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New ways of looking at old issues: inequality and growth

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Abstract

The paper uses new cross-country data on income and asset (land) distribution to show that (i) there is a strong negative relationship between initial inequality in the asset distribution and long-term growth; (ii) inequality reduces income growth for the poor, but not for the rich; and (iii) available longitudinal data provide little support for the Kuznets hypothesis. Policies that increase aggregate investment and facilitate acquisition of assets by the poor might thus be doubly beneficial for growth and poverty reduction. © 1998 Elsevier Science B.V. All rights reserved.

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1. Introduction

This paper uses two new data sets to examine the interactions between growth and inequality and how they in turn affect efforts to reduce poverty in the course of economic development. While these are well-researched issues, two factors motivate a fresh empirical look. First, the data on inequality that have been used in many of the existing studies are of doubtful quality. We therefore re-examine these questions using what we believe are better and more comprehensive data. Second, and more importantly, our new data sets allow us to look at these issues in new ways. Thus, we are able to utilize data on the distribution of *land* as a proxy for the distribution of assets rather than measures of *income* distribution that have

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traditionally been used to substantiate a negative relationship between initial inequality and growth. We also are able incorporate at least some *intertemporal* element into the analysis of the aggregate inequality–income relationship as summarized in Kuznets' hypothesis, in contrast to past efforts that have relied on cross-country information. And, we make use of data on *income shares* to measure the change in income of the bottom 20 or 40% of the population (as well as higher quintiles). This allows us to recognize the fact that growth and the distribution of income evolve simultaneously and, for example, examine directly the impact of growth–inequality interactions on the poorest groups in society. The results can be used to make more specific inferences regarding the effect of initial conditions on the poor, thus generating information that could be of greater policy relevance than analyses relying on aggregate relationships. While we draw on the theoretical literature to justify the approach chosen and the variables used, the paper's main contribution is clearly empirical.

The paper has four sections. In Section 2, we briefly detail the two data sets—one on income inequality and one on land inequality—used in this paper and highlight some deficiencies of the data sets used in much of the empirical analysis to date.

Section 3 explores the possibility of a systematic relationship between *initial* inequality and subsequent growth. If confirmed, this would imply that unequal economies will experience lower rates of growth and—assuming that all groups in society benefit equally from aggregate growth—lower rates of poverty reduction. While a negative relationship between initial inequality and subsequent growth has recently been confirmed in the literature, the inequality data used are deficient with respect to their quality, their comparability over time and across countries, and their geographical and temporal coverage. Using our data set, we find that initial income inequality is not a robust determinant of future growth. By contrast, initial inequality of assets, as proxied by the distribution of land, has a significant effect on subsequent growth both in the overall sample and for developing countries separately. Only two of the 15 developing countries with a Gini coefficient for the distribution of land in excess of 70 managed to grow at more than 2.5% over the 1960–1992 period.

Theoretical explanations for this relationship proceed through two channels. One possibility is that credit rationing in the presence of indivisible investments—in schooling, for example,—may prevent the asset-poor from making economically profitable investments. Another possible channel is through the effect of an individual's asset position on her ability to participate in political bargaining or his preferences regarding the political outcome realized through a voting mechanism. The fact that inequality is not a significant determinant of future growth in democratic countries suggests that redistribution to the poor as a result of democratic voting is unlikely to be at the root of this phenomenon. While the relationship between initial asset inequality and future growth disappears in high-income economies, initial land inequality has a significant effect on aggregate

schooling attainment in the population. Investment, in turn, while not affected by initial land inequality, is significantly increased by higher levels of education.

While this suggests that initial conditions do affect subsequent growth, the impact of growth, at whatever rate, on the poor will also depend on how its benefits are distributed throughout the population. Accordingly, we investigate in Section 4 whether there is any systematic, *contemporaneous* relationship between inequality and growth. The vast empirical literature on the Kuznets hypothesis—that inequality increases with income in the early stages of development and only decreases in the later stages—was in many cases motivated by the fear that, at low levels of per capita income, the poor might lose out from development. The work on this issue is, however, flawed not only because it relies on questionable data but, more importantly, because this intrinsically intertemporal relationship has usually been tested using cross-country data. Our longitudinal data indicate that per capita income fails to be significantly associated with changes of inequality in the vast majority of countries. Many countries that started with low levels of per capita income grew rapidly without experiencing an increase in inequality, while countries that failed to grow were not immune against possibly considerable swings in aggregate measures of inequality. In the few countries where a significant relationship emerges, it contradicts the Kuznets hypothesis almost as often as confirming it. We interpret this as an indication that, rather than being governed by an unmoveable universal law, the evolution of income and inequality is affected by initial conditions and possibly policies.

