17
Ireland	103.82	99.12	101.61	102.10	101.50	102.30	101.90
Sweden	101.65	100.50	101.42	101.31	99.95	101.73	100.85
Austria	99.88	101.49	102.80	101.01	103.81	106.10	104.96
Belgium	99.88	102.47	98.25	100.12	105.06	97.07	101.04
France	99.49	101.88	99.04	99.98	103.61	94.70	99.13
Denmark	97.91	101.29	95.28	98.10	96.68	90.89	93.77
Spain	97.12	93.80	97.26	96.33	93.79	98.31	96.06
Germany	95.35	96.56	96.46	95.93	98.03	100.97	99.51
Italy	95.94	90.06	94.68	94.16			
Greece	93.38	88.09	91.31	91.54	93.21	94.51	93.86
Portugal	92.59	89.47	90.92	91.39	87.44	91.27	89.37
Netherlands					104.19	106.49	105.35
Japan	102.84	109.76	108.94	106.09	116.52	108.58	112.52
Canada	105.20	105.03	104.78	105.06	101.50	100.97	101.23
N. Zealand	104.22	105.82	104.59	104.71	97.84	99.83	98.84
Australia	104.02	105.03	104.59	104.41	102.07	103.63	102.86
Switzerland	97.32	104.25	98.25	99.28	104.96	99.26	102.09
Norway	99.49	98.33	99.04	99.09	96.87	100.21	98.55
US	99.29	97.15	98.84	98.64	96.30	101.54	98.93
Czech Rep.	96.93	98.14	101.22	98.30	108.62	109.15	108.89
Bulgaria					104.00	107.44	105.73
Slovenia					104.19	106.49	105.35
Hungary	94.56	96.17	98.25	95.88	103.42	105.34	104.39
Slovakia					105.35	103.44	104.39
Latvia	90.23	91.24	91.12	90.70	94.95	92.22	93.58
Romania					92.83	92.41	92.62
Lithuania					91.87	90.51	91.18
Cyprus					91.29	88.04	89.65
Poland	94.37	92.62	95.67	94.25			
avge. OECD21	100.00	100.00	100.00	100.00	100.00	100.00	100.00
in levels	507.60	507.45	504.85	506.88	519.24	525.89	522.57
avge. EU14	*99.05*	*98.05*	*98.54*	*98.67*	*98.66*	*98.83*	*98.75*
avge. cands	94.02	94.54	96.56	94.78	99.61	99.45	99.53

Notes
- TIMSS: scores are for 13-year olds. In the cases of the UK and Belgium, reported figures are based on the unweighted average of mean regional scores (for England and Scotland and Flanders and Wallonia respectively).
- Countries ranked by average PISA scores (with weights 0.5 for reading and 0.25 for mathematics and science), except candidate countries, which are ranked by their average TIMSS score (with equal weights for mathematics and science).
- Sources: OECD (2001a) and *The Economist* (1997).

Table A4.12: Communications and technology indicators

	telephone lines	*mobile phones*	*personal computers*	*internet hosts*	*R&D personnel*
Denmark	120.26	127.91	133.46	130.31	127.14
Sweden	122.81	163.05	127.80	138.36	149.26
Finland	100.95	201.00	123.49	251.31	109.19
Netherlands	108.05	74.85	112.31	106.64	86.57
Germany	103.31	59.74	107.75	42.82	110.44
Belgium	91.11	60.79	101.14	64.65	88.63
Ireland	79.26	90.31	96.08	32.83	90.47
UK	101.49	88.55	93.01	66.29	95.50
Austria	89.47	99.10	82.54	69.87	63.47
France	103.86	66.06	73.49	27.12	103.73
Italy	82.18	124.75	61.32	23.60	51.42
Spain	75.44	62.90	51.21	21.73	50.91
Portugal	75.25	108.58	28.75	18.70	46.11
Greece	95.11	68.17	18.35	15.23	30.16
US	120.44	89.96	162.18	400.13	143.40
Switzerland	122.99	82.58	149.16	88.49	117.27
Australia	93.29	100.50	145.56	117.01	130.96
Norway	120.26	166.57	132.05	185.52	142.94
Canada	115.52	61.85	116.70	111.41	106.07
N. Zealand	87.28	71.34	99.76	145.07	64.88
Japan	91.65	131.43	83.88	42.91	191.50
Slovenia	68.33	29.52	88.73	21.39	87.81
Czech Rep.	66.33	33.03	34.41	22.64	47.67
Slovakia	52.11	30.57	23.02	9.89	72.79
Hungary	61.22	36.90	20.83	23.39	42.87
Lithuania	54.66	25.30	19.10	7.10	79.11
Poland	41.54	17.57	15.52	9.75	52.98
Estonia	62.50	59.74	12.17	42.66	78.69
Turkey	46.28	18.62	8.20	2.87	11.35
Romania	29.52	10.19	3.61	2.27	54.11
Bulgaria	59.95	5.27		2.99	68.15
Latvia	55.03	23.90		11.82	40.92
avge. OECD21	100.00	100.00	100.00	100.00	100.00
in levels	548.81	284.57	282.78	484.83	2563.38
avge. EU14	*96.33*	*99.70*	*86.48*	*72.10*	*85.93*
avge. candidates	54.32	26.42	25.06	14.25	57.86

Definitions:
- main *telephone lines* per 1.000 people in 1998
- number of *mobile telephones* per 1.000 people in 1998
- *personal computers* per 1000 people in 1998
- *internet hosts* per 10.000 people in 2000
- *R&D personnel* = scientists and engineers employed in R&D per million, most recent year available (ranges between 1987 and 1997).
- *Source*: World Bank, *2000/01 World Development Report.*

