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Development Accounting and the Rise of TFP

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Abstract

The paper presents evidence that the contribution of differences in total factor productivity (TFP) to income differences across countries steadily increased between 1970 and 2000. We verify that our finding is neither imputable to measurement errors in input factors nor dependent on the assumption of factor neutral differences in technology. We conclude that theories explaining cross-country income differences based on institutions or on forces that are constant over time, such as geography or legal origin, should be reconsidered in the light of their consistency with the rise of the explanatory power of TFP.

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Contents

I. INTRODUCTION

Development accounting seeks to explain cross-country differences in income level. More specifically, development accounting consists in calculating the relative contribution of input factors and total factor productivity (TFP) differences in explaining cross-country differences in income levels. The main finding of the development accounting literature is that TFP differences explain a large fraction of the observed cross-country income differences. Caselli (2005) finds that as much as 60 percent of the observed variance in cross-country income in the year 1996 is associated with TFP differences. The focus on income levels instead of growth rates has strong justification in the literature as laid down in Hall and Jones (1999). Measuring differences in levels captures better long-run differences in welfare, while growth rates differences are considered as transitory. So, if the interesting differences to explain are long-run differences in levels, it is natural that a standard practice in the development accounting literature is to focus on a single year. The contribution of the present paper is to break with such a practice and to investigate how the contribution of TFP differences in explaining cross-country income level differences has evolved through time.

To do so, we perform a standard development accounting exercise using a sample of up to 94 countries from the period 1970 to 2000. The share of TFP differences in explaining income inequality across countries increases steadily over the period studied. We show that measurement errors in input factors, in particular in physical capital, are unlikely to be responsible for our observation. Moreover, we show that departing from the assumption of Cobb-Douglas technology and factor neutrality in efficiency does not alter our finding.

This paper shows that the measure of our "ignorance," or the share of the variance in income explained by the variance in TFP, increased significantly between 1970 and 2000. Although we do not intend to provide a theoretical explanation, we argue that our finding may challenge our understanding of cross-country income level differences by reconsidering the main theories through its lens. In particular, we study the evolution of differences in the quality of institutions and its consistency with a narrative where most TFP variance is explained by differences in the quality of institutions and the rise of TFP inequality.

The remainder of the paper is as follows. Section II briefly describes the methodology and data used to conduct development accounting. Section III describes the results. Section IV discusses the sensitivity of our results to measurement errors in input factors and the introduction of non-factor neutral difference in technology. Section V studies the evolution of differences in the quality of institutions and its implications. Section VI concludes.

II. Measuring Our Ignorance Through Time

This section describes the methodology and data used to conduct development accounting. More detailed presentations are provided by Hall and Jones (1999) and Caselli (2005).

A. Benchmark

Following Hall and Jones (1999), we adopt a benchmark Cobb-Douglas technology such that:

$$Y = AK^{\alpha}(Lh)^{1-\alpha}, \tag{1}$$

where Y is GDP, A is TFP, K is the aggregate capital stock, L is the number of workers, h is the average human capital such that Lh represents the "quality adjusted" workforce. α is a constant in [0,1].

The production function can be rewritten in per-worker terms as:

$$y = Ak^{\alpha}h^{1-\alpha}, \tag{2}$$

where y is GDP per worker and k is the capital-labor ratio, i.e., $k = \frac{K}{L}$.

The goal of the exercise is to determine how much of the variation in income is driven by variation in the factors of production k and h on an annual basis between 1970 and 2000. Therefore, the cross-country data requirement to conduct development accounting for each year is the following: a measure of output; a measure of the stock of physical capital; and a measure of the "quality adjusted" workforce. We also need a value for α.

B. Data

Following Caselli (2005), we construct a measure of the stock of physical capital, output, and the number of workers from version 6.1 of the Penn World Tables [PWT61-Heston, Summers and Aten (2002)]. From Barro and Lee (2001), we obtain educational attainment data.

We measure *y* from PWT61 (real GDP per worker in PPP terms). To compute the stock of capital, *K*, at each point in time, we use the perpetual inventory equation:

$$K_t = I_t + (1 - \delta)K_{t-1}, \quad (3)$$

where I_t is investment measured from PWT61 (real aggregate investment in PPP terms) and δ is the depreciation rate set to 6 percent.[2] We assume that the initial capital stock K_0 is equal to $I_0 /(g+ \delta)$, where I_0 is investment in the first year it is available, and *g* is the geometric average growth rate of investment between the first year the series is available and 1970.[3] Below, we present several types of evidence that this imputation of the initial capital stock does not drive our results.

Following Hall and Jones (1999), we compute a measure of *h* using the formula:

$$h = e^{\Phi(s)}, \quad (4)$$

where *s* is the average years of schooling in the population over 25 years of age, taken from Barro and Lee (2001) and the function $\Phi(s)$ is piecewise linear.[4] The goal of using this functional form is to produce a log-linear relationship between years of schooling and wage given the production function chosen and under perfect competition both in factor and goods

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[2] Both GDP and investment are measured using the chain method. The results are very similar if we use GDP measured using the chain method and investment measured with Laspeyres method as in Caselli (2005).

[3] This method is a standard practice in the literature (see, for example, Hall and Jones, 1999, and Caselli, 2005); The intuition behind it is that $I /(g+ \delta)$ corresponds to capital stock in the steady state of the Solow growth model.

[4] $\Phi(s)=0.134*s$ if $s\leq4$, $\Phi(s)=0.134*4 +0.101*(s-4)$ if $4<s\leq8$, $\Phi(s)=0.134.4+0.101*4+0.068*(s-8)$ if $8<s$.

markets. In addition, Hall and Jones choose *Φ(s)* to be consistent with differences in the return to schooling across countries found in Psacharopoulos (1994).[5]

We assume α to be 1/3 for all countries, which is consistent with U.S. time-series data.

As a result, we construct an annual series for physical capital stock, output, and education for a large number of countries between 1970 and 2000.[6] The sample of countries with data simultaneously available for the three series starts with 88 countries in 1970 and reaches its minimum with 82 countries in 2000. The year 1996 has the maximum number of countries and is therefore the natural choice by Caselli (2005) to perform development accounting.

C. Measuring Our Ignorance

We now have annual data on k, h, and y and made an assumption on the value of α. Annual values of A, as defined in equation (3) are thus the only unknown. We define the factor-only model as follows:

$$y_{factor} = k^{\alpha} h^{1-\alpha}, \qquad (5)$$

$$y = A\, y_{factor}\,, \qquad (6)$$

and A can thus be computed from equation (6) given y and y_{factor}.