In Section 5, we bring the two strands of the paper (and the literature) together and, acknowledging the simultaneous evolution of growth and changes in inequality, both of which can be affected by the initial distribution of assets, examine how both *initial* inequality and *contemporaneous* changes in inequality influence the evolution of poverty. To do this, we use our data on income shares to examine directly the factors influencing the growth in income of selected population subgroups, including the bottom 20%.

The general conclusion emerging from this analysis is that the poor, defined as the bottom 20% of the population, who do most clearly suffer from the growth-reducing effects of inequality also benefit from measures that promote aggregate growth, at least in the medium term. Three aspects are important. First, initial inequality hurts mainly the poor, but not the rich, a finding that is in line with explanations advanced in the theoretical literature which have emphasized credit rationing and the inability of the poor to undertake productive investments as mechanisms through which the effects of initial inequality and growth may be transmitted. Low initial inequality is thus doubly beneficial—it is associated with higher aggregate growth the benefits of which accrue disproportionately to the poor. Second, investment is associated with higher levels of growth for all quintile groups, but appears to benefit the poor more than the rich. Third, other policy variables appear to affect the growth of individual quintile groups' income mainly through their effect on investment. This would imply that—if initial endowments

are reasonably well-distributed—there is little justification for concern regarding a negative impact of investment-oriented policies on the poor.

These results allow us to comment on the relative merits of redistributing existing assets (land) vs. creating new assets (investment) as a means of reducing poverty. Comparing the predicted effects from both measures, one finds that a one standard deviation change in initial land inequality (16.3 points; a fairly major change by historical standards) would affect the growth rate of income for the bottom quintile by 1.05 percentage points a year. By comparison, an increase in the rate of investment by one standard deviation (9.4 percentage points) would lead to a predicted annual increase of 1.8 percentage points in the bottom quintile's income. This suggests that, especially considering political feasibility, creation of new assets will in many cases have a greater impact on poverty reduction and growth than redistribution of existing ones. Although redistribution of assets that is associated with increased aggregate investment is likely to lead to a considerable increase in the welfare of the poor, our results also caution that attempts at asset-redistribution or land reform that are associated with a decrease in aggregate investment may slow down overall growth and hurt both the poor and the rich.

Apart from the specific quantitative results reported in this paper, we believe that a shift away from aggregate analysis and from structural relationships towards analysis of specific groups and of specific policy measures offers promise for future research. The approach pursued here—analysis of growth by income group—allows us to treat inequality and growth as *joint* outcomes of the development process. It also allows us to begin—but not complete—the task of identifying policies have promise of generating aggregate growth and favoring the poor disproportionately, an issue that motivated many of the empirical investigations on this issue.

2. The data

Whether an aggregate relationship between growth and inequality exists, and, if so, through what channels it might materialize, are essentially empirical questions. Given that relatively little attention has traditionally been devoted to the quality of the data utilized to support such inferences, we start by briefly discussing the data utilized and comparing them to those used in the literature.

2.1. Income inequality

To provide a valid basis for inferences on issues of inequality and growth, data on income inequality should:

- be based on household surveys, rather than estimates drawn from national accounts statistics;

Table 1
Decadal medians of Gini coefficients for the income distribution, by Region 1960–1990

	1960s	1970s	1980s	1990s
Eastern Europe	22.76	21.77	24.93	28.60
South Asia	31.67	32.32	32.22	31.59
OECD and high income	32.86	33.04	32.20	33.20
East Asia and Pacific	34.57	34.40	34.42	34.80
Middle East and North Africa	41.88	43.63	40.80	39.72
Sub-Saharan Africa	49.90	48.50	39.63	42.30
Latin America	53.00	49.86	51.00	50.00

Regions are ordered by increasing inequality in the 1990s.
Source: Deininger and Squire (1996).