According to most of these indicators, mean EU performance is sligthly below the OECD average and significantly above that of the group of candidate countries (except for the TIMSS study, where the last group does slightly better than the EU14 on average). A particularly worrisome finding of IALS is that a large fraction of the population (43.54% for the OECD21 and over 60% in the candidate countries) lacks basic literacy and quantitative skills that are likely to be important both on the job and in everyday life. The corresponding figures for the (reading literacy of the) student population are only somewhat better (36.70% and 48.75% respectively) according to the PISA study. It is also interesting to note that the cross-country variation in skill levels (as measured by the standard deviation of normalized scores) is much higher at the bottom of the distribution (5th percentile score) than for mean or top performance levels, and that there is essentially no correlation between mean national performance and the range of scores in the PISA study. This suggests that the quality of the educational system is particularly important for disadvantaged individuals, and that the performance of this group can be improved without lowering average standards.

Finally, Table A4.12 collects various indicators of the penetration of ICT technologies and of R&D effort. In terms of these indicators, most candidate countries are lagging well behind the EU which is, in turn, far below US standards. Within the EU, there is a clear divide between the north and the south, with the Scandinavian countries at one end of the scale and Spain, Portugal and Greece at the other, in terms of indices of computer and internet use and R&D investment.

5. Social capital: a survey of the theoretical and empirical literature

The attention paid by economists to social capital has been rapidly increasing in the last decade. The term social capital was rendered popular by the contributions of Coleman (1988, 1990) and Putnam (1993, 1995) and by now the World Bank (2002) has an excellent internet site with an entire electronic library on the subject. Coleman starts with the consideration that social interaction brings about long lasting patterns of relations, which constitute a resource available to individual actors. Such a resource may be accumulated or depleted over time and is defined by its productive function: it allows actors to reach goals otherwise not reachable or it diminishes the cost of reaching them. Thus, it may be thought of as a peculiar form of capital, namely a 'social capital', whose specific characteristic consists in the fact that it is not incorporated in physical goods or in single human beings, as physical and human capital, but rather in social relations: it is an attribute of social structures. Examples of social capital are the level of trust and the information potential incorporated in relations, the existence of civic norms with effective sanctions, and the presence of hierarchical and horizontal relations and organizations.

A critical difference between social capital and other forms of capital, stressed by Coleman, is that it presents a key aspect of public goods: ‘As an attribute of the social structure in which a person is embedded, it is not the private property of any of the persons who benefit from it’. This poses a problem of under-investment, since ‘there will be in society an imbalance between the relative investment in organizations that produce private goods for the market and in organizations (often voluntary associations) from which the benefits are not captured – an imbalance in the sense that if the positive externalities created by such social capital could be internalized, it would come to exist in greater quantity’. Thus, private investment in social capital could fall short of the social optimum; on the other hand, if social capital is accumulated through interaction among individuals, public provision cannot be a solution either. One of the key contributions of social capital, according to Coleman, is to the accumulation of human capital: it is much easier to develop individual skills in a socially rich environment than in a socially poor one. Since human capital accumulation constitutes an engine of growth in advanced economies, social capital appears in a way as a deep root of growth processes.

Putnam (1993a) investigates the link between social capital and economic and political performance in Italy and finds that a great part of the difference in development between Southern and Northern Italian regions is ‘explained’ by the different presence of networks of horizontal organizations, which is a historical heritage and constitutes a form social capital. In particular, he shows that local governments are more efficient where civic engagement is stronger, and argues that civic engagement is strictly related to the presence of horizontal associational networks. In other works (1993b, 1995a, 1995b, 1996, 2000) he extends the analysis of social capital. In particular, applying it to the U.S., he argues that the stock of American social capital has been declining in the late Twentieth Century, mainly due to the disappearance of the ‘civic generation’, come to age between the Great Depression and World War II, and to television, that keeps individuals apart from one another. Nevertheless, his last work also discusses some signs of revival.

Already in these contributions, the authors do not always refer the term social capital to the same thing: Putnam’s definition is relatively narrow, whereas Coleman’s one is broader. The World Bank now defines social capital at the broadest level as ‘the norms and networks that enable collective action’. Different authors have proposed still different definitions, so that by now ‘social capital’ denotes more a whole strand of research than a single concept. Our first step is consequently to review the various theoretical definitions and to provide a conceptual clarification. Next, we consider the empirical problem of measuring social capital and its effects. The subsequent step is to analyze the process of social capital accumulation. Finally, we consider some policy implications, with particular attention to Europe.

a. What is social capital?

Let us start with a rather general definition of social capital – adapted from the World Bank – as the norms and social relations embedded in the social structure of a group that enable people to coordinate action to achieve desired goals. This definition deserves some comments. First of all, the group considered might consist of only one individual, at one extreme, as well as of the whole society, at the opposite extreme; correspondingly, we can define social capital at the individual as well as at the aggregate level, and we can choose between focusing on a specific group or on the society as a whole. Secondly, social capital consists of norms and social relations, which are attributes of the social structure. They can be reinforced or weakened over time, but at a given point in time they constitute a stock. Third, this stock is 'productive', in the sense that it allows group members to reach their goals. Such goals may concern standard output and income, but may also concern socially provided goods, like status and friendship. Moreover, the goals pursued by one group may be in accordance or contrary to those of other groups, so that social capital may display both positive and negative externalities (for instance, it may serve cooperative as well as rent-extracting purposes). Fourth, social capital is both accumulated and displays its effects through social interaction: it is this way that norms and relations are reinforced or weakened and it is this way that coordination among people is achieved. Such coordination may take place at two levels: either within the group members ('bonding social capital'), or with non-members ('bridging social capital'). There is an intrinsic difficulty in the aggregation of social capital, because what is productive for a group may either hurt or benefit a different group: if we collect together groups with a strong 'bonding' social capital, we do not necessarily end up with a high aggregate level of social capital; 'bridging' links play a crucial role. For this reason, it is useful to work both with an individual-level definition of social capital and with a group-level one. In the literature both are present. Let us consider them in turn.