Hall and Jones (1999) use an alternative representation of output per capita in terms of capital-output ratio instead of capital-labor ratio:

$$y = A^{\frac{1}{1-\alpha}} \left(\frac{k}{y}\right)^{\frac{\alpha}{1-\alpha}} h\,,$$

such that the associated factor-only model, noted $y^{*}{}_{factor}$, is given by:

$$y^{*}{}_{factor} = \left(\frac{k}{y}\right)^{\frac{\alpha}{1-\alpha}} h.$$

[5] Estimated by Mincerian wage regressions, see Mincer (1974).

[6] Schooling data from Barro and Lee (2001) are available only every five years so we use linear interpolation in between.

In doing so, Hall and Jones (1999) follow Klenow and Rodriguez-Clare (1997) and Mankiw, Romer, and Weil (1992). They justify their choice by essentially two reasons: along a balanced growth path, the capital-output ratio is proportional to the investment rate and therefore the representation has a more natural interpretation; a formulation in terms of capital-labor ratio is unable to discriminate between the effect of an increase in capital and ultimately output owing to a productivity increase and an increase owing to a change in the investment rate. On the other hand, Caselli (2005) favors the formulation in terms of capital-labor ratio for which productivity A does not appear in the factor-only model and hence is more appropriate to assess the effect of productivity differences on the distribution of income across countries. We do not intend to weigh in on the debate. Instead, we use both approaches to show that our main result does not depend on a particular representation of output per capita.

In order to assess the explanatory power of the factor-only model in explaining cross-country income differences, we decompose the variance of the logarithm of y as follows:

$$var[(\log(y)] = var\left[\log(y_{factor})\right] + var[\log(A)] + 2cov\left[\log(A), \log(y_{factor})\right]. \quad (7)$$

Following Caselli (2005), we rely on two indicators of "success" to perform our development accounting exercise. The first indicator is:

$$success1 = \frac{var\left[\log(y_{factor})\right]}{var[(\log(y)]}.$$

The interpretation of indicator success1 is straightforward. It quantifies the extent to which differences in the accumulation of factors measured by the variance of the logarithm of the factor-only model help explain the differences in cross-country income differences measured by the variance of the logarithm of y. In a world where the technology is the same, success1 would take a value of 1.

The second indicator is as follows:

$$success2 = \frac{y_{factor}^{90}/y_{factor}^{10}}{y^{90}/y^{10}}.$$

This alternative measure success2 allows us to address the issue of outliers. It is defined as the 90^{th}-to-10^{th} percentile ratio of the income distribution derived from the factor-only model divided by the 90^{th}-to-10^{th} percentile ratio of the income distribution that is actually observed. In a world where the technology is the same across countries, success2 would take a value of 1.[7]

Alternatively, we measure the indicators of success (noted success1* and success2*) associated with the factor-only model $y^*{}_{factor}$.

III. RESULTS

In this section, we describe the results of our development accounting exercise. First, we present the basic results. Second, we study the decomposition of the success indicators. Third, we perform robustness checks for the changing sample. Finally, we compare the evolution of the success indicators between OECD and non-OECD countries.

Basic Results

Figure 1 describes the evolution of both indicators of success of the factor-only model y_{factor} a la Caselli (2005) using all available data at each period. It should be noted that the benchmark result for the year 1996 in Caselli (2005) is a special case of our present exercise.[8]

The main finding is that the explanatory power of the factor-only model decreases steadily over time. Indeed, in 1970 according to our indicator success1, combining human and physical capital allows us to explain 52 percent of the total variance of the output. In 2000, the factor-only model explains only 39 percent of the variance in output according to

[7] It should be noted that both indicators of success are not bounded by 1 as $var[\log(A)] + 2cov[\log(A), \log(y_{factor})]$ is not constrained to be positive in equation (7).

[8] For the year 1996, our measures of success are very close to Caselli's: success1 is 0.38 and success2 is 0.35. When we use a measure of investment computed with Laspeyres method as in Caselli (2005) and impute schooling data from 1995 we obtain identical results: 0.39 for success1 and 0.34 for success2.

success1. The results are even more pronounced when we use success2 as our measure of the contribution of the factor-only model.

The alternative only-factor model y^*_{factor} a la Hall and Jones (1999) produces a similar pattern. Figure 2 shows that both indicators of success, success1* and success2*, have decreasing trends albeit at systematically lower levels compared with success1 and success2.

Figure 1. The Evolution of the Success Indicators Following Caselli (2005)

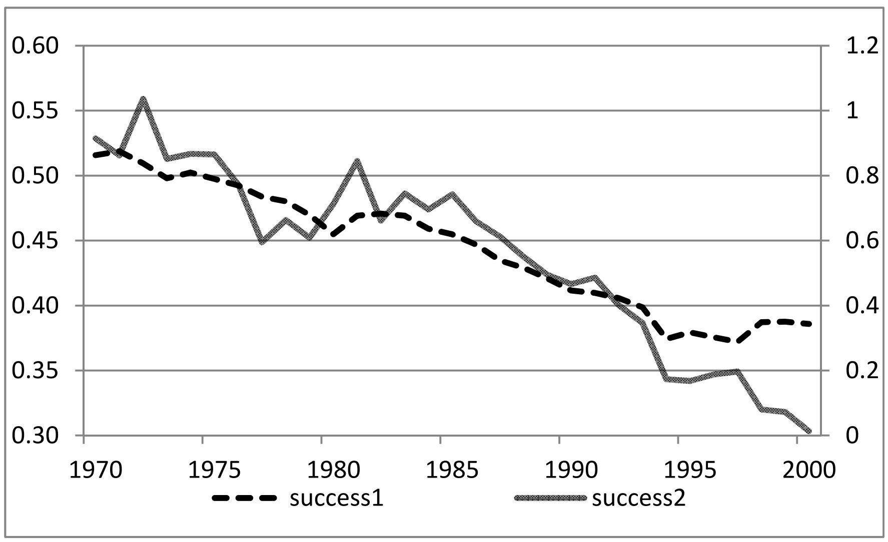
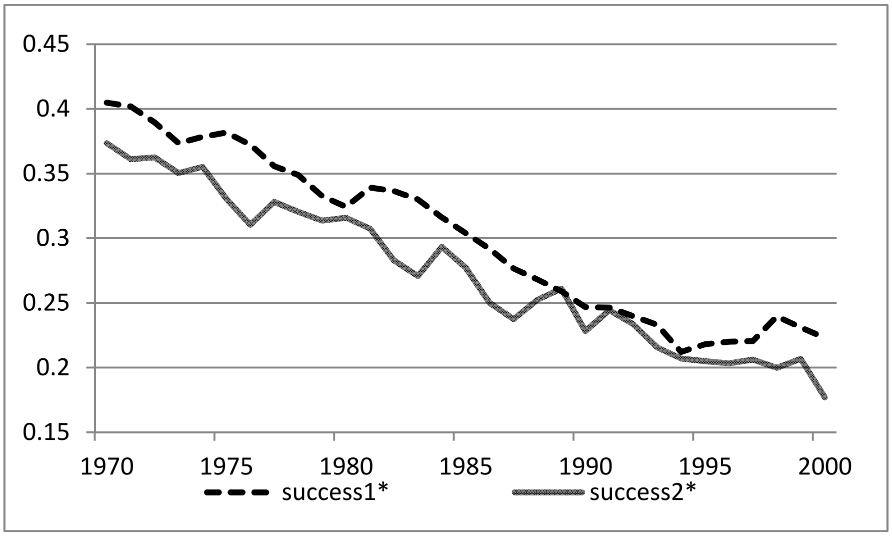


