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The contribution of schooling in development accounting: Results from a nonparametric upper bound☆

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ABSTRACT

How much would output increase if underdeveloped economies were to increase their levels of schooling? We contribute to the development accounting literature by describing a nonparametric upper bound on the increase in output that can be generated by more schooling. The advantage of our approach is that the upper bound is valid for any number of schooling levels with arbitrary patterns of substitution/complementarity. Another advantage is that the upper bound is robust to certain forms of endogenous technology response to changes in schooling. We also quantify the upper bound for all economies with the necessary data, compare our results with the standard development accounting approach, and provide an update on the results using the standard approach for a large sample of countries.

1. Introduction

Low GDP per worker goes together with low schooling. For example, in the country with the lowest output per worker in 2005, half of the adult population has no schooling at all and only 5% has a college degree (Barro and Lee, 2010). In the country with output per worker at the 10th percentile, 32% of the population has no schooling and less than 1% a college degree. In the country at the 25th percentile, the population shares without schooling and with a college degree are 22% and 1% respectively. On the other hand, in the US, the share of the population without schooling is less than 0.5% and 16% have a college degree.

To some extent, such differences in attainment could reflect efficient schooling decisions in response to international differences in technology or institutional quality (e.g. Foster and Rosenzweig, 1995; Jensen, 2010; Munshi and Rosenzweig, 2006). On the other hand, it seems highly plausible that schooling attainment in poor countries is also limited by lack of access to schools (particularly in rural areas), and credit constraints that force parents to send children to work in order to provide for current consumption. Credit constraints also limit poor parents' capacity to cover tuition, uniforms, and meals. Consistent with the view that there are barriers to investment in schooling, Duflo (2001) finds large enrollment effects from an expansion in public school provision, and Schultz (2004) from the introduction of a conditional cash transfer program. The crucial importance of public funding (and other government policies) to enable mass schooling is discussed at length in Goldin and Katz (2008). It is also consistent with the existence of barriers to attainment that the returns to schooling are higher in poor countries than in rich ones (e.g. Bils and Klenow, 2000). The view that schooling attainment is in part limited by lack of access and credit constraints has led national governments, bilateral and multilateral donors, and civil-society NGOs to prioritize schooling attainment among their development goals for several decades. For example, one of the millennium development goals is universal education.

But how much of the output gap between developing and rich countries can be accounted for by differences in the quantity of schooling? Early empirical attempts to answer this question using cross-country data focused on regressions of growth (or GDP levels) on measures of educational enrollment or attainment (e.g. Barro, 1991; Benhabib and Spiegel, 1994; Caselli et al., 1996; Mankiw et al., 1992; see Krueger and Lindhal, 2001 for a survey and evaluation of this literature). One difficulty with this literature is that results on the impact of schooling did not prove robust to alternative measures of the education variable, the sample, or the estimation method. Also, it proved difficult to tackle the problem of endogeneity of schooling.

In part in response to these difficulties with the regression approach, a second wave of studies focused on calibration rather than estimation (e.g. Hall and Jones, 1999; Hendricks, 2002; Klenow

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and Rodriguez-Claire, 1997), giving rise to a thriving new literature known as development accounting. A robust result in the development accounting literature is that only a relatively small fraction of the output gap between developing and rich countries can be attributed to differences in the quantity of schooling.[1] This result appears to dampen expectations that current efforts at boosting schooling in poor countries, even if successful, will do much to close the gaps in living standards.[2]

The somewhat negative result from development accounting is obtained using a parametric approach. Technology differences across countries are assumed to be skill neutral, and workers with different attainment are perfect substitutes. Relative wages are then used to gauge the relative efficiency in production of workers with different attainment. A potential concern is that there is by now a consensus that differences in technology across countries or over time are generally not Hicks-neutral, and that perfect substitutability among different schooling levels is rejected by the empirical evidence (e.g. Angrist, 1995; Autor and Katz, 1999; Caselli and Coleman, 2006; Ciccone and Peri, 2005; Goldin and Katz, 1998; Katz and Murphy, 1992; and Krusell et al., 2000). Once the assumptions of perfect substitutability among schooling levels and Hicks-neutral technology differences are discarded, can we still say something about the output gap between developing and rich countries attributable to schooling?

Answering this question while sticking to a parametric approach requires assuming that there are only two imperfectly substitutable skill types, that the elasticity of substitution between these skill types is the same in all countries, and that this elasticity of substitution is equal to the elasticity of substitution in countries where instrumental-variable estimates are available (e.g. Angrist, 1995; Ciccone and Peri, 2005). These assumptions are quite strong. For example, the evidence indicates that dividing the labor force in just two skill groups misses out on important margins of substitution (Autor et al., 2006; Goos and Manning, 2007). Once there are more than 3 skill types, estimation of elasticities of substitution becomes notoriously difficult for two main reasons. First, there are multiple, non-nested ways of capturing patterns of substitutability/complementarity and this make it difficult to avoid misspecification (e.g. Duffy et al., 2004). Second, relative skill supplies and relative wages are jointly determined in equilibrium and estimation therefore requires instruments for relative supplies. It is already challenging to find convincing instruments for two skill types and we are not aware of instrumental-variable estimates when there are 3 or more imperfectly substitutable skills groups.

We explore an alternative to the parametric production function approach. In particular, we make the observation that when aggregate production functions are weakly concave in inputs, assuming perfect substitutability among different schooling levels yields an upper bound on the increase in output that can be generated by more schooling. This is true irrespective of the pattern of substitutability/complementarity among schooling levels, as well as the pattern of cross-country non-neutrality in technology. This basic observation does not appear to have been made in the development accounting literature. It is worthwhile noting that the production functions used in the development accounting literature satisfy the assumption of weak concavity in inputs. Hence, our approach yields an upper bound on the increase one would obtain using the production functions in the literature. Moreover, the assumption of weakly concave aggregate production functions is fundamental for the development accounting approach as it is clear that without it, inferring marginal productivities from market prices cannot yield interesting insights into the factors accounting for differences in economic development.

The intuition for why the assumption of perfect substitutability yields an upper bound on the increase in output generated by more schooling is easiest to explain in a model with two schooling levels, schooled and unschooled. In this case, an increase in the share of schooled workers has, in general, two types of effects on output. The first effect is that more schooling increases the share of more productive workers, which increases output. The second effect is that more schooling raises the marginal productivity of unschooled workers and lowers the marginal productivity of schooled workers. When assuming perfect substitutability between schooling levels, one rules out the second effect. This implies an overstatement of the output increase when the production function is weakly concave, because the increase in the marginal productivity of unschooled workers is more than offset by the decrease in the marginal productivity of schooled workers. The result that increases in marginal productivities produced by more schooling are more than offset by decreases in marginal productivities continues to hold for an arbitrary number of schooling types with any pattern of substitutability/complementarity as long as the production function is weakly concave. Hence, assuming perfect substitutability among different schooling levels yields an upper bound on the increase in output generated by more schooling.

From the basic observation that assuming perfect substitutability among schooling levels yields an upper bound on output increases, and with a few ancillary assumptions – mainly that physical capital adjusts to the change in schooling so as to keep the marginal product of physical capital unchanged – we derive a formula that computes the upper bound using exclusively data on the structure of relative wages of workers with different schooling levels. We apply our upper-bound calculations to two data sets. In one data set of 9 countries we have detailed wage data for up to 10 schooling-attainment groups for various years between 1960 and 2005. In another data set of about 90 countries we use evidence on Mincerian returns to proxy for the structure of relative wages among 7 attainment groups. Our calculations yield output gains from reaching a distribution of schooling attainment similar to the US that are sizeable as a proportion of initial output. However, these gains are much smaller when measured as a proportion of the existing output gap with the US. These results are in line with the conclusions from development accounting (e.g. Caselli, 2005; Hall and Jones, 1999; Klenow and Rodriguez-Claire, 1997). This is not surprising as these studies assume that workers with different schooling attainment are perfect substitutes and therefore end up working with a formula that is very similar to our upper bound.

A potential limitation of the parametric approach to development accounting is that it typically assumes that changes in schooling attainment leave technology unchanged.[3] This assumption would be wrong if there were important schooling externalities or significant appropriate-technology effects. We discuss the extent to which our nonparametric upper bound is robust to endogenous technology

[1] Recently this result has been challenged by Gennaioli et al. (forthcoming), who argue that much of top managers' and entrepreneurs' returns to schooling are formally earned as profits, and therefore unaccounted for by standard microeconomic estimates of the returns to schooling – a key ingredient in most development-accounting calculations. After accounting for managers' returns to schooling, they argue that the average Mincerian return to schooling is around 20%, about double what is usually found in the literature. Using this higher return leads to a large increase in the explanatory power of human capital for income differences. Gennaioli et al.'s estimate of managers' returns to schooling is based on firm-level valued-added regressions that do not control for manager characteristics other than schooling. As such characteristics may be correlated with managers' schooling, it is difficult to know what part of the return can be attributed to schooling only.