- have comprehensive coverage of all sources of income or uses of expenditure, rather than covering, say, wages only; and
- be representative of the population at the national level, rather than dealing with only the rural or urban population, or with taxpayers. [1]

Based on these criteria, we have assembled a data set on income inequality that contains at least one observation on the Gini index for 108 countries and information on shares received by different quintiles in the population for 103 countries (Deininger and Squire, 1996). There are 54 countries with four or more observations and 32 countries with eight or more observations. Table 1 shows the median values of the Gini coefficient by region and decade. [2]

To determine whether this effort has added any value, we compare our data with those used in the empirical literature. With few exceptions—for example, Fields and Jakubson (1995) and Ravallion (1995)—that are based on 'growth spells'—nearly all empirical investigations of the Kuznets curve, from Ahluwalia (1976) to Anand and Kanbur (1993) have been based on the data set assembled by Jain (1975) which, despite a relatively large number (405) of observations, contains only a modest number of data points (61) satisfying the above minimum standards. [3] The larger number of observations that is available for individual countries contained in our data set allows us to investigate the potential for country-specific Kuznets curves (see below), a type of analysis that has, to our knowledge, not been possible using existing data.

[1] A more detailed justification of these points, together with some examples, is provided in Deininger and Squire (1996).

[2] We use medians rather than means as they are less sensitive to the addition or deletion of individual countries—a relatively common event in our still unbalanced data set.

[3] These limitations have been recognized in part of the literature (e.g., Anand and Kanbur, 1993) and generally been addressed by utilizing a more limited set of 'high quality' data points.

Our data on income inequality also offer a considerable improvement over the ones used in the literature testing for a negative relationship between initial inequality and growth. Persson and Tabellini (1994) utilize share data from Paukert (1973), some of them 'of rather doubtful value' (Paukert, 1973, p. 125). Application of our minimum quality standards leads to a considerable reduction in the number of data points—from 55 observations included in their analysis to only 18—that satisfy the minimum criteria outlined above. Compared to their data, the 69 observations utilized in the high-quality data set of Alesina and Rodrik (1994), and to some degree the data used by Perotti (1995) are of better quality, especially insofar as developing countries are concerned. [4] The inclusion of observations from developed countries—from the compilation of Jain (1975)—is, however, not without problems. First, some of the data for developed countries refer to wage incomes, creating the impression of very high disparities in countries with low inequality of net or gross household income—for example, Sweden. Second, the relatively limited number of observations from developing countries—36 countries in Alesina and Rodrik; and 50 out of 67 data points in Perotti—gives rise to the suspicion that the statistical analysis captures structural differences between developed and developing countries more than regularities that are equally valid for both groups. No sensitivity analysis is performed to dispel this concern. And third, even if data of adequate quality are available, the desire to utilize *initial* income inequality often leads to the use of non-representative surveys that had been conducted in early years. For 16 countries included in the Persson and Tabellini sample, the first measure of income inequality of acceptable quality (in our data set) is dated more than 10 years later than the data point reported by them.

2.2. Land distribution

The literature has long recognized that it may be the distribution of assets, rather than income, that underlies a systematic effect of inequality on growth, for example, by restricting access to credit markets and thus the ability to finance productive, but indivisible investments. Nevertheless, data on the distribution of assets have rarely been used in empirical analysis. To partially remedy this shortcoming, we have assembled data on the initial distribution of operational holdings of agricultural land from the decennial FAO World Census of Agricul-

[4] For developing countries, Alesina and Rodrik rely on Fields (1989) who uses quality standards that are very similar to ours—the only difference being his inclusion of distributional data that refer to the wage-earning population only. Perotti uses Lecaillon et al. (1984), a slightly improved version of the data assembled by Paukert (1973).

ture [5] and other sources for 261 observations from 103 countries. The data suggest that—as is the case with other assets—the distribution of land is more concentrated and characterized by greater cross-country variation than that of income (with mean Gini coefficients of 63 and 37, and standard deviations of 19 and 9, respectively).

Data on land holdings are attractive for a number of reasons. First, possession of land could be a major determinant of individuals' productive capacity and their ability to invest, especially in agrarian economies where land is a major asset. [6] Second, in contrast to income, the measurement of which is often associated with large errors, the distribution of land is relatively easily ascertained [7] and does not require assumptions regarding the mapping from income flows into stocks of assets. The available data, however, refer to the operational rather than the ownership distribution of land. Nevertheless, we note that these data constitute a lower bound for the latter in that the rental market generally seems to contribute to a more equal distribution of land holdings. [8] Using these data we find that, indeed, the assumption of a one-to-one mapping from the distribution of income to the distribution of assets that has been used in much of the literature receives little empirical support—the correlation between the Gini coefficients for initial distribution of land and income is relatively low (0.39). Finally, coverage is more equal both geographically and over time than for data on income distribution. In most cases, observations on land distribution are available for earlier dates than estimates on income distribution and for countries in which no nationally represen-