Figure 2. The Evolution of the Success Indicators Following Hall and Jones (1999)



The Decomposition of the Success Indicators

Figure 3 disentangles the success ratios success1 and success1* into the variance of the logarithm of the associated factor-only model (y_{factor} and $y^{*}{}_{factor}$) and the variance of the logarithm of output. It shows that the variance of output has substantially increased over time while the variance of the logarithm of the factor only model has been relatively flat for y_{factor} and decreasing for $y^{*}{}_{factor}$ over the same period. In other words, income divergence has not been matched by factor accumulation divergence.

The relative stability of differences in the factor-only models could hide stark changes in the relative contribution of human and physical capital in explaining per capita income differences. A simple illustration indicates that this is not the case: for each year between 1970 and 2000, we compute, for the five richest and five poorest countries respectively in per capita terms, the geometric average of the GDP per capita.[9] Figure A1 in the appendix portrays ratio of the averages. Figure A1 indicates that the five richest countries were on

[9] The exercise is an extension of the one used in Hall and Jones (1999) to illustrate the magnitudes of physical and human capital contributions to income differences.

average 25 times richer than the five poorest ones in 1970. The ratio reaches 40 in the 1990s. We perform the same exercise for each term of the income per capita representation, k^{α} and $h^{1-\alpha}$ (respectively $\left(\frac{k}{y}\right)^{\frac{\alpha}{1-\alpha}}$ and h), in equation (4) and equation (5) respectively. Figures A2 and A3 in the appendix show that the relative contributions are relatively stable. The figures also show how the relative contributions of the factors shift depending on the representation. In Figure A2, using the representation a la Caselli (2005), the term associated with human capital contributes about 16 percent of total income difference while the term associated with physical capital contributes about 50 percent. In Figure A3, or in the representation a la Hall and Jones (1990), the contribution of human capital is about 20 percent of the total while the term associated with capital-income ratio contributes for about 18 percent of the total.

Figure 3. The Decomposition of the Success Indicators

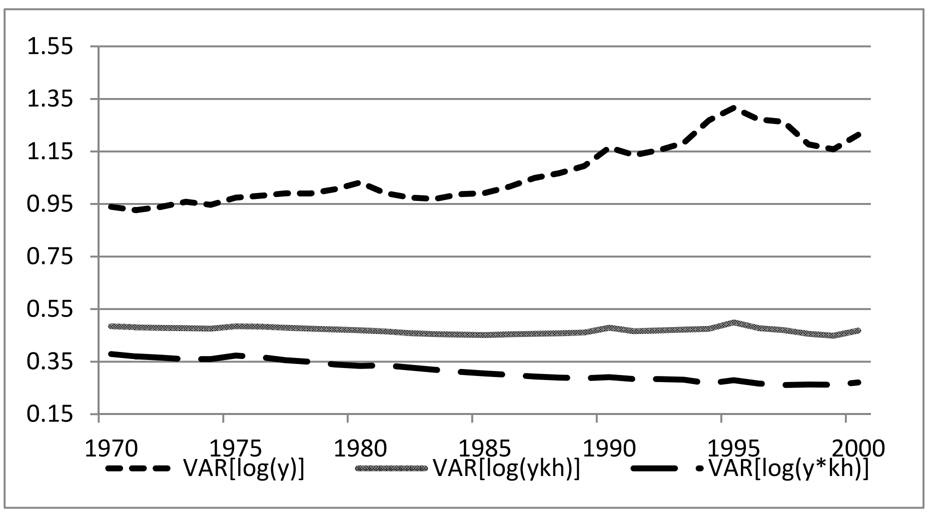


Figure 4. The Variance of the Factors as a Share of their Value in 1970

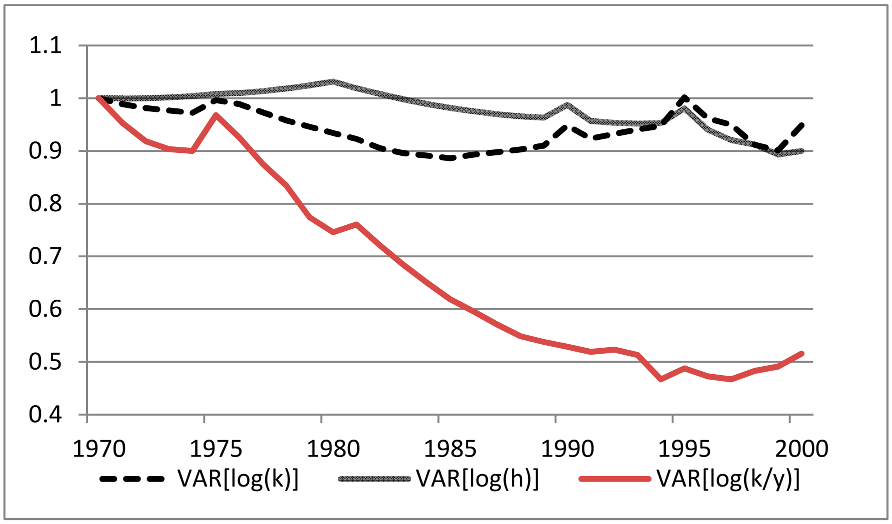


Figure 4 shows the variance of the logarithm of physical and human capital per capita as well as capital-output ratio as a share of their value in 1970. Both the variances of physical and human capital per capita are relatively constant, contributing to the relative stability of the variance of $log\ (y_{factor})$, while the variance of capital-output declines substantially to reach 50 percent of its initial value by the end of the period.