i. Individual social capital

Glaeser, Laibson and Sacerdote (2000) propose to define individual social capital as an individual's social skills, which are partly innate ('e.g., being extroverted and charismatic'), but partly cultivated (e.g., popularity), i.e., they are the result of an investment. Social skills enable an individual to 'to reap market and non-market returns from interaction with others. As such, individual social capital might be seen as the social component of human capital'. Not all of the social skills which are beneficial to an individual are also beneficial to the aggregate outcome of social interaction: for instance, the ability to persuade others that you are trustworthy when you are not generates a negative externality (think e.g. of some sellers of encyclopedias or of used cars), whereas the ability to induce others to participate to a socially beneficial project generates a positive one. Moreover, the same social skills may be used sometimes to increase aggregate outcome, but sometimes only to increase the slice reaped by their owner, with a possible aggregate loss. This problem makes it difficult to aggregate individual social capital over a whole economy (or even over a group), since one should incorporate 'all of the cross-person externalities generated by the different types of individual social capital'. The consequence is that 'the determinants of social capital at the individual level may not always determine social capital at the society-level'. On the other side, the big advantage of this framework is that it allows studying individual decisions of investment in social capital with standard investment models, which provide predictions that can be confronted with the data. Glaeser, Laibson and Sacerdote perform such exercise and find that individuals invest in social skills in the same way as they invest in human capital.

Two remarks are in order. On one side, Glaeser, Laibson and Sacerdote's definition of social capital does not really fit the definition we have given above, since they focus on individual characteristics and not on traits of the social structure. As they recognize, what they are analyzing is the social component of human capital, which, for the sake of clarity, should perhaps be kept separated from the concept of social capital. On the other side, the amount of social skills belonging to an individual is highly correlated with the amount of his or her social connections, an aspect that is better compatible with our definition.

In this spirit, DiPasquale and Glaeser (1999) define individual social capital as an individual's connections to others and argue that it matters much for private provision of local amenities and of local public goods. They also investigate empirically whether homeownership increases investment in local amenities and social capital and find that indeed it does, especially because it reduces individual mobility. We discuss this last point in Section 4.

ii. Group social capital

At the aggregate level, definitions of social capital tend to focus either on the density of trust, which facilitates collective action and reduces free-riding, or on networks of civic engagement and of horizontal associations, following Putnam. Although these two aspects overlap to some extent, so that it is often not easy to distinguish between them, they have given rise to two strands of the literature.

Trust

Although at first sight very intuitive, the notion of trust is quite hard to define theoretically in a clear-cut way. There is a huge literature on this topic, but its subtleties are probably not so relevant for an aggregate theory of social capital, especially when it comes to the empirical side. A relevant feature of trusting behavior seems to be that it exposes an individual to the risk of being worse off, if others behave in a purely selfish way. A key effect is that trusting others may make them more trustworthy. If this happens, the advantages of cooperation may be exploited, if it does not, trusting people may be exploited by non-trustworthy ones.

Paldam and Svendsen (2000) define social capital as 'the density of trust within a group' and notice that 'the group may be extended to the whole society', consistently with the definition we gave above. They discuss the link between social capital theories dealing with goodwill (management), credibility (macroeconomic policy), cooperative solutions (game theory) and group norms (anthropology and psychology), and point out three possible, non mutually exclusive approaches to social capital: as a factor in a production function, as a factor that reduces transaction costs and as determinant of monitoring costs.

Fukuyama (1995a, 1995b) identifies social capital with trust and argues that it determines the industrial structure of an economy. Germany, Japan and the United States, for instance, are high trust societies, where trust is not restricted to the family, but rather generalized, whereas Taiwan, Hong-Kong, Italy and France are examples of low-trust societies. In the former group of countries it is easy to find giant, professionally managed corporations, because people are better able to cooperate on an enlarged scale, whereas, in the latter group, smaller, family-owned and -managed firms dominate the industrial structure. In general, Fukuyama argues that the strength of family ties may be detrimental to the emergence of large organizations, and that, where familism is not accompanied by a strong culture of work and education, it may lead to stagnation, as pointed out, e.g., by Banfield (1958). This does not automatically imply that high trust, and hence large companies, are *per se* better performing or even better for aggregate growth, since what they gain in scale may be lost in flexibility and rapidity of decision making. The economic success of Northern Italy provides a good example.

The theoretical relationship between trust and growth is investigated by Zak and Knack (2001) through a moral hazard model, in which formal and informal institutions determine the amount of monitoring that a principal needs to exercise over an agent. They argue that 'informal sanctions depend on, or are facilitated by, social ties', which can be captured by a notion of social distance, and that monitoring costs and risk aversion may make low trust societies have lower income and lower investments, and thereby lower growth. Moreover, they add that trust is lower in more heterogeneous societies because a higher social distance among actors weakens informal controls. As a consequence, in such societies growth may be lower as well: there may be a 'low trust poverty trap'.

Indeed, one can observe that their model deals more with informal sanctions than with trust: once we consider the incentives induced by such sanctions, we can avoid any reference to trust without conceptually losing anything (in Williamson's words, 'calculative trust is a contradiction in terms'). The point is that trust is the complex product of a structure of social relations, of the interactions that take place in it, and of how these shape individual identities and motivations, and finally behaviors. So let us now turn to a more structural point of view.