[5] We are extremely grateful to Gustavo Gordillo de Anda who, in addition to making accessible the Statistical Bulletins from the 1980 World Census of Agriculture, provided us with the relevant material from individual countries' statistical Yearbooks in FAO's Statistical Library for the 1990 World Census. Martien van Nieuwkoop and Ernesto Franco-Temple provided valuable inputs for this analysis. Being able to use these data enabled us to considerably expand on the data sets assembled and/or used by other authors (Koo and Quan, 1985; Persson and Tabellini, 1994; Alesina and Rodrik, 1994). All FAO data are based on official 'Agricultural Censuses', conducted at the beginning of each decade; we therefore do not have to deal with data problems of the kind encountered for income distribution data.

[6] By holding income per capita constant, one can at least partially adjust for the fact that the usefulness of the distribution of land is likely to decline with rising wealth.

[7] There are of course problems arising from the fact that aggregate measures of land distribution do not adjust for soil quality or land improvements (e.g., irrigation), rarely account accurately for land held under communal tenure arrangements, and that—especially in regions such as Sub-Saharan Africa where population density is still relatively low-land may not have scarcity value.

[8] We checked the incidence of land rental for about 10 countries (including Brazil, El Salvador, Guatemala, India, the Philippines, Thailand, the Netherlands and Belgium) where such data are available. For all developing countries for which data were available, the share of land rented was below 10% of the total, the Gini coefficient for the distribution of operated land lower than the Gini coefficient for the distribution of owned land (implying that the rental market mainly redistributed land from very large owners to medium-sized operators), and the difference between the two measures did not exceed five Gini points. While use of data on land ownership would certainly be desirable, we conclude that the data used here constitute a reasonable proxy that can be used until better data become available.

Table 2
Decadal medians of Gini coefficients for the initial land distribution, by Region 1950–1990

	1950s	1960s	1970s	1980s	1990s
Sub-Saharan Africa		48.60	56.88	46.73	49.00
East Asia and Pacific	44.84	47.32	48.86	46.94	41.12
OECD and high income	58.43	59.43	52.26	54.62	59.03
South Asia	67.81	59.56	61.96	61.44	58.35
Middle East and North Africa	78.30	64.56	71.90	67.53	
Latin America	82.00	81.19	81.33	80.47	77.42
Eastern Europe	62.03	52.41	75.13	97.97	91.95

Regions are ordered by increasing inequality in the 1980s.
Source: Own calculations as described in the text.

tative data on income inequality are available. [9] Table 2 presents median values of the land distribution for the major regions.

3. Does initial inequality reduce long-run growth?

In this section, we investigate whether the strong relationship between initial inequality and long-term growth that has recently received attention in the literature holds up to the scrutiny of better data. We then check whether additional insights can be gained by using an asset-based measure of inequality and study some of the potential channels through which initial inequality might affect future growth. Three main results emerge. First, the effect of the initial income distribution on subsequent growth is not very robust. Second, inequality in the initial distribution of land is significantly associated with lower subsequent growth, an effect that is more robust than that of initial income inequality. Third, the fact that a significant relationship between inequality and growth does not exist in democratic societies suggests that explanations other than democratic voting may be at work. Wealth constraints that prevent the poor from accessing credit markets and which may be important determinants of the poor's ability to invest in physical and human capital and to substitute for insurance, could provide an alternative explanation.

3.1. Theoretical basis for an inequality–growth relationship

Relationships between the level of income and its distribution (Kuznets, 1955) and the forces shaping the latter (Kaldor, 1956) have long been discussed in the

[9] Data on initial land distribution are available for a number of countries (Argentina, Austria, Belize Burma, Haiti, Iraq, Iran, Israel, Lebanon, Libya, Mali, Malta, Paraguay, Syria, Togo, Uruguay, Yemen and Zaire) for which no nationally representative estimates on income distribution are available.

economic literature. A negative relationship between inequality and growth could emerge if investments in human or physical capital are lumpy and have to be financed through credit. Where information is costly and imperfect, equilibrium credit rationing (Stiglitz and Weiss, 1981) will arise—that is, agents will be able to obtain credit only if they own assets that can be used as collateral. A more unequal distribution of assets would then imply that, for any given level of per capita income, [10] a greater number of people are credit-constrained. In an economy where individuals make indivisible investments—in schooling and education, for example—that have to be financed through borrowing, this would imply lower aggregate growth (Chatterjee, 1991; Tsiddon, 1992). Investment possibilities may be limited not only by individuals' stock of collateralizable assets collaterable, but also by neighborhood effects and social capital with even more pronounced effects in an intertemporal context through the possible impact on societies' ability to take advantage of exogenous technological possibilities (Galor and Zeira, 1993). Under these conditions, inequality in the initial distribution of assets could be maintained over time through intergenerational bequests (Banerjee and Newman, 1993), a hypothesis that would be in line with the stylized fact of relatively high intertemporal stability of income inequality within countries, compared to great variation across countries (Li et al., 1996).