Robustness Check for the Changing Sample

In the above, we have used all available data at each period to address the following question: what would we find if we were to perform a development accounting exercise in 1970 instead of 1996 (or 1971, or 1972...)? We take an alternative approach by confining our attention to the same sample of 87 countries for which data are available throughout the period 1970-96. We find that our results are virtually unchanged for success1 and success1*, while the trends are unchanged for success2 and success2* (see Figure A4 and Figure A5 in the appendix). In the remainder of the paper, we use the approach that maximizes the number of countries covered at each point in time, as it extends to the period 1996-2000 without limiting the sample for previous years.[10] It should also be noted that all the results

[10] The sample size reaches its minimum in 2000 and its maximum in 1996.

presented in the paper are qualitatively similar when using recently released PWT63 dataset [PWT63-Heston, Summers, and Aten (2009)].[11]

OECD vs. non-OECD

Further, we replicate our development accounting exercise separately for two groups of countries, namely OECD and non-OECD countries. Figures 5 and 6 show the evolution of success1 and success1* for the two groups respectively.[12] We find that for both groups of countries, the factor-only model explains less and less of the variance in income per capita following the pattern found for the whole sample of countries. It should be noted that: (i) the trend of the decrease in the success indicator is less pronounced for the OECD group; (ii) the success indicators for the OECD group are greater than for the non-OECD group (iii) and more importantly the success indicators are very close. In contrast, the pattern and magnitude of the success ratios in the non-OECD group are similar to the whole sample (see Figures 1 and 2). The fact that the success indicators are relatively close for the OECD countries is mainly driven by a much lower covariance between the logarithm of the factor model $log\ (y_{factor})$ and the logarithm of TFP, log(A). It can be shown that:

$$y^*{}_{factor} = y_{factor} A^{\frac{\alpha}{\alpha-1}},$$

and therefore:

$$success1^* - success1 = [0.25 * var(\log(A)) - cov(\log(A), \log(y_{factor})]/var(\log(y)).$$

Figure 7 graphs the difference, of each term of the RHS of the equation above, between OECD and non-OECD. Figure 5 shows that the difference between success1 and success1* is smaller in the OECD mainly because the covariance between the logarithm of TFP and the logarithm of the factor model y_{factor} is substantially smaller.[13]

[11] Results are not shown but are available from the authors upon request.

[12] The small sample size of OECD countries makes the use of success2 less appropriate.

[13] Hsieh and Klenow (2010) provide some evidence that TFP affects output not only directly but also indirectly through its impact on factor accumulation.

Figure 5. Success Indicators in OECD Countries

Figure 6. Success Indicators in Non-OECD Countries

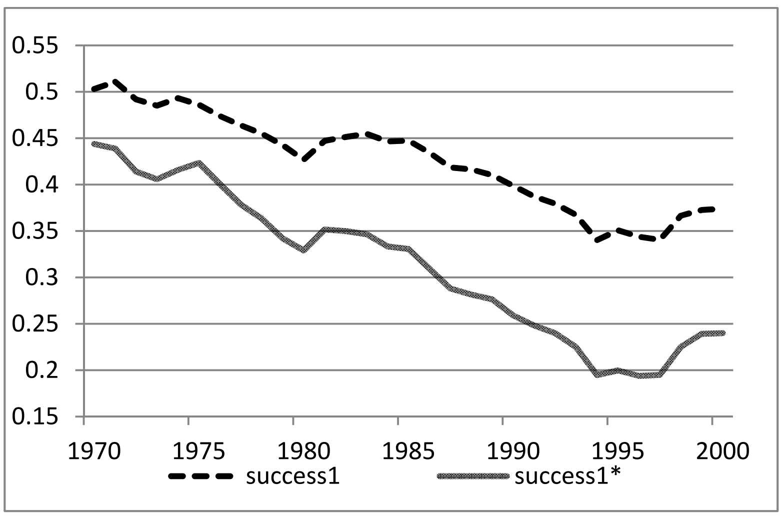
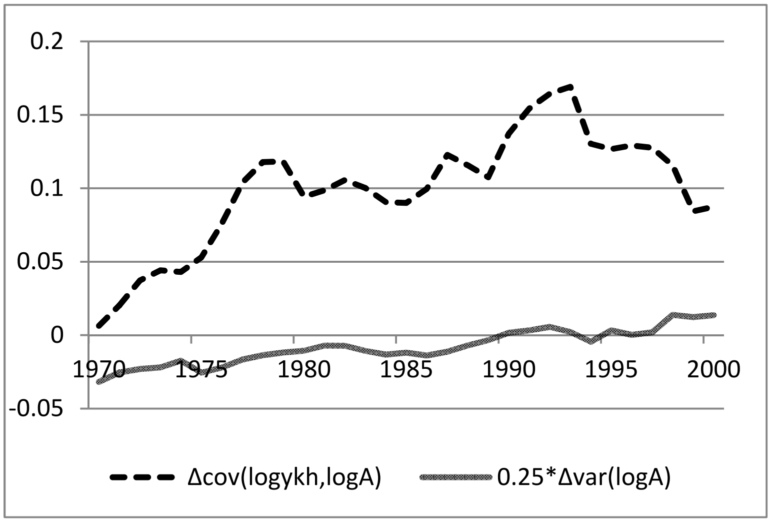


Figure 7. Difference Between OECD and Non-OECD Countries as a Share of var[log(y)]



We have seen the continuous rise of the power of TFP differences in explaining income inequality in a standard development accounting framework. Yet, the trend uncovered may simply stem from measurement errors, in particular in the measurement of the stock of physical capital in the 1970s, or it could hinge on the factor neutrality of the Cobb-Douglas production function.

IV. Is It Measurement Error Or Factor Neutrality?

In this section, we investigate two possible explanations of our central finding, namely measurement error or factor neutrality.

A. Is It Measurement Error?

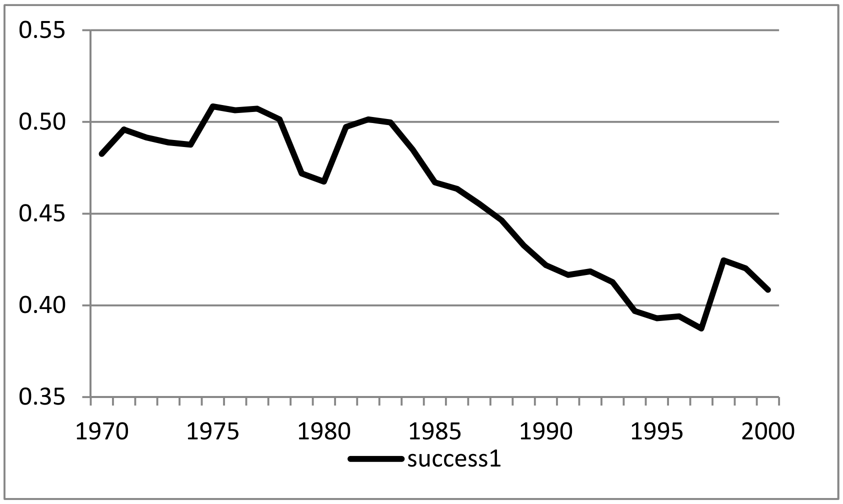
One common criticism of development accounting is targeted toward the perpetual inventory method used to construct the physical capital stock.[14] Caselli (2005) performs robustness checks that lead him to conclude that the baseline result, a measure of "ignorance" around 40