Social norms and networks

As we pointed out above, Putnam defines social capital in terms of networks of civic engagement and of horizontal associations. Norms and associations are a relatively stable attribute of a social structure, and can be thought of as a stock. They arise through social interaction and they shape the way individuals interact with one another, so that social interaction (a flow) is both a source of social capital and the means through which it displays its productive services. If a norm of cooperation or of participation is effective, those behaviors that are in accordance with it will also appear quite stable. This has generated some confusion in the theoretical definition of social capital, since the term is sometimes referred to the stock of social norms and networks and sometimes to the specific form of interaction that arises out of it. This has led some author, for instance Bowles e Gintis (2000), to abandon the term social capital in favor of something they perceive as more precise. In particular, Bowles and Gintis prefer to speak of community governance, arguing that it is often the case in the literature that the term social capital is referred to what groups *do* rather than to what they *own*, and such aspect is better captured by the notion of community governance – as opposed to the governance mechanisms of the state and of the market – than by the notion of social capital. Notice, however, that considering just the community of direct and frequent interactions, expressed by Bowles and Gintis' idea of community governance, is restrictive, since it may overlook the strength of weak ties, stressed e.g. by Granovetter (1973) and by Narayan (1999), and the relevance of generalized trust, as we have discussed above.

Aware of such conceptual problems, Fukuyama (1999) proposes to change his previous definition of social capital in terms of trust into the following one: 'social capital is an instantiated informal norm that promotes cooperation between two or more individuals'. He argues that 'by this definition, trust, networks, civil society, and the like which have been associated with social capital are all epiphenomenal, arising as a result of social capital but not constituting social capital itself'. One crucial aspect of such definition is the extent of validity of the norms considered (also referred to by Fukuyama as the 'radius of trust, that is, the circle of people among whom cooperative norms are operative'). This leads to a more precise specification of the group (or institution) to which one refers the term social capital.

A second crucial aspect is that cooperation within a certain group may have positive as well as negative external effects on other groups. For instance, the degree of participation to associational activities does not necessarily increase aggregate (society-level) social capital, as hypothesized by Putnam: Olson (1982) emphasizes that the purpose of some groups is to exert a distributive pressure, i.e. to seek rents, and that active participation to such groups indeed increases the level of distributive struggle in a society and decreases social capital.
Both these aspects – the extension of the group and the kind of external effects – are captured by Collier (1998). He starts with a definition of social capital in terms of those externality-generating social interactions which are either themselves durable or whose effects are durable, and he carefully distinguishes among the various institutional levels at which social capital may be present: the family, the firm, the government and the civil society.

Social capital at the firm level is the easiest one to study. As already noticed by Coleman, the internal organization of a firm is intentionally designed to make profits, so that this is one of the few cases in which social capital is the product of a specific investment and not just the by-product of other activities. Such aspects are widely studied in management and business disciplines, although without any reference to the notion of social capital. A proof of their relevance is the amount of money that firms spend not only to design internal structures, but also to train managers and workers to work in groups: management consultants and labor psychologists are often very well paid to provide such training, evidently because it pays off. Inter-firm linkages, typical, for instance, of industrial districts, constitute a second form of firm-level social capital. Signorini (2000) presents a very detailed analysis of the Italian case, which helps to understand how the success of many small Italian firms relies upon external economies that compensate the scale disadvantage.

Coming to the family, we have noticed above that Fukuyama and Banfield, among others, emphasize the possible contrast between strong family ties and more aggregate levels of social capital. Family is indeed the primary source of narrow trust, i.e. trust in peer or primary groups, but whether or not trust generalizes and extends beyond kinship relations depends to a high degree both upon the kind of interaction that takes place in the intermediate structures of the civil society, and upon the well functioning of the government, which can provide, for instance, a reliable judicial system.

As we have seen, Putnam emphasizes the first aspect, i.e. participation in associational networks at the level of the civil society. However, whether trust remains confined within certain groups or generalizes beyond their scope depends to a high degree on whether groups form along social cleavages or across them: one needs to look at the specific kind of social participation and not just at the density of associations, although the latter one may be sometimes the best empirical proxy available.

As far as the link between social capital and the well functioning of government is concerned, Narayan (1999) points out it is not univocally of substitution or of complementarity, since when either of them is poor, the other one may work as a substitute, but if both of them are rich, they indeed work as complements (he also provides a detailed discussion of the empirical evidence available). Exactly the fact that formal institution (market and state) are not working properly may increase reliance on primary groups: what Rose (1998) claims happened in Russia after the collapse of Soviet Union, but he also points out that such reliance on primary groups had been previously fostered by the extreme centralization and had emerged as a way of defending oneself from the invasion of the state. Another interesting example of how government, family and civil society interplay to shape trust, norms and connections (social capital) at the level of some groups, but with troublesome extensions to the whole society, is Gambetta's (1993) analysis of the Sicilian Mafia.

The problem is that social capital tends to exert positive aggregate effects when trust, norms and networks that foster cooperation extend beyond primary, ethnic, linguistic or even income groups and form 'bridges' among different groups. This last point is made with particular strength by Narayan (1999), who observes that the same links that keep together the members of a group may also exclude the non-members, and who displays an analytical framework to study 'bonding' and 'bridging' (i.e. intra-group and inter-group) social capital at the level of the civil society, together with its connections to the functioning of the state.

b. Empirical evidence on social capital and aggregate performance

There is by now a wide empirical literature on the effects of social capital on aggregate performance. The World Bank considers a list of eleven broad topics to which social capital is relevant. Here we analyze only some of them: in particular, we consider empirical evidence on the effects of social capital on growth, trade and migration, finance, government performance, education, crime and violence.

i. Social capital and growth

Knack and Keefer (1997) examine various possible empirical proxies for social capital, corresponding to the different aspects emphasized by the theoretical literature, and assess their impact on growth. They discuss three main relationships: between trust and civic norms and economic growth; between associational activity and growth; and between trust and civic norms and their determinants, including associational activity and formal institutions. On the latter aspect we shall come back in the next section. Let us consider here the first two ones.