[2] Partially in response to these findings, some authors have advocated a shift to cross-country differences in the quality of schooling (e.g. Erosa et al., 2010; Hanushek and Woessmann, 2008, 2011; Manuelli and Sheshadri, 2010). Other authors have emphasized aspects of human capital such as health (Weil, 2007) and experience (Lagakos et al., 2012).

[3] This is not always the case however. For example, a recent paper by Jones (2011) computes rich–poor human capital ratios using relative wages in poor as well as rich countries. His approach implies that computed human capital ratios will also reflect differences in human capital quality and – to the extent they affect relative wages – differences in technology. In Jones' framework, the perfect substitution case yields a lower bound on the income increase that can be achieved by raising human capital in poor countries.

responses. Perhaps surprisingly, we find that our approach also works in the appropriate-technology framework developed and estimated by Caselli and Coleman (2006). On the other hand, and less surprisingly, our approach does not yield an upper bound in the presence of aggregate schooling externalities. However, the empirical evidence suggests that such externalities are not large enough for our upper bound to be far off. We therefore conclude that our upper-bound calculations could well continue to be useful even in a world where technology responds endogenously to relative skill supplies.

The rest of the paper is organized as follows. Section 2 derives the upper bound. Section 3 shows the results from our calculations. Section 4 discusses the robustness of our upper-bound calculation to making technology endogenous to schooling. Section 5 concludes.

2. Derivation of the upper bound

Suppose that output Y is produced with physical capital K and workers with different levels of schooling attainment,

$$Y = F(K, L_0, L_1, \ldots L_m) \tag{1}$$

where L_i denotes workers with schooling attainment $i = 0,..,m$. The (country-specific) production function F is assumed to be increasing in all arguments, subject to constant returns to scale, and weakly concave in inputs. Moreover, F is taken to be twice continuously differentiable.

The question we want to answer is: how much would output per worker in a country increase if workers were to have more schooling. Specifically, define s_i as the share of the labor force with schooling attainment i, and $\mathbf{s} = [s_0, s_1.., s_{i,...}s_m]$ as the vector collecting all the shares. We want to know the increase in output per worker if schooling were to change from the current schooling distribution $\mathbf{s}^1$ to a schooling distribution $\mathbf{s}^2$ with more weight on higher schooling attainment. For example, $\mathbf{s}^1$ could be the current distribution of schooling attainment in India and $\mathbf{s}^2$ the distribution in the US. Our problem is that we do not know the production function F.

To start deriving an upper bound for the increase in output per worker that can be generated by additional schooling, denote physical capital per worker by k and note that constant returns to scale and weak concavity of the production function in Eq. (1) imply that changing inputs from $(k^1, \mathbf{s}^1)$ to $(k^2, \mathbf{s}^2)$ generates a change in output per worker $y^2 - y^1$ that satisfies

$$y^2 - y^1 \le F_k\left(k^1, \mathbf{s}^1\right)\left(k^2 - k^1\right) + \sum_{i=0}^{m} F_i\left(k^1, \mathbf{s}^1\right)\left(s_i^2 - s_i^1\right) \tag{2}$$

where $F_k(k^1, \mathbf{s}^1)$ is the marginal product of physical capital given inputs $(k^1, \mathbf{s}^1)$ and $F_i(k^1, \mathbf{s}^1)$ is the marginal product of labor with schooling attainment i given inputs $(k^1, \mathbf{s}^1)$. Hence, the linear expansion of the production function is an upper bound for the increase in output per worker generated by changing inputs from $(k^1, \mathbf{s}^1)$ to $(k^2, \mathbf{s}^2)$.

We will be interested in percentage changes in output per worker and therefore divide both sides of Eq. (2) by y^1,

$$\frac{y^2 - y^1}{y^1} \le \frac{F_k\left(k^1, \mathbf{s}^1\right)k^1}{y^1}\left(\frac{k^2 - k^1}{k^1}\right) + \sum_{i=0}^{m} \frac{F_i\left(k^1, \mathbf{s}^1\right)}{y^1}\left(s_i^2 - s_i^1\right). \tag{3}$$

Assume now that factor markets are approximately competitive. Then Eq. (3) can be rewritten as

$$\frac{y^2 - y^1}{y^1} \le \alpha^1\left(\frac{k^2 - k^1}{k^1}\right) + \left(1 - \alpha^1\right)\left(\sum_{i=0}^{m}\left(\frac{w_i^1}{\sum_{i=0}^{m} w_i^1 s_i^1}\right)\left(s_i^2 - s_i^1\right)\right) \tag{4}$$

where α^1 is the physical capital share in output and w_i^1 is the wage of workers with schooling attainment i given inputs $(k^1, \mathbf{s}^1)$. Since schooling shares must sum up to unity we have $\sum_{i=0}^{m} w_i^1(s_i^2 - s_i^1) = \sum_{i=1}^{m}(w_i^1 - w_0^1)(s_i^2 - s_i^1)$ and $w^1 = w_0^1 + \sum_{i=1}^{m}(w_i^1 - w_0^1)s_i^1$, Eq. (4) becomes

$$\frac{y^2 - y^1}{y^1} \le \alpha^1\left(\frac{k^2 - k^1}{k^1}\right) + \left(1 - \alpha^1\right)\left(\frac{\sum_{i=1}^{m}\left(\frac{w_i^1}{w_0^1} - 1\right)\left(s_i^2 - s_i^1\right)}{1 + \sum_{i=1}^{m}\left(\frac{w_i^1}{w_0^1} - 1\right)s_i^1}\right). \tag{5}$$

Hence, the increase in output per worker that can be generated by additional schooling and physical capital is below a bound that depends on the physical capital income share and the wage premia of different schooling groups relative to a schooling baseline.

2.1. Optimal adjustment of physical capital

In Eq. (5), we consider an arbitrary change in the physical capital intensity. As a result, the upper bound on the increase in output that can be generated by additional schooling may be off because the change in physical capital considered is suboptimal given schooling attainment. We now derive an upper bound that allows physical capital to adjust optimally (in a sense to be made clear shortly) to the increase in schooling. To do so, we have to distinguish two scenarios. A first scenario where the production function is weakly separable in physical capital and schooling, and a second scenario where schooling and physical capital are not weakly separable. In this section we develop the first of these cases, while in the appendix we develop the latter.

Assume that the production function for output can be written as

$$Y = F(K, G(L_0, L_1, \ldots L_m)) \tag{6}$$

with F and G characterized by constant returns to scale and weak concavity. This formulation implies that the marginal rate of substitution in production between workers with different schooling is independent of the physical capital intensity. While this separability assumption is not innocuous, it is weaker than the assumption made in most of the development accounting literature.[4]

We also assume that as the schooling distribution changes from the original schooling distribution $\mathbf{s}^1$ to a schooling distribution $\mathbf{s}^2$, physical capital adjusts to leave the marginal product of capital unchanged, $MPK^2 = MPK^1$. This could be because physical capital is mobile internationally or because of physical capital accumulation in a closed economy.[5] With these two assumptions we can develop an upper bound for the increase in output per worker that can be generated by additional schooling, that depends on the wage premia of different schooling groups only. To see this, note that separability of the production function implies

$$\frac{y^2 - y^1}{y^1} \le \alpha^1\left(\frac{k^2 - k^1}{k^1}\right) + \left(1 - \alpha^1\right)\left(\frac{G\left(\mathbf{s}^2\right) - G\left(\mathbf{s}^1\right)}{G\left(\mathbf{s}^1\right)}\right). \tag{7}$$

The assumption that physical capital adjusts to leave the marginal product unchanged implies that $F_1(k^1/G(\mathbf{s}^1), 1) = F_1(k^2/G(\mathbf{s}^2), 1)$ and therefore $k^2/G(\mathbf{s}^2) = k^1/G(\mathbf{s}^1)$. Substituting in Eq. (7),

$$\frac{y^2 - y^1}{y^1} \le \frac{G\left(\mathbf{s}^2\right) - G\left(\mathbf{s}^1\right)}{G\left(\mathbf{s}^1\right)}. \tag{8}$$

[4] Which assumes that the function F in Eq. (6) is Cobb–Douglas, often based on Gollin's (2002) finding that the physical capital income share does not appear to vary systematically with the level of economic development. In the Appendix we show that our approach can be extended to the case where physical capital displays stronger complementarities with higher levels of schooling.

[5] See Caselli and Feyrer (2007) for evidence that the marginal product of capital is not systematically related to the level of economic development.