[14] Also, Pritchett (2000) and Pritchett (2006) provide insightful criticisms of both the schooling based measures of human capital and a commonly used measure of physical capital that does not separate out government owned from privately owned capital.

percent in 1996, is unlikely to be revised if initial capital mismeasurement were to be corrected. In the same fashion, we show that it is not realistic to assume that measurement error, in particular in earlier years, is the driver of the secular decrease in the success indicators we observed. We focus here on measurement error associated with physical capital stock. The argument also applies to human capital and income per capita, although there are fewer issues about potential differences between magnitudes of measurement errors in the 1970s as compared with the 2000s. To do so, we rely on three different approaches. First, we limit the sample to countries that have data starting in 1950 and therefore have the "best" measured capital stock in 1970. Our main result is robust to the use of that restricted sample. Second, we show the magnitude of the measurement error needed to fully offset the decreasing trend in the success1 indicator is unrealistic. Third, we show that replacing capital-output ratios by investment rates (which are proportional along the balanced growth path) in the computation of success1* does not affect its value, in particular in the 1970s. As there is much less concern about measurement errors in investment compared with capital stock, we conclude that the decreasing trend of success1* is unlikely to be corrected if exact measures of capital stock were obtained.

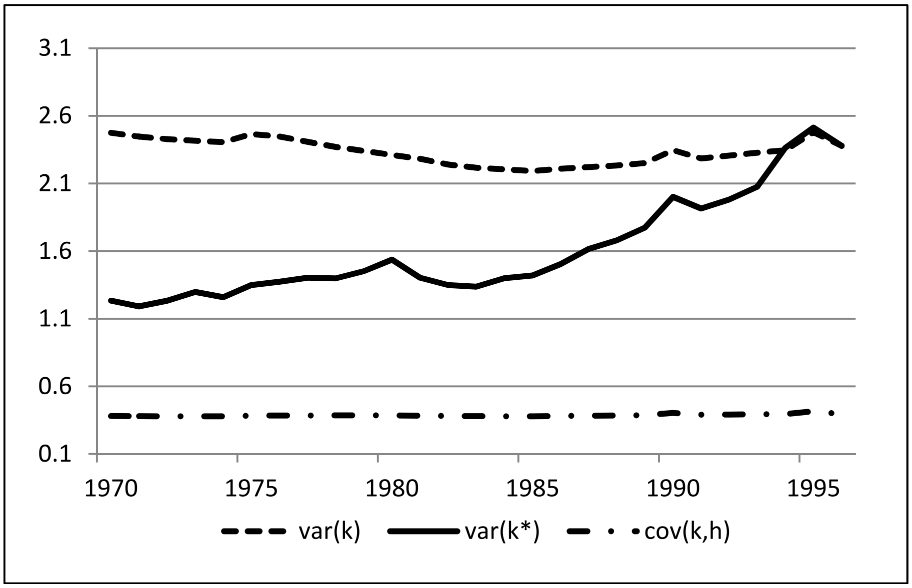
We find that 51 countries have data available on the 1950-2000 period. In particular, investment data on the 1950-1970 period are available and we can produce the most accurate estimation of the stock of physical capital using the PWT6.2 table. We calculate the success indicator success1 using the same 51 countries for each year and we report the results in Figure 8. The sample is too narrow to make the success2 indicator interpretable. Figure 8 shows that the pattern we found for the whole sample remains and that the success indicator decreases from about 50 percent in 1970 to about 40 percent in 2000.

Figure 8. Success1 for 51 Countries for which Data is Available from 1950 Onwards



We calculate the required level of variance of physical capital that would achieve a constant level of success1 and would equal its observed value in 1996. It is labeled *var(k*)*. The idea is to assume that there are no measurement errors in 1996 and if we were to correct for measurement errors in earlier years, we would obtain a constant indicator of success. *var(k*)* is computed assuming that the covariance between physical and human capital is constant and equal to its value in 1996. Figure 9 shows *var(k*);* the observed variance of the logarithm of physical capital *var(k)*; the observed covariance between the logarithm of physical capital and the logarithm of human capital *var(k,h)*. The assumption of a constant covariance is justified by the relative stability of the observed covariance in the most recent years where we assume that it is correctly measured.

Figure 9. Required Variance of Capital to Achieve Constant Success Indicator



We find that the required level of variance of the logarithm of physical capital, *var(k*)*, in 1970 that would achieve a constant success1 is less than half of the measured variance in 1970. The magnitude of the measurement error required to offset the decreasing trend observed is huge. Therefore, even a substantial measurement error leading to an over-estimation of the true variance by 20 percent would not fully offset the decreasing trend of the success indicator. Moreover, *var(k*)* is about 40 percent of the observed value of the observed variance *var(k)* in 1996. It is difficult to believe that cross-country differences in physical capital per capita were halved between 1970 and 2000 for such a large sample.

The rationale for using a representation of income per capita of the form, $y = A^{\frac{1}{1-\alpha}} \left(\frac{k}{y}\right)^{\frac{\alpha}{1-\alpha}} h$, is that the capital-output ratio is proportional to the investment rate along the balanced growth path. In theory, replacing the capital-output *(k/y)* by the investment rate should produce the same value of the success indicators success1* and success2*.

We compute success1** and success2** based on the factor-only model:

$$y^{**}{}_{factor} = \left(\frac{I}{Y}\right)^{\frac{\alpha}{1-\alpha}} h.$$

Figures 10 and 11 show that the success indicators, success1* and success1** and success2* and success2**, are close. The proximity reflects the fact that the proportionality of capital-output and investment rate is a good approximation in the sample. Therefore, the secular decrease in the success indicators is robust to the use of investment rates instead of capital-output ratios, while investment data are more reliable compared with the estimated stock of physical capital.

Figure 10. Success1* Indicator Using Alternative Measures of Capital Stock

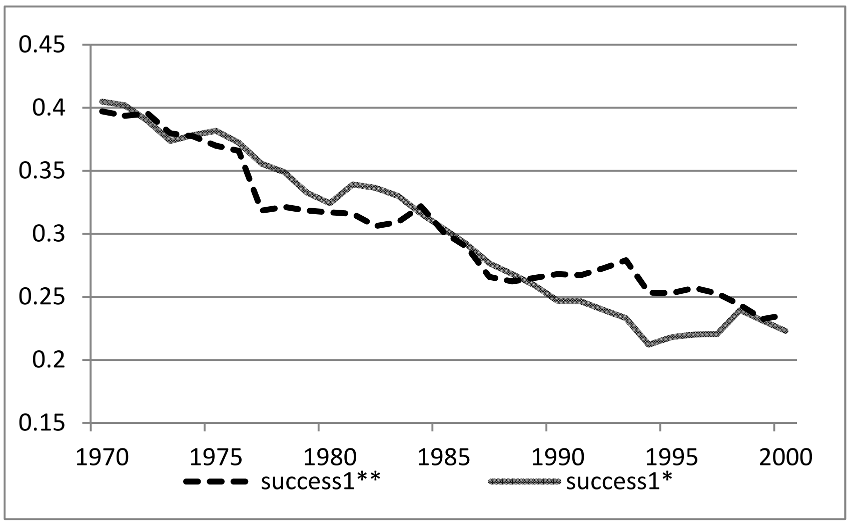


Figure 11. Success2* Indicator Using Alternative Measures of Capital Stock

We have used three different approaches, none of which is an unchallengeable proof. But together they lead us to believe that obtaining the true values of the stock of physical capital will not affect the decreasing trend in the success indicators.