Knack and Keefer consider data from the World Value Survey for 29 market economies between 1981 and 1991. As a proxy for trust (TRUST) they take for each nation the percentage of respondents that most people can be trusted (after deleting the 'don't know' answers) to the following question: 'Generally speaking, would you say that most people can be trusted, or that you can't be too careful in dealing with people?'. This measure of trust exhibits a high cross-country variance and high serial autocorrelation within each country. To capture the strength of norms of civic cooperation, they construct a variable (CIVIC) on the base of the answers to various questions about how individuals evaluate some anti-civic behaviors. These two variables are highly positively correlated and both of them are designed to capture generalized trust and cooperative attitudes, rather than social capital at the level of a specific group. The first main finding of Knack and Keefer is that 'trust and civic cooperation are associated with stronger economic performance'. In particular, they find that one standard deviation change in TRUST is associated with a change in growth of more than half of a standard deviation. This result seems to be quite robust. The second question they address concerns the effects of associational activities, about which, as noticed above, Olson and Putnam have contrasting hypotheses. As a proxy for the density of horizontal networks in a society (GROUPS), they consider the average number of groups cited per respondent when faced with the question of whether they belong to any of a list of groups of ten kinds. The second main result is that 'associational activity is not correlated with economic performance – contrary to Putnam's (1993) findings across Italian regions'. They also split the data to identify the possibly contrasting effects of 'Putnamesque' and 'Olsonian' groups, i.e., of groups that 'involve interactions that can build trust and cooperative habits' and of groups with redistributive goals, respectively. The results are contrary to what the theory predicts, but, by admission of the authors, they should be regarded as only preliminary. Their relevance, rather than substantial, is methodological.

Zak and Knack (2001) perform a similar analysis, using the same variable for trust, but with more data. In particular, while Knack and Keefer's investigation concerns 29 OECD countries, Zak and Knack add to the sample 12 additional countries. The effect of the larger sample is basically that it reinforces the statistical impact of trust on investments and growth. Moreover, they investigate the impact of formal institutions and social homogeneity, finding that they 'increase growth in part by building trust'.

A related empirical contribution is due to Temple and Johnson (1998), who show that indexes of 'social capability' constructed in the early Sixties, adapted from the work of Adelman and Morris (1967), are good predictors of long run growth for a wide set of developing countries. In particular, they find that a mass communication index is robustly correlated with growth and they argue that this may be due to the fact that 'it captures the social capital of developing countries'. Although these results are striking, it is a bit hard to understand exactly how one should evaluate them, because the social capability index used is quite composite and not so straightforward to interpret, and because it is not very clear how the index of mass communication is related to social capital.

Taken together, this evidence consistently shows that social capital, especially in the form captured by the variable TRUST, has a relevant impact on growth. Glaeser, Laibson, Sheinkman and Soutter (2000) address the question of what exactly TRUST measures. To this purpose, they use two experiments and a survey, and assess that standard questions about trust, such as the one reflected in TRUST, provide a better measure of the level of trustworthiness in a society rather than of trusting behavior. Nevertheless, they also assess the possibility to gain robust measures of social capital (trust) as an individual-level variable. In particular, measures of past trusting behavior predict an individual's trust better than abstract questions.

ii. Social capital and government performance

Hall and Jones (1999) explain a relevant part of country productivity as due to institutions and government policies (what they call social infrastructures). Since these characteristics are endogenous, they propose a set of instruments. A growing amount of evidence is now showing that the quality of government is positively influenced by social capital. An in-depth investigation of the determinants of government quality is due to La Porta et al (1999). They evaluate empirically the ability of economic, political and cultural theories to explain the observed quality of governments, according to different measures. Broadly speaking, they find that economic theories focusing on efficiency are rejected by the data; political theories focusing on redistribution are highly and robustly supported by the evidence (as instrument for redistributive tendencies they use ethnolinguistic heterogeneity and legal system); finally, cultural theories focusing on trust, social norms of tolerance and work ethic cannot be rejected. In particular, as an instrument for such cultural characteristics they use religion, in the spirit of Weber (1958), and find essentially that 'predominantly Protestant countries have better government than either predominantly Catholic or predominantly Muslim countries'. Such results prove to be robust to many alternative specifications and confirm earlier findings of the same authors.

One of the main functions of governments is to provide public goods. Alesina, Baqir and Easterly (1999) relate spending in public goods to ethnic division at the level of cities, metropolitan areas and urban counties in the United States. Their finding that 'more ethnically diverse jurisdictions in the United States have higher spending and higher deficits/debt per capita, and yet devote lower shares of spending to core public goods like education and roads' is consistent with the idea that 'heterogeneous and polarized societies will value public goods less'.