Weak concavity and constant returns to scale of G imply, respectively, $G(\mathbf{s}^2)-G(\mathbf{s}^1)\le\sum_{i=0}^{m}G_i(\mathbf{s}^1)(s_i^2-s_i^1)$ and $G(\mathbf{s}^1)=\sum_{i=0}^{m}G_i(\mathbf{s}^1)s_i^1$, where G_i denotes the derivative with respect to schooling level i. Combined with Eq. (7), this yields

$$\frac{y^2-y^1}{y^1}\le\frac{\sum_{i=0}^{m}G_i\left(\mathbf{s}^1\right)\left(s_i^2-s_i^1\right)}{\sum_{i=0}^{m}G_i\left(\mathbf{s}^1\right)s_i^1}=\frac{\sum_{i=1}^{m}\left(\frac{w_i^1}{w_0^1}-1\right)\left(s_i^2-s_i^1\right)}{1+\sum_{i=1}^{m}\left(\frac{w_i^1}{w_0^1}-1\right)s_i^1}\tag{9}$$

where the equality makes use of the fact that separability of the production function and competitive factor markets imply

$$\frac{G_i\left(\mathbf{s}^1\right)}{G_0\left(\mathbf{s}^1\right)}=\frac{F_2\left(k^1,G\left(\mathbf{s}^1\right)\right)G_i\left(\mathbf{s}^1\right)}{F_2\left(k^1,G\left(\mathbf{s}^1\right)\right)G_0\left(\mathbf{s}^1\right)}=\frac{w_i^1}{w_0^1}.\tag{10}$$

Hence, assuming weak separability between physical capital and schooling, the increase in output per worker that can be generated by additional schooling is below a bound that depends on the wage premia of different schooling groups relative to a schooling baseline.

2.2. The upper bound with a constant marginal return to schooling

The upper bound on the increase in output per worker that can be generated by additional schooling in Eq. (9) becomes especially simple when the wage structure entails a constant return to each additional year of schooling, $(w_i-w_{i-1})/w_{i-1}=\gamma$. This assumption is often made in development accounting, because for many countries the only data on the return to schooling available is the return to schooling estimated using Mincerian wage regressions (which implicitly assume $(w_i-w_{i-1})/w_{i-1}=\gamma$). In this case, the upper bound for the case of weak separability between schooling and physical capital in Eq. (9) becomes

$$\frac{y^2-y^1}{y^1}\le\frac{\sum_{i=1}^{m}\left((1+\gamma)^{x_i}-1\right)\left(s_i^2-s_i^1\right)}{1+\sum_{i=1}^{m}\left((1+\gamma)^{x_i}-1\right)s_i}.\tag{11}$$

where x_i is years of schooling corresponding to schooling attainment i (schooling attainment 0 is assumed to entail zero years of schooling).

The upper-bound calculation using Eq. (11) is closely related to analogous calculations in the development accounting literature. In development accounting, a country's human capital is typically calculated as

$$(1+\gamma)^S\tag{12}$$

where S is average years of schooling and the average marginal return to schooling γ is calculated using evidence on Mincerian coefficients.[6]One difference with our approach is therefore that development accounting calculations identify a country's schooling capital with the schooling capital of the average worker, while our upper-bound calculation uses the (more theoretically grounded) average of the schooling capital of all workers. The difference is Jensen's inequality.[7] Another difference is that we use country-specific Mincerian returns while development accounting often uses a common value (or function) for all countries.

2.3. Link to development accounting and graphical intuition

At this point it is worthwhile discussing the relationship between our analysis of schooling's potential contribution to output per worker differences across countries and the analysis in development accounting. Following Klenow and Rodriguez-Claire (1997), development accounting usually assesses the role of schooling for output per worker under the assumption that workers with different schooling are perfect substitutes in production. This assumption has been made because it is necessary to explain the absence of large cross-country differences in the return to schooling when technology is Hick-neutral (e.g. Hendricks, 2002; Klenow and Rodriguez-Claire, 1997). But there is now a consensus that differences in technology across countries or over time are generally not Hicks-neutral and that perfect substitutability among different schooling levels is rejected by the empirical evidence, see Katz and Murphy (1992), Angrist (1995), Goldin and Katz (1998), Autor and Katz (1999), Krusell et al. (2000), Ciccone and Peri (2005), Caselli and Coleman (2006). Moreover, the elasticity of substitution between more and less educated workers found in this literature is rather low (between 1.3 and 2, see Ciccone and Peri, 2005 for a summary).

Hence, the assumption of perfect substitutability among different schooling levels often made in development accounting should be discarded. But this does not mean that the findings in the development accounting literature have to be discarded also. To understand why note that the right-hand side of Eq. (9) – our upper bound on the increase in output per worker generated by more schooling – is exactly equal to the output increase one would have obtained under the assumption that different schooling levels are perfect substitutes in production, $G(L_0,L_1,\ldots,L_m)=a_0L_0+a_1L_1+\ldots+a_mL_m$. Hence, although rejected empirically, the assumption of perfect substitutability among different schooling levels remains useful in that it yields an upper bound on the output increase that can be generated by more schooling.

To develop an intuition for these results, consider the case of just two labor types, skilled and unskilled, and no capital,

$$Y=G(L_U,L_H)\tag{13}$$

where G is taken to be subject to constant returns to scale and weakly concave. Suppose we observe the economy when the share of skilled labor in total employment is s^1 and want to assess the increase in output per worker generated by increasing the skilled-worker share to s^2. The implied increase in output per worker can be written as

$$\begin{aligned}y\left(s^2\right)-y\left(s^1\right)&=G\left(1-s^2,s^2\right)-G\left(1-s^1,s^1\right)\\&=\int_{s^1}^{s^2}\frac{\partial G(1-s,s)}{\partial s}ds\\&=\int_{s^1}^{s^2}[G_2(1-s,s)-G_1(1-s,s)]ds.\end{aligned}\tag{14}$$

Weak concavity of G implies that $G_2(1-s,s)-G_1(1-s,s)$ is either flat or downward sloping in s. Hence, Eq. (14) implies that $y(s^2)-y(s^1)\le[G_2(1-s^1,s^1)-G_1(1-s^1,s^1)](s^2-s^1)$. Moreover, when factor markets are perfectly competitive, the difference between the observed skilled and unskilled wage in the economy $w_H^1-w_U^1$ is equal to $G_2(1-s^1,s^1)-G_1(1-s^1,s^1)$. As a result, $y(s^2)-y(s^1)\le(w_H^1-w_U^1)(s^2-s^1)$. As $(w_H^1-w_U^1)(s^2-s^1)$ is also the output increase one would have obtained under the assumption that the two skill types are perfect substitutes, it follows that our upper bound is equal to the increase in output assuming perfect substitutability between

[6] More accurately, human capital is usually calculated as $exp(\gamma S)$, but the two expressions are approximately equivalent and the one in the text is more in keeping with our previous notation.

[7] To see the relation more explicitly, for small γ, $(1+\gamma)^{x_i}$ is approximately linear and the right-hand side of (11) can be written in terms of average years of schooling $S=\sum_{i=1}^{m}x_is_i$, as we do not miss much by ignoring Jensen's inequality and assuming $\sum_{i=0}^{m}(1+\gamma)^{x_i}s_i\approx(1+\gamma)^S$. As a result, if the Mincerian return to schooling is small, the upper bound on the increase in output per worker that can be generated by more schooling depends on the Mincerian return and average schooling only

$$\frac{y^2-y^1}{y^1}\le\frac{(1+\gamma)^{S^2}-(1+\gamma)^{S^1}}{(1+\gamma)^{S^1}}.$$

Another approximation of the right-hand side of (11) for small γ that is useful for relating our upper bound to the development accounting literature is $\gamma(S^2-S^1)/(1+\gamma S^1)$.