B. Is It Factor Neutrality?

In the following sub-section, we investigate whether or not our main findings depend on the assumption of factor neutral differences in technology[15]. To do so, we replace the Cobb-Douglas production function by a CES (which allows for non-factor neutrality) as follows:

$$Y = [\alpha(A_k K)^\sigma + (1-\alpha)(A_h Lh)^\sigma]^{\frac{1}{\sigma}},\ \alpha \in (0,1),\ \sigma < 1. \quad (8)$$

Where A_k and A_h are the efficiency units associated with one unit of physical capital and one unit of quality-adjusted labor, respectively. Both α and σ are constant across countries. The elasticity of substitution, denoted η, is as follows:

$$\eta = \frac{1}{1-\sigma}.$$

[15] The very same exercise is performed, with a much richer analysis, in Caselli (2005) for a single year. While his goal is to show that the level of the success indicator can be substantially increased by choosing an elasticity of substitution close to 0.5, our goal in this paper is to show that the decreasing trend of the success indicator is unaffected of the choice of the value of the elasticity of substitution.

When η approaches 1 (σ approaching 0), the Cobb-Douglas case used in our development accounting exercise becomes the limit. We assume that factor markets are competitive as in Caselli and Coleman (2005) such that:

$$r = \alpha y^{1-\sigma} k^{\sigma-1} A_k^{\sigma}, \tag{9}$$

$$w = (1-\alpha) y^{1-\sigma} h^{\sigma-1} A_k^{\sigma}.$$

Where *r* is the cost of physical capital and *w* is the cost of quality-adjusted labor. Given parameter values α and σ, we can pin down A_k and A_h. From equations (9), we obtain the following equations:

$$A_k = \left(\frac{rk}{y}\frac{1}{\alpha}\right)^{1/\sigma} \frac{y}{k} = \left(\frac{\pi_k}{\alpha}\right)^{1/\sigma} \frac{y}{k}, \tag{10}$$

$$A_h = \left(\frac{wh}{y}\frac{1}{1-\alpha}\right)^{1/\sigma} \frac{y}{k} = \left(\frac{\pi_h}{1-\alpha}\right)^{1/\sigma} \frac{y}{h}.$$

Where π_k and π_h are respectively the shares of physical and human capital in income. In order to perform our development accounting exercise with this new framework, we calculate the implied income distribution without appealing to differences in technologies. To do so, we infer A_k and A_h for the U.S. from equation (11). Namely, we calculate $A_{k,US} = \left(\frac{\pi_{k,US}}{\alpha}\right)^{1/\sigma} \frac{y_{US}}{k_{US}}$ and $A_{h,US} = \left(\frac{\pi_{h,US}}{1-\alpha}\right)^{1/\sigma} \frac{y_{US}}{h_{US}}$. We then impose for all countries that $A_k = A_{k,US}$ and A_h=$A_{h,US}$ and plug those values in equation (9) to compute the various success indicators. In other words, we calculate the share of the observed variance of income explained by the factor-only model.

Figure 12. Success1 Indicator for a Class of Non-Neutral Production Functions

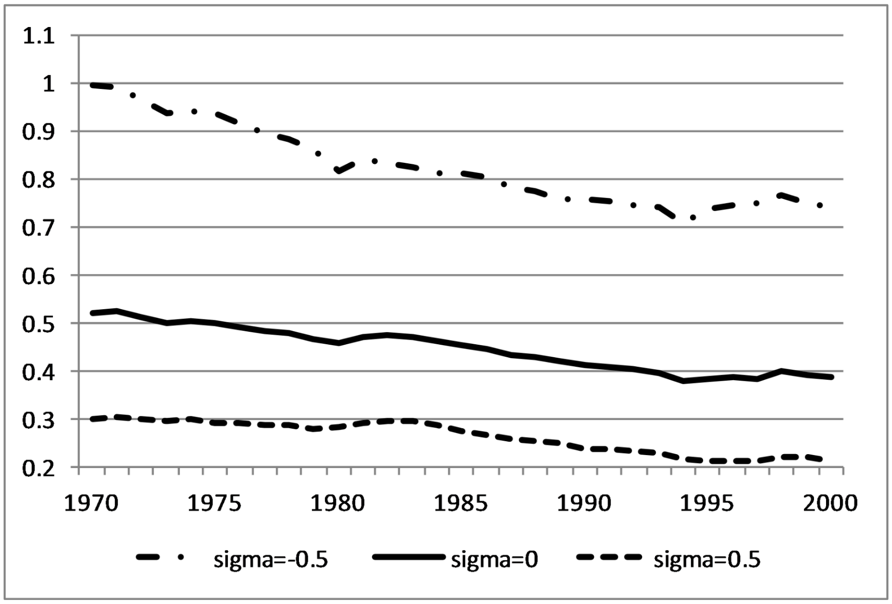


Figure 12 displays the evolution of the indicator of success 1 through time, for a class of production functions parameterized by σ in [-0.5,+0.5].[16] The negative trend seems always present although the level of the explanatory power of the only-factor model and more importantly its rate of change is substantially lower for σ close to 0.5.[17] Therefore, assuming that σ is greater or equal to 0.5 would explain most of our result. However, such an assumption would lead to an extremely high elasticity of substitution (greater than 2) compared to what is usually admitted in the literature (between 0 and 1).[18] Thus, we can conclude that for reasonable values of the elasticity of substitution, the assumption of factor neutral differences in technology is not essential to the main finding of this paper.

[16] Note that for sigma equal to zero, we use the value obtained from the Cobb-Douglas based success indicator.

[17] As noted in Caselli (2005), the success of the factor-only model decreases with the value of the elasticity of substitution. The intuition is that a smaller elasticity of substitution leads to a production function closer to a Leontief. Meanwhile, poor countries are relatively abundant in human capital (in terms of human to physical capital ratio) while the U.S. technology is relatively efficient in its use of human capital. Therefore, a Leontief would "waste" a big share of efficient human capital and produce a low GDP for poor countries and increase the success indicator.

[18] Note that for σ close to 0.5, the explanatory power of the only-factor model would be only 20 percent in 2000.