A second way in which inequality could possibly affect future growth is through political channels. The degree of inequality could affect the median voter's desired pattern of policies or it could determine individuals' ability to access political markets and participate in costly lobbying. Empirical models that have utilized this argument generally rely on some version of the median voter theorem (Persson and Tabellini, 1994; Bertola, 1993), which in its simplest (and most widespread) version, relies on democratic determination of tax rates. As the median voter's distance from the average capital endowment in the economy increases with the aggregate inequality of wealth, he or she will be led to approve a higher tax rate. This in turn could reduce incentives for (productive) investment, resulting in lower growth. If this is correct, democratic societies with a more unequal distribution of wealth should be characterized by 'exploitation of the rich by the poor—that is, high taxes and, consequently, low investment and growth, whereas undemocratic ones with similar characteristics would not.

3.2. What do the data show?

Given the deficiencies in the data used by much of the earlier literature, it is of interest to test the robustness of the inequality–growth relationship. The traditional

[10] The degree of credit rationing would, for any given income distribution, decrease as aggregate income increases and markets become more complete.

regression that has been estimated is some variant of the following specification:

$$\text{Growth}_{it} = \text{A} + \text{B}\,\text{IGDP}_{it} + \text{C}\,\text{IGINI}_{it} + \text{D}\,\text{INV}_{it} + \text{E}\,\text{BMP}_{it} + \text{F}\,\text{EDU}_{it} + \text{error}, \quad (1)$$

where *i* denotes countries, *t* denotes time, IGDP denotes initial GDP, IGINI is a measure of initial income inequality, INV indicates investment, BMP represents the black market premium, and EDU is education as measured by either average attainment in the population or enrolment rates. [11] Often only a subset of the right hand side variables is included to avoid simultaneity. In the regression results reported below, we do not include the black market premium, a variable that one would expect to affect investment. We also run separate regressions for investment and education (reported below) to determine to what degree these variables might be affected by other right hand side variables. Based on the finding that initial inequality affects education, but not investment, we report results with investment, but not education included. However, the substantive conclusions are little affected if either investment is dropped or other variables such as different measures of education, the black market premium, or measures of financial market development are added.

Given the low intertemporal persistence of growth rates within countries (Easterly et al., 1993), it is desirable to use growth over long periods as the dependent variable. Based on the availability of data from the Summers–Heston data set, we choose the period 1960–1992. Unfortunately, as the length of period increases, the quality or availability of distributional data available at the beginning of the period decreases. Short of using only a very small sample of high-quality data, there are two ways to deal with this problem. One way—admittedly imperfect—is to utilize the average Gini coefficient of the income distribution for whatever high-quality observations are available during the whole period under concern. There is some justification for this procedure in that Gini coefficients are relatively stable within countries—for the 44 countries with more than four observations, the average coefficient of variation is about 0.03, suggesting that these coefficients vary only very little around their mean. A second and preferred way is to utilize data on the initial distribution of land because here, high-quality data are available prior to 1960.

To facilitate comparison with the existing literature, we first run the regression using averages of high-quality observations on income inequality as described

[11] Data used in subsequent estimations are taken from Summers and Heston (investment and GDP), Nehru et al. (1995) (education) and King and Levine (1993) (black market premium) throughout. None of the results reported below is significantly affected by the use of data on income distribution that have been adjusted—as suggested in Deininger and Squire (1996)—to account for the fact that some of them are based on expenditures and others on income. Use of educational data from Barro and Lee (1993) led to similar conclusions.

above. The base result (Table 3, column 1) [12] indicates that the main finding of the existing empirical research is not affected by the use of our high-quality data —initial income inequality indeed affects future growth negatively. Although the quantitative effect of initial inequality is not unimportant, it is far from sufficient to explain the large differences in growth rates observed across countries. A difference in the initial Gini coefficient of about one standard deviation (nine points) would, according to the regression results, be associated with a difference in growth rates of about 0.4 percentage points. [13] Note, however, that the coefficient on initial inequality ceases to be significant once regional dummies are introduced (Table 3, column 2). This holds for all the specifications and data sources used above and leads us to question the robustness and validity of the negative association between inequality and growth. It suggests that region-specific characteristics which may, but need not, include *income* inequality, could be at the root of the relationship observed in much of the literature.