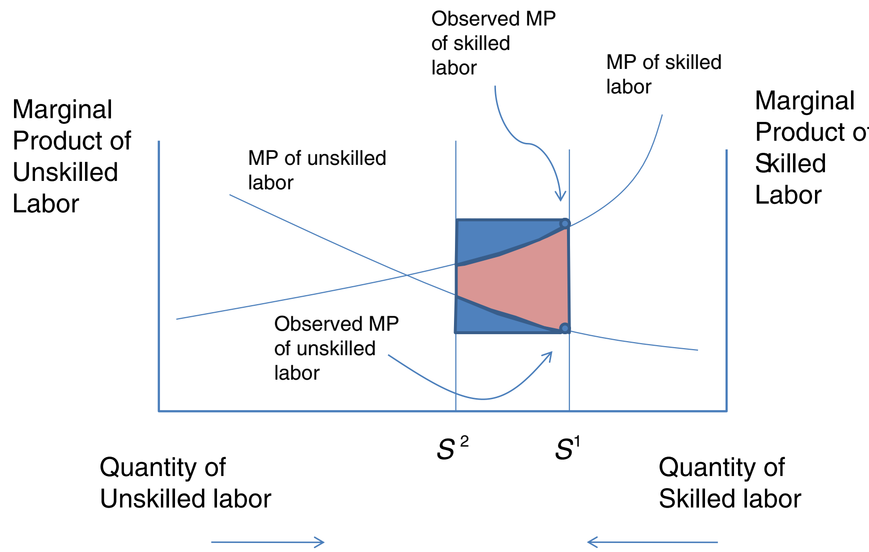

Fig. 1. Change in income from change in schooling.

skill types. Fig. 1 illustrates this calculation graphically.[8] The increase in output is the area between the vertical lines through S_1 and S_2, and the two MP lines. The upper bound is the entire rectangle that contains this area. The figure also illustrates that the difference between our upper bound and the true output gain is larger – making our upper bound less tight – the larger the increase in schooling considered.[9]

It is important to note that once perfect substitutability has been discarded, the particular way the development accounting question is asked becomes very important. There are two main ways that the question has been asked in the literature. The first is the one in Hall and Jones (1999) which is also the focus of our paper: by how much would the output gap shrink if the capital inputs of a factor scarce country increased to the level of a factor rich country? The second way is to follow Klenow and Rodriguez-Claire (1997) and ask what fraction of the variance in income is attributable to variation in a certain factor. Such decompositions are difficult once skill-biased technology and imperfect substitutability among skills are allowed for. What we have shown is that the perfect substitution case provides an upper bound for the answer to the Hall and Jones question.[10]

3. Estimating the upper bounds

We now estimate the maximum increase in output that could be generated by increasing schooling to US levels. We first do this for a subsample of countries and years for which we have data allowing us to perform the calculation in Eq. (9). For these countries we can also compare the results obtained using Eq. (9) with those using Eq. (11), which assume a constant return to extra schooling. These comparisons put in perspective the reliability of the estimates that are possible for larger samples, where only Mincerian returns are available. We also report such calculations for a large cross-section of countries in 1990.

3.1. Using group-specific wages

We implement the upper-bound calculation in Eq. (9) for 9 countries for which we are able to estimate wages by education attainment level using national censa data from the international IPUMS (Minnesota Population Center, 2011). The countries are Brazil, Colombia, Jamaica, India, Mexico, Panama, Puerto Rico, South Africa, and Venezuela, with data for multiple years between 1960 and 2007 for most countries. The details vary somewhat from country to country as (i) schooling attainment is reported in varying degrees of detail across countries; (ii) the concept of income varies across countries; and (iii) the control variables available also vary across countries. See Appendix A Tables 1–3 for a summary of the micro data (e.g. income concepts; number of attainment levels; control variables available; number of observations) and our Supplementary Appendix A for country-by-country data and estimation results. These data allow us to estimate attainment-specific returns to schooling and implement Eq. (9) using the observed country-year specific distribution of educational attainments and the US distribution of educational attainment *in the corresponding year* as the arrival value.

It is worthwhile noting that in implementing Eq. (9) – and also Eq. (11) below – we estimate and apply returns to schooling that vary both across countries and over time. Given our setup, the most immediate interpretation of the variation in returns to schooling would be that there is imperfect substitutability between workers with different schooling attainments and that the supply of different schooling attainments varies over time and across countries. It is exactly the presence of imperfect substitutability among different schooling levels that motivates our upper-bound approach. Another reason why returns to schooling might vary could be that there are differences in technology. Our upper-bound approach does not require us to put structure on such (possibly attainment-specific) technology differences. As we discuss in Section 4, our upper-bound calculation may continue to be correct even under particular ways in which technology changes in response to changes in schooling.[11]

The results of implementing the upper-bound calculation in Eq. (9) for each country-year are presented (in bold face) in Table 1. For this group of countries applying the upper-bound calculation leads to conclusions that vary significantly both across countries and over time. The largest computed upper-bound gain is for Brazil in 1970, which

[8] We thank David Weil for suggesting this figure.

[9] Our implementation of the upper bound below considers US schooling levels as the arrival value. As a result, the increase in schooling considered is large for many developing countries and our upper bound could be substantially larger than the true output gain.

[10] In principle there is a third way of phrasing the question of development accounting: by how much would output in a factor-rich country fall if its factor endowment fell to the level of a factor-poor country. Our main result implies that when the production function is weakly concave in inputs, the decrease in output generated by a fall in schooling is always greater than the decrease predicted under the assumption of perfect substitutability.

[11] Another possible source of differences in schooling returns across countries is sampling variation. However our estimates of both attainment specific and Mincerian returns are extremely precise, so we think that this explanation is unlikely.

Table 1
Upper-bound income increase from moving to US attainment.

	1960	1970	1980	1990	1995	2000	2005
Brazil		**1.576**	**1.201**	**1.020**		**0.901**	
		0.441	0.567	0.304		0.224	
Colombia		**0.901**					
		0.159					
Jamaica			**0.620**	**0.242**		**0.469**	
			0.209	0.076		0.135	
India			**0.908**	**0.945**	**0.769**	**0.792**	**0.769**
			0.053	0.056	0.047	0.054	0.06
Mexico	**1.238**	**0.916**		**0.439**	**0.543**	**0.543**	
	0.524	0.411		0.169	0.187	0.201	
Panama		**0.434**	**0.408**	**0.331**		**0.255**	
		0.088	0.109	0.072		0.055	
Puerto Rico		**0.202**	**0.108**	**0.045**		**−0.003**	**−0.012**
		0.209	0.111	0.061		−0.006	−0.019
South Africa					**0.745**	**0.708**	**0.609**
					0.140	0.129	0.130
Venezuela		**0.757**	**0.604**	**0.403**		**0.860**	
		0.568	0.353	0.132		0.235	

Figures in bold type are changes in income divided by initial income (multiply by 100 to obtain percent change), based on Eq. (9) [i.e. use attainment-specific returns to education].
Figures in normal type are income increases as share of income gap with the US.
1970 figure refers to 1971 for Venezuela and 1973 for Colombia;
1980 figure refers to 1981 for Venezuela, 1982 for Jamaica, and 1983 for India;
1990 figure refers to 1987 for India and 1991 for Brazil and Jamaica;
1995 figure refers to 1993 for India and 1996 for South Africa;
2000 figure refers to 1999 for India and 2001 for Jamaica, South Africa, and Venezuela;
2005 figure refers to 2004 for India and 2007 for South Africa.

is of the order of 150%. This result largely reflects the huge gap in schooling between the US and Brazil in that year (average years of schooling in Brazil was less than 4 in 1970). The smallest upper bound is for Puerto Rico in 2005, which reflects the high schooling attainment achieved by that country in that year (average years of schooling is almost 13). The average is 0.59.

A different metric is the fraction of the overall output gap with the US that reaching US attainment levels can cover. This calculation is also reported in Table 1 (characters in normal type). As a proportion of the output gap, the largest upper-bound gain is for Brazil in 1980 (57%), while the smallest is again for Puerto Rico in 2005 (virtually zero). On average, at the upper bound, attaining the US education distribution allows countries to cover 21% of their output gap with the US.

The shortcoming of the results in Table 1 is that they refer to a quite likely unrepresentative sample. For this reason, we now ask whether using the approach in Eq. (11) leads to an acceptable approximation of Eq. (9). As we show in the next section, data to implement Eq. (11) is readily available for a much larger (and arguably representative) sample of countries, so if Eq. (11) offers an acceptable approximation to Eq. (9) we can be more confident on results from larger samples.

To implement Eq. (11), we first use our micro data to estimate Mincerian returns for each country-year. This is done with an OLS regression using the same control variables employed to estimate the attainment-specific returns to schooling above.[12]See Appendix A Table 2 for point estimates and standard errors of Mincerian returns for each country-year. Once we have the Mincerian return we can apply Eq. (11) to assess the upper-bound output gains of increasing the supply of schooling (assuming that technology remains unchanged). The results are reported, as a fraction of the results using Eq. (9), in the first row of Table 2 (bold type). This exercise reveals differences between the calculations in Eqs. (9) and (11). On average, the calculation that imposes a constant proportional wage gain yields only 77% of the calculation that uses attainment-specific returns to schooling. Therefore, the first message from this comparison is that, on average, basing the calculation on Mincerian coefficients leads to a significant underestimate of the upper-bound output increase associated with attainment gains. However, there is enormous heterogeneity in the gap between the two estimates, and the results from Eq. (11) are not uniformly below those from Eq. (9). Almost one third of the estimates based on Eq. (11) are larger. The significant average difference in estimates and the great variation in this difference strongly suggest that whenever possible it would be advisable to use detailed data on the wage structure rather than a single Mincerian return coefficient. It is interesting to note that the ratio of Eq. (11) to Eq. (9) is virtually uncorrelated with per-worker GDP. To put it differently, while estimates based on (11) are clearly imprecise, the error relative to Eq. (9) is not systematically related to per-worker output. Hence, one may conclude that – provided the appropriate allowance is made for the *average* gap between Eqs. (11) and (9) – some broad conclusions using Eq. (11) are still possible.