The relationship between the variables considered in these studies, like ethnolinguistic heterogeneity and religion, and social capital will be considered in the next section. Here, in turn, we pass to the analysis of the impact of social capital on education.

iii. Social capital and education

One of the possibly most relevant contributions of social capital is to the formation of human capital. This was very early recognized by Coleman (1988), who argued that the same basic individual skills have much better chances of being well cultivated and developed in a socially rich environment than in a socially poor one. Goldin and Katz (1999), in a study on the development of secondary education in the United States and in particular in Iowa, acknowledge that, 'because educational decisions are made primarily at a local level in the United States, the production of human capital depends largely on social capital lodged in small communities'. As a measure of community-level social capital they use 'the amount of public resources committed to education as a fraction of the total resources of the community, given by income'. It is interesting to see that this 'indicator of educational commitment rises steeply during the 1910S and for most of the 1920S' and then rises again in the 150S, but it is harder to take it as a direct measure of community-level social capital. However, one further empirical observation supports this interpretation: 'one good reason for building schools in rural America was to stop the drift of the population to the cities', i.e., to save and promote community cohesion. The almost ubiquitous public provision of schooling is consistent with the view 'public funding was part of an intergenerational loan. According to this view, homogeneous communities, in which people tend to remain and take an active interest in each other, would be more likely to provide intergenerational loans'. Indeed, such communities were present in Iowa, one of the leading states in the development of schooling. In particular, 'smaller towns of Iowa had the highest rates of secondary school attendance', even though a more precise assessment of why this was the case turns out to be difficult.

A relevant problem in empirical analyses of the link between social capital and education is that there is an issue of reverse causation. Goldin and Katz find a strong correlation 'between an index of social capital today (combining measures of associational activities, social trust, and political/civic participation)' and 'the high school graduation rate in 1928'. They conclude that social capital has a double role of condition for accumulation of human capital and of handmaiden of human capital. The issue of how education determines social capital is also tackled by Helliwell and Putnam (1999), to whom we will turn in the next section.

iv. Social capital and crime

It is intuitive that social capital, determining the degree of social cohesion, may have a relevant influence on the rates of crime and violence. Coleman (1990) already stresses this point. Glaeser, Sacerdote and Scheinkman (1996) explore this issue. In face of several possible empirical explanations of the high variance of crime across time and space, they take a sharp interactionist view, assessing that 'positive covariance across agents' decisions about crime is the only explanation for variance in crime rates higher than the variance predicted by differences in local conditions'. Patterns of local interaction thus seem to drive crime to a relevant extent, the more so as far as young people and petty crimes are concerned.

v. Social capital and financial development

Guiso, Sapienza and Zingales (2000) investigate the impact of trust on financial development. They argue that 'financing is nothing but an exchange of a sum of money today for a promise to return more money in the future. Whether such an exchange will take place depends upon not only the enforceability of contracts, but also the extent the financier trusts the financee. In fact, financial contracts are trust intensive *par excellence*. Thus, if trust matters, it should matter most for the development of financial markets'. Their proxy for trust is different from standard survey measures, since they consider participation in elections and blood donation. They use data on Italian regions, which present the advantage of having the same 'legal, administrative, judiciary, regulatory and tax system', but at the same time very different levels of social capital, and assess that higher trust increases investment in stocks, access to credit and use of checks, whereas it reduces investment in cash and resorting to informal credit channels. Moreover, such effects appear to be more relevant where legal enforcement is weaker and among less-educated people.

c. Social capital accumulation

The theoretical and empirical literature considered so far shows that social capital, defined and measured in several ways, matters for a great variety of economic outcomes. This finding raises the questions of how social capital is accumulated and of whether its accumulation may be enhanced by policy intervention. We address the first question in this section and the second in the next one.

i. Theory

There is not much theoretical work discussing the determinants of social capital. According to Glaeser, Laibson and Sacerdote (2000), as we noticed above, this lack is due to the fact that most definitions and measures of social capital are aggregate ones, whereas economists are used to think of capital accumulation as a result of individual investments. They therefore define social capital in terms of individual social skills, i.e. as the social component of human capital, and apply a standard model of individual investment. Such model implies that investment in social capital should increase with patience and with the relevance of positive externalities in the return to social capital investment (e.g., individuals invest more in social skills in those occupations where returns to social skills are higher), whereas it decreases when are higher expected mobility (e.g., homeowners should invest more in social skills), the opportunity cost of time, the rate of depreciation and the degree of community-specificity of social capital. Moreover, investment should decrease with age, but, assuming that individual endowment at birth is sufficiently low, the stock of individual social capital should first increase and then decrease with age.

This model is theoretically very clear, but it does not solve the problem of aggregation, so that aggregate determinants of social capital might be quite different from the determinants of investment in individual social skills. In the authors' words, 'understanding the link between individual and aggregate social capital is important, difficult, and best left to future research'.

If we consider group-level definitions of social capital, both in terms of trust and of social networks, the theory of social capital accumulation focuses on the individual problems of whether to trust or not and of whether to join a group or not.

As far as trust is concerned, Alesina and La Ferrara (2000b) admit that 'the theory of what determines trust is sketchy at best'. They consider 'five broad factors influencing how much people trust others: 1) individual culture, traditions and religion; 2) how long an individual has lived in a community with a stable composition; 3) recent personal history of misfortune; 4) the perception of being part of a discriminated group; 5) several characteristics of the composition of one's community, including its racial and income heterogeneity'. However, they do not display any formal model.

As far as participation in groups and associational activities is concerned, Alesina and La Ferrara (2000a) focus on population heterogeneity and argue that its link with social participation is theoretically ambiguous. On one hand, heterogeneity could increase the number of associations, since each group would like to have its own ones. On the other hand, heterogeneity may also increase the likelihood of mixed groups being formed. This, in turn, may reduce participation if individuals prefer to interact with others similar to them (e.g. in terms of income, 'race' or ethnicity).

ii. Evidence

Let us now consider the empirical evidence on the accumulation of the different kinds of social capital (individual social skills, trust and social participation) and then on the extent of the decline of social capital assessed by Putnam .