A more satisfactory alternative is to use data on the initial distribution of land which are available for a total of 66 countries. Indeed, the coefficient on the initial land distribution is highly significant and negative. Moreover, while addition of regional dummies reduces its statistical significance, it does not result in its elimination (Table 3, columns 3 and 4). This suggests that the initial distribution of assets may capture economic characteristics that are only imperfectly reflected in standard measures of income distribution. There are a number of possible explanations for this phenomenon and further research would be needed to distinguish between them. On the one hand, higher significance of the asset compared to the income distribution is consistent with the view that collateral-related constraints limit the ability of the poor to access credit markets and thus accumulate human and physical capital or avoid depletion of assets during crises. Alternatively, it could point towards that wealth-related limitations on access to political markets as the root of a negative relationship between inequality and growth.

Given that measures of inequality for income and land are only moderately correlated—a correlation of 0.39 for the 57 countries where both are available—it

[12] Using only the 31 countries for which information on the income distribution before 1970 is available slightly increases the significance of the coefficient on income distribution the quantitative value of which remains basically unchanged.

[13] For example, if Japan in the 1970s would have had an income distribution as unequal as Brazil's (Gini of 59 instead of 34), it would have grown by about 2, rather than 3.9%. By contrast, initial income distribution is not the only—and rarely the most important—variable affecting subsequent growth. Low investment contributed more to the negative growth experienced in Zambia in the 1980s than the relatively unequal distribution of income; in this case, even a move to OECD-levels of inequality would not have been enough to ensure positive growth. We interpret this as an indication that initial inequality can well make the difference between mediocre performance and rapid growth—possibly by ensuring higher investment and greater political stability—but that low initial inequality is not a sufficient condition for high growth.

Table 3
Growth regression (1960–1992) with income and land inequality

	All countries						Developing countries[a]	
Intercept	2.614	1.346	2.949	2.379	4.738	3.389	4.246	3.906
	(2.94)	(1.40)	(4.12)	(2.39)	(4.47)	(2.17)	(2.93)	(1.51)
Investment	0.132	0.122	0.134	0.123	0.107	0.115	0.130	0.148
	(6.15)	(5.09)	(6.38)	(4.77)	(4.68)	(4.00)	(3.94)	(3.59)
Initial GDP	−0.302	−0.205	−0.288	−0.264	−0.308	−0.248	−0.301	−0.338
	(3.70)	(2.23)	(4.39)	(3.49)	(4.50)	(3.06)	(1.39)	(1.54)
Income Gini	−0.047	−0.019			−0.025	−0.019	−0.018	−0.045
	(2.80)	(0.95)			(1.34)	(0.86)	(0.60)	(1.27)
Land Gini			−0.034	−0.022	−0.037	−0.027	−0.039	−0.053
			(4.07)	(1.95)	(3.85)	(2.09)	(2.43)	(2.10)
Latin Dummy		−0.530		−0.432		0.018		2.765
		(0.85)		(0.87)		(0.03)		(1.83)
Africa Dummy		−0.214		−0.254		0.324		2.191
		(0.32)		(0.46)		(0.46)		(1.52)
Asia Dummy		1.320		0.668		0.798		1.882
		(2.32)		(1.36)		(1.46)		(1.51)
R2 adj	0.3781	0.468	0.549	0.564	0.550	0.547	0.576	0.585
No. Obs.	87	87	64	64	55	55	27	27

[a]Only developing countries with a population of more than two million have been included.
Here and in all subsequent tables, figures in brackets denote *t*-values.