We can also compare the results of our approach in Eq. (9) to the calculation combining average years of schooling with a single Mincerian return in Eq. (11). The results are reported in the second rows of Table 2. On average, the results are extremely close to those using Eq. (11), suggesting that ignoring Jensen's inequality is not a major source of error in the calculations. However, the variation around this average is substantial.

3.2. Using Mincerian returns only

The kind of detailed data on the distribution of wages that is required to implement the calculation in Eq. (9) is not often available. However, there are estimates of the Mincerian return to schooling for many countries and years. For such countries, it is possible to implement the approximation in Eq. (11).

We begin by choosing 1990 as the reference year. For Mincerian returns we use a collection of published estimates assembled by Caselli (2010). This starts from previous collections, most recently by Bils and Klenow (2000), and adds additional observations from other countries and other periods. Only very few of the estimates apply exactly to the year 1990, so for each country we pick the estimate prior and closest to 1990. In total, there are approximately 90 countries with an estimate of the Mincerian return prior to 1990. Country-specific Mincerian returns and their date are shown in Appendix A Table 4. For schooling attainment, we use the latest installment of

[12] The empirical labor literature finds that OLS estimates of Mincerian returns to schooling are often close to causal estimates, see Card (1999).

Table 2
Comparison of alternative measures of upper-bound income increase from moving to US attainment.

	1960	1970	1980	1990	1995	2000	2005
Brazil		**0.828**	**0.749**	**0.743**		**0.657**	
		0.816	0.821	0.880		0.773	
Colombia		**0.839**					
		0.873					
Jamaica			**1.052**	**1.269**		**0.439**	
			1.092	1.255		0.431	
India			**0.915**	**0.954**	**0.907**	**0.866**	**0.842**
			1.037	1.100	1.042	1.017	1.000
Mexico	**1.137**	**1.195**		**0.983**	**1.109**	**0.886**	
	1.049	1.105		1.055	1.311	1.024	
Panama		**0.934**	**0.984**	**0.978**		**1.017**	
		1.065	1.202	1.231		1.278	
Puerto Rico		**0.996**	**1.023**	**0.992**		**−1.748**	**0.134**
		1.237	1.285	1.369		−4.333	−0.479
South Africa					**0.711**	**0.612**	**0.694**
					0.861	0.739	0.855
Venezuela		**0.693**	**0.917**	**1.112**		**0.283**	
		0.612	0.958	1.172		0.283	

Figures in bold type are ratios of upper-bound gains assuming constant-returns to schooling years (based on Eq. (11)to upper-bound gains assuming attainment-specific returns (based on Eq. (9)). Figures in normal type further compute the numerator by assigning to all workers the average years of schooling (based on Eq. (12)).
1970 figure refers to 1971 for Venezuela and 1973 for Colombia;
1980 figure refers to 1981 for Venezuela, 1982 for Jamaica, and 1983 for India;
1990 figure refers to 1987 for India and 1991 for Brazil and Jamaica;
1995 figure refers to 1993 for India and 1996 for South Africa;
2000 figure refers to 1999 for India and 2001 for Jamaica, South Africa, and Venezuela;
2005 figure refers to 2004 for India and 2007 for South Africa.

the Barro and Lee data set (Barro and Lee, 2010), which breaks the labor force down into 7 attainment groups, no education, some primary school, primary school completed, some secondary school, secondary school completed, some college, and college completed. These are observed in 1990 for all countries. For the reference country, we again take the US.[13]

Fig. 2 shows the results of implementing Eq. (11) on our sample of 90 countries. For each country, we plot the upper bound on the right side of Eq. (11) against real output per worker in PPP in 1995 (from the Penn World Tables). Not surprisingly, poorer countries experience larger upper-bound increases in output when bringing their educational attainment in line with US levels. The detailed country-by-country numbers are reported in Appendix A Table 4.

Table 3 shows summary statistics from implementing Eq. (11) on our sample of 90 countries. In general, compared to their starting point, several countries have seemingly large upper-bound increases in output associated with attaining US schooling levels (and the physical capital that goes with them). The largest upper bound is 3.66, meaning that output almost quadruples. At the 90th percentile of output gain, output roughly doubles, and at the 75th percentile there is still a sizable increase by three quarters. The median increase is roughly 45%. The average country has an upper-bound increase of 60%.

Fig. 3 plots the estimated upper bounds obtained using Eq. (11) as a percentage of the initial output gap with the US.[14]Clearly the upper-bound output gains for the poorest countries in the sample are small as a fraction of the gap with the US. For the poorest country the upper-bound output gain is less than 1% of the gap with the US. For the country with the 10th percentile level of output per worker, the upper-bound gain covers about 5% of the output gap. At the 25th percentile of the output per worker distribution, the upper-bound gain covers about 7% of the output gap, and at the median it is around 20%. The average upper-bound closing of the gap is 74%, but this is driven by some very large outliers.

[13] To implement Eq. (11) we also need the average years of schooling of each of the attainment groups. This is also available in the Barro and Lee data set.

[14] For the purpose of this figure the sample has been trimmed at an income level of $60,000 because the four countries above this level had very large values that visually dominated the picture.

4. Development accounting and endogenous technology

A possible concern with the approach that characterizes the development accounting literature is that the production function is assumed to be invariant to changes in factor inputs. This may lead development accounting to misjudge the output gap that can be accounted for by input differences. The literature points to two main ways in which inputs may affect the production function. First, there may be a positive external effect of human capital on the overall efficiency of the economy (e.g. Lucas, 1988; Nelson and Phelps, 1966; Romer, 1986). For example, a larger number of schooled workers may make it more likely that the adaptation of an advanced technology to a particular country is profitable, which would lead countries with more schooling to have higher levels of TFP. Second, firms' technology choices may depend on the relative prices of different factors, which in turn depend on relative supplies. Such appropriate-technology considerations may lead the factor bias of the production function to change as the relative supply of workers with different quantities of schooling changes (e.g. Acemoglu, 1998, 2002; Basu and Weil, 1998; Caselli and Coleman, 2006).

We can formalize these concerns as follows. Denote the mapping from a vector of input quantities X to output Y by

$$Y = F(X, \theta(X)) = H(X), \tag{15}$$

where θ is a vector of parameters that depend on X if there are externalities or if technology choice is affected by factor inputs. For example, in models with externalities that work through total factor productivity, the function F may take the form $A(X)G(X)$, where $A(X)$ would capture that total factor productivity changes when factor quantities change. In models of appropriate technology F may be written as $G(A(X)\otimes X)$, where $A(X)$ is a vector of input-specific efficiencies, which in turn may depend on the relative supplies of different inputs.[15]

Development accounting is often understood as asking about the effect of an increase in input quantities X on output Y holding θ

[15] The symbol $\otimes$ denotes the Kronecker product.

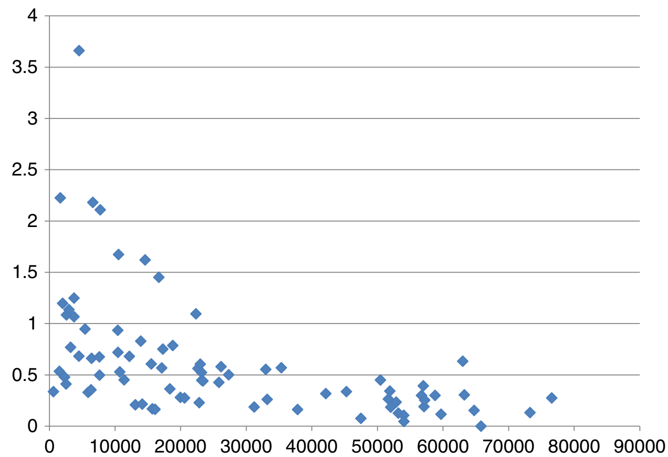

Fig. 2. Upper-bound income increase from moving to US attainment.

constant at the *initial* level of X. That is, development accounting is about quantifying

$$Y^{'} - Y = F\left(X^{'}, \theta(X)\right) - F(X, \theta(X)). \tag{16}$$

This is a well-defined exercise, but strong believers in externalities or appropriate technology may feel that it is of limited practical value if θ changes significantly with X. Such critics would find a calculation of $F(X', \theta(X')) - F(X, \theta(X)) = H(X') - H(X)$ more informative.