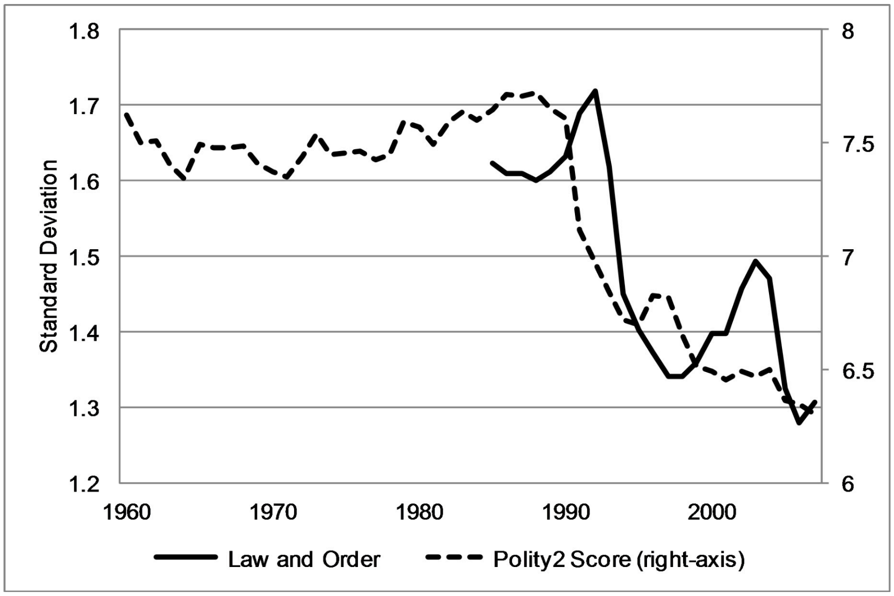
V. Are Cross-Country Differences in the Quality of Institutional Arrangements Consistent with our Finding?

We have shown evidence that the role of TFP in explaining income differences rose over the period 1950-2000. We do not intend to produce a theory to explain our finding. Nonetheless, we take a first look at the consistency between leading theories of income differences and our main finding. In general, any theory relying on predetermined factors cannot be consistent with a secular rise in the explanatory power of TFP over the last decades. In particular, explanations of income per capita differences based on differences in geography (see Sachs and Warner, 2007), ethnic fragmentation (see Easterly and Levine, 1997), and legal origin (see La Porta et al., 1999) are challenged by the rise of TFP.

We turn to an illustrative exercise whose goal is to assess whether cross-country differences in the quality of institutions are consistent with the rising role of TFP in explaining cross-country income differences over the past decades. Building on North (1990), recent research has moved beyond the purely economic factors to consider the role of institutions in shaping economic outcomes. For instance, Keefer and Knack (1995), Quinn and Wooley (2001), and Acemoglu, Johnson, and Robinson (2001) have found relationships between some measures of political institutions and macroeconomic outcomes.[19] In a context of a development accounting exercise, Hall and Jones (1999) provide cross-sectional evidence that the quality of institutional arrangements (aiming at limiting the expropriation risk investors face) have a statistically significant and economically large impact on cross-country TFP differences.

[19] However, Glaeser et al. (2004) argue that convincingly identifying the causal effects of institutions is difficult.

Figure 13. Standard Deviation of Rule of Law Indicator and Polity2 Score



Our main finding is that our main finding is inconsistent with the evolution of the cross-country differences in the quality of institutions, when assuming a stable relationship between institutions and TFP over time. This can be inferred from Figure 13 which shows that the cross-country variance of the indicator of rule of law from Political Risk Services (2009) displays a decrease over the past two decades for which data are available. The observed decrease indicates that cross-country qualities of institutional arrangements have converged in past the decades. Therefore, the evolution of differences in the quality of institutions is a priori not consistent with an increase in cross-country TFP differences.[20]

To circumvent the unavailability of data prior to 1985 from the Political Risk Services (2009), we also resort to using an indicator related to political institutions from Polity IV database (Marshall and Jaggers, 2005). The rationale for using this type of data is that economic institutions could be considered an outcome derived from political institutions.

[20] It should be noted that this results holds whether we restrict the set of countries used in the computation of the variance to be constant over time, or whether we restrict the set of countries to be a subset of the set of countries used in our development accounting exercise. Moreover, the decline in the variance is robust to the use of alternative indicators of the quality of institutions from the Political Risk Services (2009).

Figure 13 displays the evolution of Polity2 score's variance over time.[21] Interestingly, from 1985 onwards the decreasing pattern of Polity2 score's variance is similar to the pattern of the evolution of the variance of the indicator of the rule of law. Prior to 1985, the variance of the Polity2 score is relatively stable.[22] Again, developments related to cross-country differences in political institutions appear inconsistent with a secular rise in the role of TFP differences in explaining cross-country income differences, if we assume a stable relationship between institutions and TFP over time.

As production processes are more complex today compared with 50 years ago, one could reasonably assume that there are relatively more complementarities between the quality of institutions and output today compared with the past. An important direction for research would thus be to test whether the nature of the relationship between potential factors shaping cross-country TFP differences, including institutions, and output has changed over time, so as to reconcile the patterns of development of those factors with the rise of TFP.

VI. Conclusion

The paper has presented evidence that the contribution of differences in total factor productivity (TFP) to income differences across countries steadily increased between 1970 and 2000. To do so, we performed the traditional development accounting exercise for each year of a period covering 30 years. We provided evidence that our main finding neither stems from measurement errors in input factors nor hinges on the assumption of factor neutral technology differences.

The scope of this paper does not include building a theory to explain the rising role of TFP. Nevertheless, we argue that our finding can help deepen our understanding of cross-country income differences by reassessing the main theories in its light. We thus took a first look at

[21] We have rescaled Polity2 score so that it now ranges from 0 to 20 instead of from -10 to 10. A higher score indicates a more democratic system.

[22] The pattern of the evolution of the variance of Polity2 score shown in Figure (13) holds against the same robustness checks performed when using the indicator of rule of law from Political Risk Services (2009).

one of the leading theories of income differences, namely the role of institutional arrangements. We find that the observed convergence in the quality of institutions is inconsistent with the rise of TFP if one assumes a stable relationship between the institutional arrangement and output. Meanwhile, such a change in the relationship is plausible. As production processes are more complex today compared with 50 years ago, one could reasonably assume that there are relatively more complementarities between the quality of institution and output today compared with the past. More generally, any theory relying on constant forces, such as geography, ethnic fractionalization, or legal origin, cannot be consistent with a secular rise in the explanatory power of TFP over the last decades. An important direction for research would thus be to test whether the nature of the relationship between potential factors shaping cross-country TFP differences, including institutions, and output has changed over time. As countries seem to converge in terms of the quality of institutions, the change in the relationship would have to be swift enough to remain consistent with the rise of TFP.