is possible to introduce them together as potential determinants of subsequent growth. Doing so confirms the earlier results—initial land inequality is more significant than income inequality and stays significant if regional dummies are introduced (Table 3, columns 5 and 6). This result is robust to the introduction of other variables often included in growth regressions which may also affect investment—black market premium, education, and various other financial variables. Jointly or individually eliminating the three countries with the lowest (Korea, Norway and Japan) or the highest (Ecuador, Venezuela and Peru) levels of initial land inequality does not affect the magnitude or significance of the corresponding coefficient. Dropping countries with high positive (Korea, Taiwan and Portugal) or negative (Philippines, Nicaragua and Senegal) residuals either individually or jointly scarcely affects the significance of the coefficient on initial land distribution. Finally, use of the mean of whatever information on land distribution is available instead of the initial land distribution in 1960 increases the sample size to 96 (72 developing) countries with the coefficient on land distribution remaining significant at the 5% level (not reported). The result also holds for the subsample of developing countries. Columns 7 and 8 of Table 3 report results for developing countries that have been obtained after elimination of 'island countries' with a population of less than 2 million (with results being similar for the whole sample), a procedure that reduces the sample to 46 countries—27 of

them developing countries. Neither for all countries (not reported) nor for developing countries alone is the significance and magnitude of the coefficient on land distribution affected.

The fact that only three non-OECD countries out of a total of 15 with a land Gini coefficient above 70 managed to grow at more than 2.5% throughout the period provides an illustration of the strength of the inequality–growth relationship. One of these countries, Israel, is not normally classified as 'developing', and anyway the presence of *kibbutzim* in Israel would bias the inequality of land upwards. And Puerto Rico is an atypical developing country. Thus there seems to be only one important case, Brazil, where a developing country with a very unequal distribution of land managed to achieve a growth rate slightly above 2.5%. While high inequality is associated with slow growth, the fact that other variables are held constant in our regressions implies that, for example, inadequate policies would still slow growth, as is illustrated by a number of countries with relatively egalitarian distributions of land that grew only slowly or not at all. [14]

Thus, our data do suggest that initial inequality of the asset distribution tends to reduce long-term growth. [15] Given the relatively strong empirical relationship, we are interested in finding out more about potential channels through which such a relationship could be transmitted.

3.3. How are the effects of initial inequality transmitted?

The question concerning the mechanism through which the effect of initial inequality may be transmitted allows us to replicate and expand the stylized analyses that have been undertaken in part of the literature to facilitate inferences regarding avenues that have traditionally been discussed in the empirical or theoretical literature.

The conventional interpretation of the median voter theorem maintains that, if the underlying assumption is true, one would expect initial inequality to affect growth in democratic, but not in undemocratic countries (e.g., Perotti, 1995). [16] One can test this by splitting the sample into democratic and undemocratic

[14] The biggest outliers are Mali, Senegal, Uganda, Poland and Hungary. In addition to these countries, our sample contains the Philippines, Iran, India and Sri Lanka as countries where, despite a relatively equal distribution of land, growth was low.

[15] The robustness of this result is confirmed by looking at determinants of growth in growth episodes of 10-year duration, an approach that also enables us to test for the effect of the *initial* income-distribution, defined as the average Gini coefficient during the previous decade (averaged over all available observations for the country under consideration).

[16] Recent work (e.g., Bertola, 1993; Bourguignon and Verdier, 1996, and Acemoglu and Robinson) has considerably expanded on the early and somewhat simplified treatment of the median voter theorem, implying that even results such as the one obtained here could be consistent with this framework.

regimes [17] and performing the above regressions separately for each set. [18] We find that initial inequality affects future growth in undemocratic countries, but not in democratic ones (Table 4), [19] providing little support for democratic voting as the root of the inequality–growth link. Similar findings are reported by Clarke (1995) and Alesina and Rodrik (1994).

The alternative based on credit market imperfections would be consistent with inequality being more important in low-income countries than high-income ones, as is indeed is the case in our data. Initial land inequality has a significant, and quantitatively important, effect on future growth performance in developing countries while the variable is insignificant if only OECD or if high-income countries are considered (not reported). The conclusion is that there is tentative support for a credit-market mediated link between the initial distribution of assets and subsequent growth.

For a number of reasons, lenders will generally be more willing to accept physical capital—even if it is loan-financed—as a collateral for a loan than being ready to lend against a future stream of earnings associated with the acquisition of human capital. Therefore, effects of initial inequality that are transmitted through credit markets would be expected to have a more important effect on the stock of human capital than on the amount of physical capital available in the economy. [20] We can investigate this conjecture empirically by testing whether, after accounting for initial levels of per capita income, initial inequality is associated with lower levels of average educational attainment and/or aggregate investment (Table 5). [21]

Turning to *schooling* first, we find a relatively consistent negative effect of initial inequality on this variable both in the 'parsimonious' specification with only initial land inequality and initial GDP, as well as in a more elaborate equation where we include in addition urbanization (averaged over all available periods) and the level of infant mortality (from the World Bank's social and economic

[17] We define a country as democratic if its Gastil index of civil liberties is below 2. The basic result is robust to the choice of different cut-off points.