Our upper-bound formula is derived for any aggregate production function featuring constant returns to scale and weak concavity. Hence, our upper-bound calculation is robust to endogenous technology if the function $H(X)$ satisfies these restrictions. Perhaps surprisingly, this is sometimes the case. Consider, in particular, the appropriate-technology framework developed and estimated, with considerable empirical success, by Caselli and Coleman (2006). The production function is

$$\left[(A_u L_u)^{\sigma} + (A_s L_s)^{\sigma}\right]^{\frac{1}{\sigma}}, \tag{17}$$

where L_u is unskilled labor, L_s is skilled labor, A_u and A_s are factor-augmenting technology terms, and $1/(1-\sigma)$ is the elasticity of substitution between skilled and unskilled labor. Perfectly competitive firms in each country choose both inputs L_u and L_s and factor-augmenting technology terms A_u and A_s, subject to the production function in (17) and a technology menu given by

$$A_s^{\omega} + \gamma A_u^{\omega} \leq B, \tag{18}$$

where γ, ω, and B are exogenous parameters. Under the parameter restriction $\omega > 1/(1-\sigma)$, which is consistent with Caselli and Coleman's estimates, it can be shown that the optimal technology choice of firms is given by

$$A_u = \frac{B}{\left(1 + \gamma^{\frac{\sigma}{\sigma-\omega}} L_s^{\frac{\omega\sigma}{\omega-\sigma}} L_u^{\frac{\omega\sigma}{\sigma-\omega}}\right)^{\frac{1}{\omega}}} \text{ and } A_s = \frac{B/\gamma}{\left(1 + \gamma^{\frac{\sigma}{\omega-\sigma}} L_u^{\frac{\omega\sigma}{\omega-\sigma}} L_s^{\frac{\omega\sigma}{\sigma-\omega}}\right)^{\frac{1}{\omega}}}. \tag{19}$$

Plugging Eq. (19) into the production function in Eq. (17) we obtain the equivalent of $H(X)$, or the full mapping from inputs to outputs

Table 3
Upper-bound income increase in a large cross section assuming constant returns to each additional year of schooling.

	Mean	Max	90th Percentile	75th Percentile	Median
% Income increase based on (11)	0.61	3.66	1.20	0.68	0.45
% Income increase based on (12)	0.80	7.59	1.48	0.82	0.54

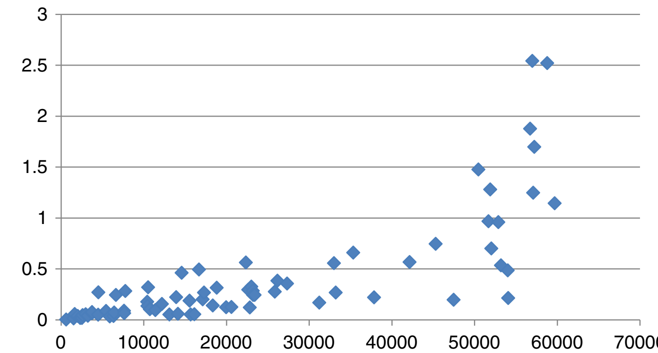

Fig. 3. Upper-bound income increase as percent of income gap with the US.

when endogenous technology is accounted for. It can be shown that this function features constant returns to scale and is concave, which implies that, at least in this case, our proposed approach still delivers an upper bound on the increase in income associated with a certain increase in schooling.

On the other hand, our upper-bound approach will generally not work if there are schooling externalities that induce aggregate increasing returns. This is clearly a limitation of our approach (and development accounting in general). On the other hand, contrary to the case of appropriate technology, the evidence for quantitatively large aggregate schooling externalities is not very strong, suggesting that such externalities are unlikely in practice to significantly affect our quantitative findings.[16] Instrumental-variables approaches suggest that there are no significant aggregate externalities to high-school attainment (e.g. Acemoglu and Angrist, 2001; Ciccone and Peri, 2006; Iranzo and Peri, 2009). And while there do appear to be some aggregate externalities to college attainment (Iranzo and Peri, 2009; Moretti, 2004), they seem to be too small to overturn our main conclusion. According to Iranzo and Peri, an additional year of schooling due to college attainment raises total factor productivity by around 5%. For the typical poor country, taking college attainment to the level of the US in 1990 would add less than 4 years to average years of schooling. Hence, schooling externalities would add around 30% to our upper bound once the induced increase in the physical capital intensity is accounted for.[17] While this is not negligible, it remains too small to significantly increase the fraction of the output gap being closed relative to our calculations.

5. Conclusion

How much of the output gap with rich countries can developing countries close by increasing their quantity of schooling? Our approach has been to look at the best-case scenario: an upper bound for the increase in output that can be achieved by more schooling. The main advantage of our approach is that the upper bound is valid for an arbitrary number of schooling levels with arbitrary patterns of substitution/complementarity. Another advantage is that the upper bound is robust to certain forms of endogenous technology response to changes in schooling. Application of our upper-bound calculations to two different data sets yields output gains from reaching a distribution of schooling attainment similar to the US that are sizeable as a proportion of initial output. However, these gains are much smaller when measured as a proportion of the existing output gap with the US. This result is in line with the conclusions from the development accounting literature, which is not surprising as many development accounting studies assume that workers with different schooling

[16] For a review of evidence on schooling externalities at the microeconomic level, see Rosenzweig (2012).

[17] This calculation assumes that the elasticity of output with respect to physical capital is 0.33, see Gollin (2002).

attainment are perfect substitutes and therefore end up employing a formula that is very similar to our upper bound.

Acknowledgments

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Appendix A. Non-separability between physical capital and schooling

Since Griliches (1969) and Fallon and Layard (1975), it has been argued that physical capital displays stronger complementarities with high-skilled than low-skilled workers (see also Caselli and Coleman, 2002, 2006; Duffy et al., 2004; Krusell et al., 2000). In this case, schooling may generate additional productivity gains through the complementarity with physical capital. We therefore extend our analysis to allow for capital-skill complementarities and derive the corresponding upper bound for the increase in output per worker that can be generated by additional schooling.

To allow for capital–skill complementarities, suppose that the production function is

$$Y = F(Q[U(L_0, .., L_{\tau-1}), H(L_\tau, .., L_m)], G[K, H(L_\tau, .., L_m)]) \tag{20}$$

where F, Q, U, and H are characterized by constant returns to scale and weak concavity, and G by constant returns to scale and $G_{12} > 0$ to ensure capital–skill complementarities. This production function encompasses the functional forms employed by Fallon and Layard (1975), Krusell et al. (2000), Caselli and Coleman (2002, 2006), and Goldin and Katz (1998) for example (who assume that F, G are constant-elasticity-of-substitution functions, that $Q(U,H) = U$, and that U, H are linear functions).[18] The main advantage of our approach is that we do not need to specify functional forms and substitution parameters, which is notoriously difficult (e.g. Duffy et al., 2004).

To develop an upper bound for the increase in output per worker that can be generated by increased schooling in the presence of capital–skill complementarities, we need an additional assumption compared to the scenario with weak separability between physical capital and schooling. The assumption is that the change in the schooling distribution from $\mathbf{s}^1$ to $\mathbf{s}^2$ does not strictly lower the skill ratio H/U, that is,

$$\frac{H\left(\mathbf{s}_2^2\right)}{U\left(\mathbf{s}_1^2\right)} \geq \frac{H\left(\mathbf{s}_2^1\right)}{U\left(\mathbf{s}_1^1\right)}, \tag{21}$$

where $\mathbf{s}_1 = [s_0, \ldots, s_{\tau-1}]$ collects the shares of workers with schooling levels strictly below τ and $\mathbf{s}_2 = [s_\tau, \ldots, s_m]$ collects the shares of workers with schooling levels equal or higher than τ (we continue to use the superscript 1 to denote the original schooling shares and the superscript 2 for the counterfactual schooling distribution). For example, this assumption will be satisfied if the counterfactual schooling distribution has lower shares of workers with schooling attainment $i < \tau$ and higher shares of workers with schooling attainment $i \geq \tau$. f U, H are linear function as in Fallon and Layard (1975), Krusell et al. (2000), Caselli and Coleman (2002, 2006), and Goldin and Katz (1998), the assumption in (21) is testable as it is equivalent to

$$\frac{\sum_{i=0}^{\tau-1} \frac{w_i^1}{w_0^1}\left(s_i^2 - s_i^1\right)}{\sum_{i=0}^{\tau-1} \frac{w_i^1}{w_0^1} s_i^1} \leq \frac{\sum_{i=\tau}^{m} \frac{w_i^1}{w_\tau^1}\left(s_i^2 - s_i^1\right)}{\sum_{i=\tau}^{m} \frac{w_i^1}{w_\tau^1} s_i^1}, \tag{22}$$

where we used that competitive factor markets and Eq. (20) imply $w_i^1/w_0^1 = F_1 Q_1 U_i / F_1 Q_1 U_0 = U_i/U_0$ for $i < \tau$ and $w_i^1/w_\tau^1 = (F_1 Q_2 + F_2 G_2) H_i / (F_1 Q_2 + F_2 G_2) H_\tau = H_i/H_\tau$ for $i \geq \tau$.