Individual social capital

Using data from the General Social Survey in the U.S. from 1972 to 1998, Glaeser, Laibson and Sacerdote's (2000) find that their theoretical model (discussed above) fits well the data. In particular, organization membership has an inverted u-shape over the life-cycle; the prediction that expected mobility reduces individual social capital seems to be consistent with the data, although they do not find a good instrument for expected mobility; more social occupations induce higher investment in social skills; the evidence on the impact of homeownership on group membership varies according to the kind of group (for instance, it is low for political groups and high for school service): in general, it seems that homeownership affects social capital more through its effect of reduced mobility than through patrimonial effects, i.e. through incentives due to expected changes in property value; investment in individual social skills might be indeed partly due to the opportunity cost of time, but it is very difficult to find a satisfying empirical assessment of this relationship; physical distance, unsurprisingly, affects negatively social connections; education and membership in organizations are positively correlated, as predicted by the theory, since patience increases both investment in human capital and in social capital; finally, the empirical evidence they find leaves the authors agnostic as to the relevance of interpersonal complementarities.

As a general point, one might notice that most of the empirical proxies used by Glaeser, Laibson and Sacerdote are more related with the rest of the literature on social capital than with their own definition as the social component of human capital. Indeed, they acknowledge that standard measures of individual trust and of organization membership do not capture in an obvious way what they define as 'social capital'.

As discussed above, DiPasquale and Glaeser (1999) also start with an individual-level definition of social capital, although they stress more an individual's social connections with others. Empirically, they study how homeownership may create incentives to social capital accumulation and to provision of local amenities. This might work either through the fact that such investments increase the value of property, or because owing a home reduces mobility and thus increases the time one expects to enjoy the fruits of such investments. They use data from the U.S. General Social Survey and from the German Socio-Economic Panel. Both in the U.S. and in Germany they find a strong correlation between homeownership and measures of civic engagement in one's community (e.g. membership in nonprofessional organizations, knowing the names of local political representatives, voting in local elections, gardening and church attendance). Such effects are weaker in Germany than in the U.S.; moreover, in the U.S. a larger fraction of the effect seems to be attributable to increased community tenure. The authors are very careful about policy conclusions, since unobserved omitted variables might play a relevant role (homeowners may be different from renters), and since they do not measure either the positive or the negative externalities linked to homeownership and decreased mobility.

A general conclusion is that individual incentives matter for social capital accumulation, but not in a naïve way. Social rewards may provide more effective incentives to social capital accumulation than material ones, a point that hints at the relevance of social capital for 'relational production' besides material production and that should be kept in mind when thinking of policy intervention.

Trust

Alesina and La Ferrara (2000b) consider both individual experiences and community characteristics as possible determinants of individual trust. Using data from the General Social Survey for the United States from 1974 to 1994, they find that the major causes of low trust are recent traumatic experiences, belonging to a discriminated group, low income, low education, living in a society with strong 'racial' cleavages or in one with high income inequality. Religious beliefs and ethnic origins, in contrast, are found not to affect trust significantly.

Glaeser et al (2000) combine survey and experimental data to separately identify the determinants of trust and of trustworthiness. Two of their findings are that a smaller social distance among individuals, for instance due to joint group membership or the same 'race' or nationality, increases both trust and trustworthiness; moreover, an individual's higher status induces others to behave in a more trustworthy manner toward him or her.

Finally, Knack and Keefer (1997) find that 'trust and norms of civic cooperation are stronger in countries with formal institutions that effectively protect property and contract rights, and in countries that are less polarized along lines of class or ethnicity'.

Social participation

Alesina and La Ferrara (2000a) study participation in associational activities like religious groups, sport groups, hobby clubs, unions, and so on (they consider participation in a list of 16 different kinds of groups). They analyze data for metropolitan areas in the U.S. from 1974 to 1994, mainly from the General Social Survey. They run a probit regression to explain the probability of social participation, controlling for individual and community characteristics. The key results are striking: social participation is higher where income inequality, 'racial' segmentation and ethnic segmentation are lower. This happens in the North/Northwest of the U.S., the opposite features appearing in the South/Southeast. Moreover, looking at participation in different kinds of groups, the authors find that heterogeneity matters less for participation in groups with a relatively *high degree of excludability* or a *low degree of close interaction* among members. Finally, they find that 'racial' segmentation matters more for individuals more averse to 'racial' mixing.

More in detail, they find that younger cohorts participate less than elder ones, providing some support to Putnam's idea of a decline in participation due to the aging of the 'older civic generation'. Years of schooling have a positive impact on participation. Women participate less than men. Black people participate more. Young children reduce parents' participation. Family income has a positive effect, 'suggesting that participation is a normal good'. Coming to community characteristics, the measures of income inequality and of racial and ethnic segmentation always have a negative impact on participation, controlling for individual variables and for year and state dummies. The authors also perform some sensitivity analysis, which confirms and even strengthen the results: they assess that an increase by one standard deviation in racial segmentation, income inequality and ethnic segmentation reduces the probability of participation by respectively eight, six and six percentage points; the impact of passing from high school dropout to high school graduate or higher is a positive increase of thirteen percentage points; moving form a full-time to a part-time job increases the propensity to participate by four percentage points; finally having a child below the age of five reduces it by 3.5 percentage points. Interestingly, the relation between participation and income seems to be increasing but not linear: convex for low levels of income and concave for high levels. Instrumenting for income inequality leaves its effect on participation highly negative and significant.