We know from Hall and Jones (1999) and Caselli (2005) that what we ignore exceeds what we know about cross-country income difference based solely on differences in input factors. It seems that we also know less and less, which is yet to be explained.

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APPENDIX

Figure A1. Ratio of Five Poorest to Five Richest Countries' Geometric Averages of Income Per Capita

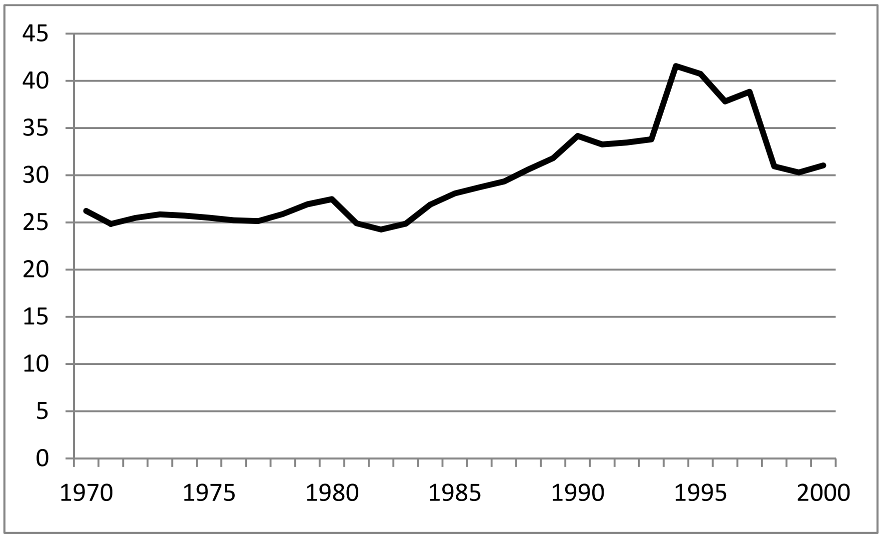

Figure A2. Decomposition of the Contribution of y_{factor} to Income Inequality

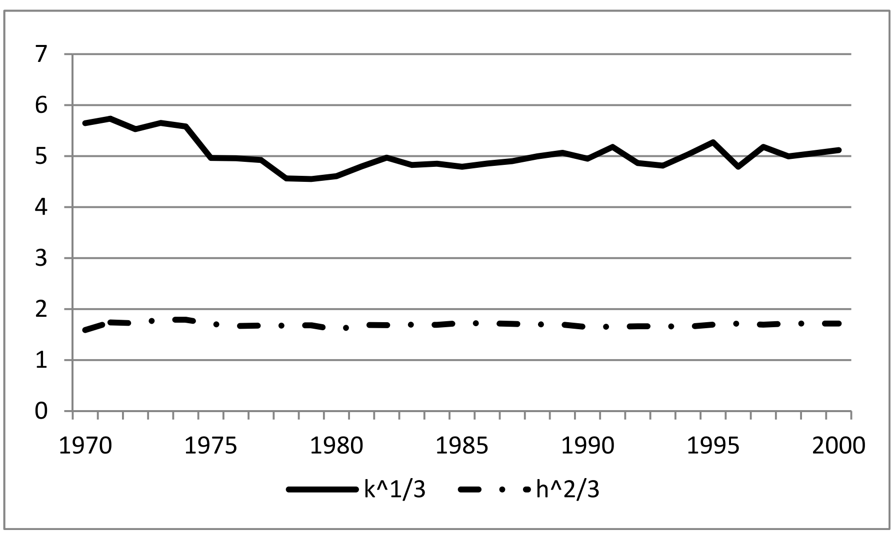

Figure A3. Decomposition of the Contribution of y^*_{factor} to Income Inequality

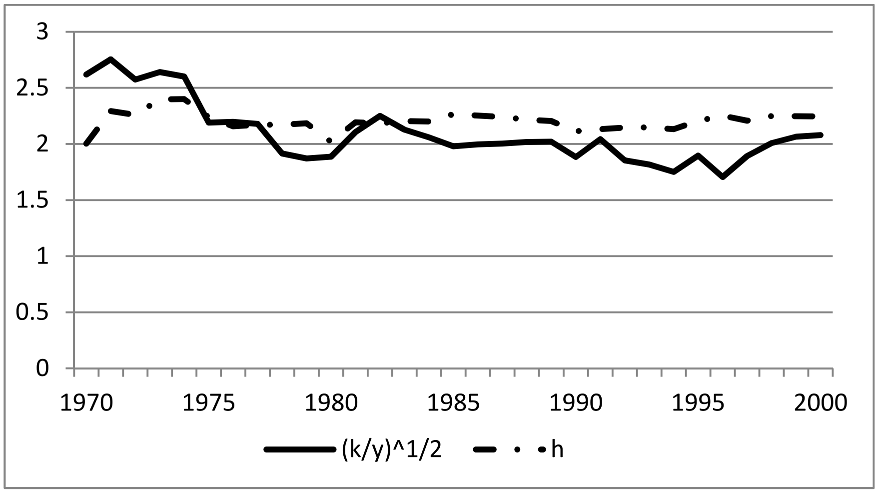

Figure A4. Success1 and Success2 with Fixed Sample

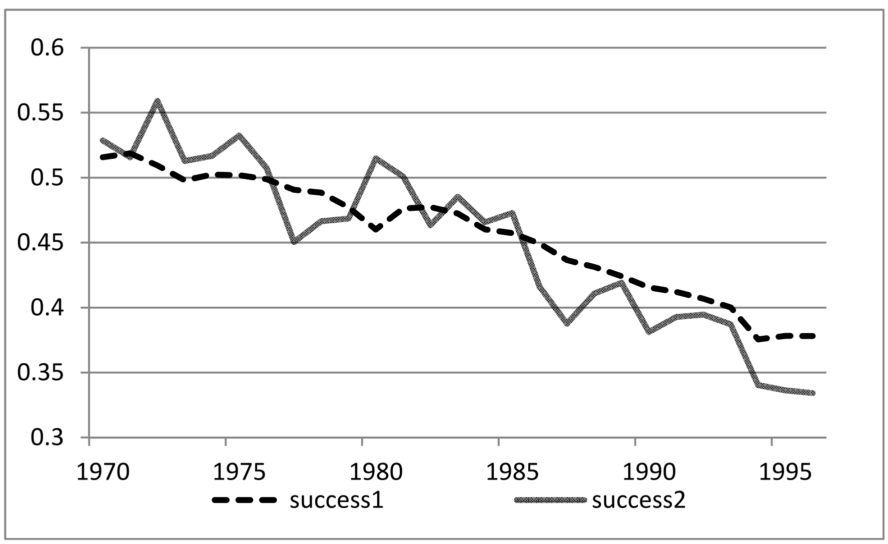

Figure A5. Success1* and Success2* with Fixed Sample