It can now be shown that the optimal physical capital adjustment implies

$$\frac{k^2 - k^1}{k^1} \leq \frac{H\left(\mathbf{s}_2^2\right) - H\left(\mathbf{s}_2^1\right)}{H\left(\mathbf{s}_2^1\right)}. \tag{23}$$

To see this, note that the marginal product of capital implied by Eq. (20) is

$$MPK = F_2\left(1, \frac{G\left[\frac{k}{H(\mathbf{s}_2)}, 1\right]}{Q\left[\frac{U(\mathbf{s}_1)}{H(\mathbf{s}_2)}, 1\right]}\right) G_1\left[\frac{k}{H(\mathbf{s}_2)}, 1\right]. \tag{24}$$

Hence, holding k/H constant, an increase in H/U either lowers the marginal product of capital or leaves it unchanged. As a result, k/H must fall or remain constant to leave the marginal product of physical capital unchanged, which implies Eq. (23).

Using steps that are similar to those in the derivation of (9) we obtain

$$\frac{U\left(\mathbf{s}_1^2\right) - U\left(\mathbf{s}_1^1\right)}{U\left(\mathbf{s}_1^1\right)} \leq \frac{\sum_{i=0}^{\tau-1} \frac{w_i^1}{w_0^1}\left(s_i^2 - s_i^1\right)}{\sum_{i=0}^{\tau-1} \frac{w_i^1}{w_0^1} s_i^1}, \tag{25}$$

where we used $w_i^1/w_0^1 = (F_1 Q_1 U_i)/(F_1 Q_1 U_0) = H_i/H_\tau$ for $i < \tau$, and

$$\frac{k^2 - k^1}{k^1} \leq \frac{H\left(\mathbf{s}_2^2\right) - H\left(\mathbf{s}_2^1\right)}{H\left(\mathbf{s}_2^1\right)} \leq \frac{\sum_{i=\tau}^{m} \frac{w_i^1}{w_\tau^1}\left(s_i^2 - s_i^1\right)}{\sum_{i=\tau}^{m} \frac{w_i^1}{w_\tau^1} s_i^1}, \tag{26}$$

where we used $w_i^1/w_\tau^1 = (F_1 Q_2 H_i + F_2 G_2 H_i)/(F_1 Q_2 H_\tau + F_2 G_2 H_\tau) = H_i/H_\tau$ for $i \geq \tau$ and (23). These last two inequalities combined with (20) imply

$$\frac{y^2 - y^1}{y^1} \leq \beta^1 \left(\frac{\sum_{i=0}^{\tau-1} \frac{w_i^1}{w_0^1}\left(s_i^2 - s_i^1\right)}{\sum_{i=0}^{\tau-1} \frac{w_i^1}{w_0^1} s_i^1}\right) + \left(1 - \beta^1\right)\left(\frac{\sum_{i=\tau}^{m} \frac{w_i^1}{w_\tau^1}\left(s_i^2 - s_i^1\right)}{\sum_{i=\tau}^{m} \frac{w_i^1}{w_\tau^1} s_i^1}\right), \tag{27}$$

where β^1 is the share of workers with schooling levels $i < \tau$ in aggregate income. Hence, with capital-skill complementarities, the increase in output per worker that can be generated by additional schooling is below a bound that depends on the income share of workers with schooling levels $i < \tau$ and the wage premia of different schooling groups relative to two schooling baselines (attainment 0 and attainment τ).

To get some intuition on the difference between the upper bound in (9) and in (27), note that the upper bound in (27) would be identical to the upper bound in (9) if, instead of β^1, we were to use the share of workers with schooling levels $i < \tau$ in aggregate wage income. As the share of workers with low schooling in aggregate wage income is greater than their share in aggregate income, (27) puts less weight on workers with low schooling and more weight on workers with more schooling than (9) (except if there is no physical capital). This is because of the stronger complementarity of better-schooled workers with physical capital.[19]

[18] Duffy et al. (2004) argue that a special case of the formulation in (20) fits the empirical evidence better than alternative formulations for capital–skill complementarities used in the literature.

[19] The main difficulty in estimating β^1 is defining threshold schooling τ. If τ was college attainment, the upper bound could be quite large because developing countries have very low college shares and the increase in college workers would be weighted by the physical capital income share plus the college-worker income share (rather than the much smaller college-worker income share only). If τ is secondary school, the difference with our calculations would be small.

Appendix Table 1
Description of individual-level data.

Brazil	*Income concept used in the analysis*: total income per hour worked for 1980, 1991, 2000; total income for 1970. *Other income concepts available*: earned income per hour worked for 1980, 1991, 2000 (yield nearly identical results as income concept used for 1991 and 2000 but a significantly negative return to schooling in 1980). *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (state) of birth, dummies for region (state) of residence, dummy for urban area, dummy for foreign born, dummies for religion, dummies for race (except 1970). *Educational attainment levels*: 8
Colombia	*Income concept used in the analysis*: total income for 1973. *Other income concepts available*: none. *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (state) of birth, dummies for region (municipality) of residence, dummy for urban area, dummy for foreign born. *Educational attainment levels*: *9*
India	*Income concept used in the analysis*: wage income for 1983, 1987, 1993, 1999, 2004. *Other income concepts available*: none. *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (state) of residence, dummy for urban area, dummies for religion. *Educational attainment levels*: 8
Jamaica	*Income concept used in the analysis*: wage income for 1982, 1991, 2001. *Other income concepts available*: none. *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (parish) of birth, dummies for region (parish) of residence, dummy for foreign born, dummies for religion, dummies for race. *Educational attainment levels*: *7*
Mexico	*Income concept used in the analysis*: earned income per hour worked for 1990, 1995, 2000; earned income for 1960; total income for 1970. Other income concepts available: total income per hour for 1995, 2000. Control variables used in the analysis: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (state) of birth, dummies for region (state) of residence, dummy for urban area, dummy for foreign born, dummies for religion (except 1995). *Educational attainment levels*: *10*
Panama	Income concept used in the analysis: wage income per hour worked for 1990, 2000; wage income for 1970; total income per hour worked for 1980. Other income concepts available: earned income per hour worked for 1990, 2000; total income per hour worked for 1990 (yield nearly identical results as income concept used). Control variables used in the analysis: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (state) of birth (except 1990), dummies for region (district) of residence, dummy for urban area (except 1990), dummy for foreign born (except 1980). *Educational attainment levels*: 8
Puerto Rico	*Income concept used in the analysis*: wage income per hour worked for 1970, 1980, 1990, 2000, 2005. *Other income concepts available*: total income per hour worked for 1970, 1980, 1990, 2000, 2005; earned income per hour worked for 1990, 2000, 2005 (yield nearly identical results as income concept used). *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (metropolitan area) of residence, dummy for foreign born, dummies for race (only 2000, 2005). *Educational attainment levels*: 8
South Africa	*Income concept used in the analysis*: total income per hour worked for 1996, 2007; total income for 2001. *Other income concepts available*: none. *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (province) of birth (except 1996), dummies for region (municipality) of residence, dummy for foreign born, dummies for religion (except 2007), dummies for race. *Educational attainment levels*: *6*
Venezuela	*Income concept used in the analysis*: earned income per hour worked for 1971, 1981, 2001; earned income for 1990. *Other income concepts available*: total income per hour worked 2001 (yields a Mincerian return to schooling of 13.7% as compared to 4.4% using earned income). *Control variables used in the analysis*: age, age squared, gender, marital status, age*marital status, gender*marital status, dummies for region (state) of birth, dummies for region (province) of residence, dummy for foreign born. *Educational attainment levels*: *10*

Note: Point estimates of the Mincerian regressions and the number of observations available are summarized in Appendix Tables 2 and 3. For more details on the variables see https://international.ipums.org/international/.

Appendix Table 2
Estimated Mincerian returns and robust standard errors in parentheses.

	1960	1970	1980	1990	1995	2000	2005
Brazil		0.124 (0.00005)	0.113 (0.00004)	0.115 (0.00004)		0.109 (0.00003)	
Colombia		0.0889 (0.0005)					
India			0.083 (0.00002)	0.0866 (0.00002)	0.074 (0.00002)	0.0776 (0.00001)	0.0788 (0.00001)
Jamaica			0.125 (0.002)	0.0573 (0.002)		0.0614 (0.001)	
Mexico	0.123 (0.0002)	0.0993 (0.0001)		0.0682 (0.0001)	0.114 (0.0001)	0.094 (0.0001)	
Panama		0.0879 (0.002)	0.0911 (0.0003)	0.0941 (0.0003)		0.0916 (0.0005)	
Puerto Rico		0.099 (0.0003)	0.088 (0.0005)	0.0938 (0.0005)		0.0985 (0.0005)	0.116 (0.0004)
South Africa					0.117 (0.0001)	0.11 (0.0002)	0.143 (0.0002)
Venezuela		0.0625 (0.0005)	0.0875 (0.0003)	0.0732 (0.0002)		0.0443 (0.0005)	

1970 figure refers to 1971 for Venezuela and 1973 for Colombia;
1980 figure refers to 1981 for Venezuela, 1982 for Jamaica, and 1983 for India;
1990 figure refers to 1987 for India and 1991 for Brazil and Jamaica;
1995 figure refers to 1993 for India and 1996 for South Africa;
2000 figure refers to 1999 for India and 2001 for Jamaica, South Africa, and Venezuela;
2005 figure refers to 2004 for India and 2007 for South Africa.