Helliwell and Putnam (1999) consider both trust and social participation at the same time. They investigate whether and how education determines social capital. They start with the observation that, although average educational levels have risen sharply in the United States in the last half century, the same did not happen to political and social participation. This is somehow puzzling, because individual education is widely acknowledged to be the best predictor of many forms of political and social engagement. Helliwell and Putnam discuss the theory trying to solve this puzzle and argue that it does not allow to reach a clear conclusion. Using data from the US General Social Survey from 1972 to 1996 and from the DDB-Needham Life Style surveys from 1975 to 1997, they assess that higher average education increases trust and does not reduce participation.

Is there a decline in social capital?

One of the main issues in the theory of social capital is the problem of a possible under-investment. Coleman (1990) raises this issue and Putnam (1995, 2000) documents empirically a decline in American social capital, identifying the main culprits in television and aging of the 'civic generation' of Americans born between 1910 and 1940. Putnam finds that television is responsible for up to a quarter of the decline in social capital and the aging of the 'civic generation' up to half of it. However, there is no widespread agreement either on the empirical relevance of such decline or on its causes.

Costa and Kahn (2001) argue that it has been overestimated by Putnam, although some forms of social capital indeed declined in the U.S. from 1952 to 1998: whereas group membership indeed diminished, the probability of volunteering did not; the largest declines are found in the time devoted to entertainment and visits with friends, relatives and neighbors. Such results are found using probit regressions with a great variety of data sources. Costa and Kahn also show that the decline in the social capital produced outside the home is mainly due to rising community heterogeneity (especially income inequality), whereas the decline in the social capital produced within the home is mainly explained by women's increased labor force participation rate (always controlling for education).

d. Policy

Policy implications are drawn in a sparse and usually very cautious way in the literature on social capital. The World Bank considers the following list of political issues, strictly connected with social capital: crime and violence, trade, education, environment, finance, health, nutrition and population, information technology, poverty and economic development, rural development, urban development and water supply and sanitation. Many of them are more relevant for developing countries than for Europe, but some of them represent hot issues in the current European political debate. Let us briefly examine some of the indications arising from the literature.

i. Individual social capital

Those contributions that emphasize individual aspects of social capital make the general point that its accumulation responds to individual incentives, but not in a naïve way. One of the difficulties here comes from the fact that intrinsic motivations may be either reinforced or crowded out by an exogenous introduction of incentive schemes. This is especially the case if incentives change the way individuals interpret and frame a situation. For instance, suppose that in a certain situation cooperation is perceived as the appropriate behavior, in accordance to a social norm, and that we now introduce a fine to sanction defective behavior; then individuals might abandon the social norm interpretation and embrace a market based one, according to which defection amounts to purchasing a good (the individual advantage arising from it) at a given price (the fine), without any remorse for a bad behavior: if the monetary cost of the fine is lower than the psychological one perceived by breaking a norm, the incentive will be counterproductive. Gneezy and Rustichini (2000) provide convincing empirical evidence of this mechanism. A second problem is that, even if incentives to individual investment in social capital were to work well, it is difficult to evaluate the aggregate impact, because one should find a way to measure interpersonal externalities.

ii. Trust

Policy indications are somewhat easier to draw if one looks at the correlates of generalized trust. In particular, policies that increase the well functioning of the state, the effective protection of property rights, a low degree of inequality in the distribution of income and a low degree of 'racial' heterogeneity create a favorable environment for the development of trust. Whether or not such policies are desirable (in particular the latter two ones) involves political issues that we do not tackle here.

The positive correlation found by Helliwell and Putnam (1999) between average education and social capital provides an additional rationale, besides the traditional ones, to invest in education even more than we are currently doing. This is especially advisable since, on one hand, there is a virtuous dynamics between human capital and social capital accumulation, and, on the other hand, trust-enhancing policies may start a multiplier mechanism. Indeed, both the theory and the experimental evidence tell us that a key effect of trust is to induce a higher trustworthiness, which in turn allows people to trust without being exploited. The role of policy may then be that of activating such mechanism, especially in low-trust environments, such as some European regions, which otherwise may remain stuck in a low-trust poverty trap, where low trust and low trustworthiness justify one another.

As we discussed above, trust-enhancing policies have a special relevance, among other things, for the purposes of long-run growth and of financial development. What may be added here is that they can play a special role in the context of the 'new economy', in which we are more and more transacting ideas (e.g. inventions, images, and so on). Unlike physical goods, whose characteristics are observable before the transaction, ideas cannot be revealed *ex ante* (once they are communicated, there is no need to purchase them any more), so that trust comes to play a prominent role. In a well operating market, reputation mechanisms may probably substitute for trust to a high degree, but in new, emerging markets such element of stability is absent, so that the level of trust and trustworthiness may determine whether some innovative, idea-intensive activities take off at all – and may in any case substantially reduce their transaction and monitoring costs.

iii. Social participation and networks

Social participation seems to be less an issue for Europe than it is for the United States. The general problem in designing participation-enhancing policies is that one cannot, by definition, force voluntary participation. With this *caveat* in mind, one can think of effective incentive schemes, which are, however, hard to formulate in general terms. Notice that the construction of networks of participation may be crucial at least at three levels. First, family- and community-level participation facilitates human capital accumulation and private provision of local amenities and of local public goods. Second, social participation at the level of the civil society generates positive externalities, at least if one focuses attention on 'Putnamesque' groups and on 'bridging' links. In affluent societies, where material needs have reached a high degree of satisfaction and relational needs assume a prominent role, these kinds of participation dynamics may be crucial for individual and social well being. Finally, cooperation networks among firms may provide at the same time those efficiency and flexibility characteristics that allow a successful adaptation in rapidly changing economies, but this is an area in which direct intervention may have positive as well as distortionary effects, so that it is hard to identify policies recommendable in general.

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