Appendix Table 3
Number of observations used in the individual-level Mincerian regressions.

	1960	1970	1980	1990	1995	2000	2005
Brazil		14,660,440	24,720,720	33,616,046		41,010,810	
Colombia		3,127,210					
India			86,928,152	45,901,965	109,703,806	133,891,583	139,597,372
Jamaica			255,720	409,100		443,629	
Mexico	4,470,106	6,183,300		14,303,270	18,762,057	21,316,086	
Panama		246,250	367,330	408,540		653,460	
Puerto Rico		653,200	775,220	698,772		732,668	1,000,738
South Africa					6,775,030	8,299,308	9,360,012
Venezuela		1,540,174	2,567,310	3,548,928		5,038,900	

1970 figure refers to 1971 for Venezuela and 1973 for Colombia;
1980 figure refers to 1981 for Venezuela, 1982 for Jamaica, and 1983 for India;
1990 figure refers to 1987 for India and 1991 for Brazil and Jamaica;
1995 figure refers to 1993 for India and 1996 for South Africa;
2000 figure refers to 1999 for India and 2001 for Jamaica, South Africa, and Venezuela;
2005 figure refers to 2004 for India and 2007 for South Africa.

Appendix Table 4
Data and results for the large sample.

	Output in 1995	% Gap with US	Mincerian return		% Gain using (11)	% Gain using (12)	% Gap closed
			Estimate	Year			
Kuwait	76,562	−0.14	4.5	1983	0.275	0.317	−1.95
Norway	73,274	−0.10	5.5	1995	0.132	0.141	−1.29
Zimbabwe	610	106.79	5.57	1994	0.337	0.370	0.00
Uganda	1525	42.13	5.1	1992	0.535	0.572	0.01
Vietnam	2532	24.99	4.8	1992	0.411	0.425	0.02
Ghana	2313	27.44	7.1	1995	0.477	0.578	0.02
Philippines	5897	10.16	12.6	1998	0.330	0.411	0.03
Nepal	2008	31.76	9.7	1999	1.197	1.518	0.04
Sri Lanka	6327	9.40	7	1981	0.355	0.408	0.04
China	3234	19.34	12.2	1993	0.769	0.964	0.04
Zambia	2595	24.35	11.5	1994	1.084	1.342	0.04
Cameroon	4490	13.65	6.45	1994	0.683	0.753	0.05
Peru	13,101	4.02	5.7	1990	0.207	0.239	0.05
Estonia	15,679	3.20	5.4	1994	0.169	0.181	0.05
Russian Federation	16,108	3.08	7.2	1996	0.165	0.172	0.05
Kenya	2979	21.08	11.39	1995	1.135	1.353	0.05
Tanzania	1640	39.10	13.84	1991	2.225	2.676	0.06
Bulgaria	14,140	3.65	5.25	1995	0.214	0.235	0.06
India	3736	16.61	10.6	1995	1.067	1.421	0.06
Bolivia	7624	7.63	10.7	1993	0.498	0.658	0.07
Indonesia	6413	9.26	7	1995	0.661	0.758	0.07
Sudan	3747	16.56	9.3	1989	1.248	1.417	0.08
Nicaragua	5433	11.11	12.1	1996	0.947	1.303	0.09
Honduras	7599	7.66	9.3	1991	0.674	0.763	0.09
Egypt	11,387	4.78	5.2	1997	0.452	0.511	0.09
Dominican Republic	10,739	5.13	9.4	1995	0.528	0.652	0.10
Slovak Republic	22,834	1.88	6.4	1995	0.229	0.265	0.12
Poland	19,960	2.30	7	1996	0.280	0.302	0.12
Croatia	20,606	2.19	5	1996	0.274	0.299	0.13
Paraguay	10,450	5.30	11.5	1990	0.719	0.851	0.14
Costa Rica	18,352	2.58	8.5	1991	0.362	0.411	0.14
El Salvador	12,182	4.40	7.6	1992	0.680	0.776	0.15
Czech Republic	31,215	1.11	5.65	1995	0.186	0.210	0.17
Thailand	10,414	5.32	11.5	1989	0.934	1.084	0.18
Ecuador	15,528	3.24	11.8	1995	0.606	0.820	0.19
Sweden	47,480	0.39	3.56	1991	0.076	0.080	0.20
Panama	17,119	2.84	13.7	1990	0.568	0.770	0.20
Australia	54,055	0.22	8	1989	0.046	0.038	0.21
Cyprus	37,843	0.74	5.2	1994	0.162	0.178	0.22
Tunisia	13,927	3.72	8	1980	0.829	1.006	0.22
Chile	23,403	1.81	12.1	1989	0.442	0.546	0.24
Pakistan	6624	8.93	15.4	1991	2.180	3.439	0.24
Argentina	23,222	1.83	10.3	1989	0.448	0.542	0.24
Korea, Rep.	33,210	0.98	13.5	1986	0.262	0.406	0.27
Botswana	17,280	2.81	12.6	1979	0.751	1.056	0.27
Cote d'Ivoire	4512	13.58	20.1	1986	3.660	7.593	0.27
Mexico	25,835	1.55	7.6	1992	0.426	0.496	0.28
Morocco	7759	7.48	15.8	1970	2.109	3.550	0.28
Malaysia	23,194	1.84	9.4	1979	0.524	0.657	0.29
South Africa	22,638	1.91	11	1993	0.562	0.668	0.29

(continued on next page)

Appendix Table 4 (*continued*)

	Output in 1995	% Gap with US	Mincerian return		% Gain using (11)	% Gain using (12)	% Gap closed
			Estimate	Year			
Colombia	18,808	2.50	14.5	1989	0.787	1.044	0.32
Guatemala	10,530	5.25	14.9	1989	1.674	2.193	0.32
Turkey	22,996	1.86	9	1994	0.605	0.736	0.32
Hungary	27,326	1.41	8.9	1995	0.501	0.588	0.36
Venezuela, RB	26,164	1.51	9.4	1992	0.579	0.689	0.38
Jamaica	14,588	3.51	28.8	1989	1.621	2.268	0.46
Canada	54,026	0.22	8.9	1989	0.106	0.108	0.49
Brazil	16,676	2.95	14.7	1989	1.451	1.903	0.49
Israel	53,203	0.24	6.2	1995	0.126	0.149	0.53
Slovenia	32,991	0.99	9.8	1995	0.553	0.693	0.56
Iran, Islamic Rep.	22,339	1.95	11.6	1975	1.095	1.483	0.56
Greece	42,141	0.56	7.6	1993	0.318	0.368	0.57
Portugal	35,336	0.86	8.73	1994	0.569	0.658	0.66
Denmark	52,032	0.26	5.14	1995	0.185	0.197	0.70
Finland	45,289	0.45	8.2	1993	0.337	0.374	0.74
Ireland	52,868	0.24	9.81	1994	0.234	0.266	0.96
Japan	51,674	0.27	13.2	1988	0.264	0.333	0.97
Netherlands	59,684	0.10	6.4	1994	0.117	0.127	1.14
Hong Kong	57,093	0.15	6.1	1981	0.190	0.229	1.25
United Kingdom	51,901	0.27	9.3	1995	0.342	0.405	1.28
Spain	50,451	0.30	7.54	1994	0.449	0.541	1.48
Switzerland	57,209	0.15	7.5	1991	0.255	0.314	1.70
Austria	56,728	0.16	7.2	1993	0.300	0.331	1.88
France	58,784	0.12	7	1995	0.300	0.347	2.52
Germany	56,992	0.15	7.85	1995	0.392	0.480	2.54
Italy	63,260	0.04	6.19	1995	0.305	0.344	7.63
Belgium	64,751	0.02	6.3	1999	0.154	0.171	9.58
Singapore	63,009	0.04	13.1	1998	0.634	0.724	14.36
United States	65,788	0.00	10	1993	0.000	0.000	n.a.
Iraq	n.a.	n.a.	6.4	1979	0.567	0.664	n.a.
Taiwan	n.a.	n.a.	6	1972	0.330	0.293	n.a.

Note: Output per worker from Penn World Tables.

Appendix B. Supplementary data

Supplementary data to this article can be found online at http://dx.doi.org/10.1016/j.jdeveco.2013.02.006.

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