[18] We report the results where the Gini coefficient of the income distribution is excluded, due to the potential problems associated with this variable. Its inclusion does not change the reported results.

[19] The greater magnitude of the coefficient in undemocratic countries may well be due to the high correlation between democracy and income noted by Perotti (1995).

[20] In addition to credit market constraints, one could think that fear of the poor gaining political power could prevent a rich elite from providing access to universal schooling even though such a step may be associated with economic gains, as modeled by Bourguignon and Verdier (1996).

[21] We utilize the stock variable constructed by Nehru et al. (1995) which is the most appropriate measure given that we are interested in the long-run effects of initial land inequality on subsequent decisions by economic agents. However, using average secondary enrolment (available from the World Bank's BESD data base) yields very similar results. The coefficient remains significant at the 5% level if only a dummy for Asia and Latin America are added, but drops to 10% if an additional dummy for Africa is included.

Table 4
Growth regression for democratic and undemocratic countries separately

	Democratic countries		Undemocratic countries	
Intercept	3.365	2.356	6.153	5.358
	(2.28)	(1.41)	(4.40)	(2.11)
Investment	0.093	0.076	0.162	0.191
	(3.28)	(2.65)	(4.45)	(3.93)
Land Gini	−0.016	−0.012	−0.041	−0.050
	(1.38)	(1.05)	(2.66)	(2.08)
Gini	−0.022	0.025	−0.046	−0.055
	(0.86)	(0.75)	(1.71)	(1.76)
Initial GDP	−0.251	−0.290	−1.162	−1.073
	(3.96)	(4.28)	(2.43)	(2.05)
Latin Dummy		−1.353		1.597
		(2.21)		(1.38)
Africa Dummy				1.467
				(1.33)
Asia Dummy		−0.596		1.055
		(1.09)		(1.06)
R2 adj	0.543	0.595	0.690	0.677
No. Obs.	28	28	25	25

The number of observations for democratic and undemocratic countries does not add up to the total as the ‘civil liberty’ variable was missing in a number of cases.

data-base). The latter variable in particular has a very significant and quantitatively not unimportant negative effect on attainment of schooling in the population while civil liberty is significant in some cases (implying that higher levels of civil liberty are associated with higher levels of education). In both cases, introduction of regional dummies does not eliminate the significance of the coefficient on the

Table 5
Determinants of education and investment

	Schooling			Investment	
Intercept	5.35	9.41	Intercept	7.42	12.88
	(3.91)	(7.30)		(8.15)	(3.11)
Land Gini	−0.04	−0.03	Education	1.77	1.34
	(2.09)	(2.08)		(5.42)	(3.58)
Initial GDP	0.87	0.03	Initial GDP		0.60
	(8.80)	(1.41)			(1.32)
Infant mortality		−0.04	Black Mkt. Prem.		0.00
		(6.18)			(0.21)
Urbanization		0.00	Land Gini		−0.04
		(0.02)			(0.80)
R2 adj	0.48	0.71	R2 adj	0.52	0.48
No. Obs	53	53	No. Obs	81	52

initial Gini coefficient for land at the 10% level (not reported). Urbanization (measured as the percentage of the population living in cities above 100,000 and taken from the Bank's Socio-economic data-base) is insignificant, as are a number of other variables (openness, population density, average arable land area, inflation). The average level of illiteracy is highly significant and results in all variables except the initial land distribution becoming insignificant. For all but the parsimonious equation, similar results are obtained if we replace initial land inequality with the average level of this variable over the whole period.

Regarding *investment*, we find that overall levels of schooling attainment in the population are the single most powerful explanatory variable (schooling levels alone explain more than 50% of the observed variation in this variable), but fail to ascertain an independent impact of initial (land or income) inequality on this variable. We find that, over the 30-year horizon considered, the average black market premium does not seem to have a statistically significant impact on investment. We conclude that the main channel through which initial inequality appears to affect aggregate growth is through schooling.

If schooling, but not investment, is affected by initial inequality, the specification reported above is indeed appropriate. Instrumental variable estimates where we instrumented for schooling and investment with the variables discussed above suggest that, even after accounting for investment and education, initial land inequality has a measurable and highly significant negative effect on aggregate growth, while the effect of education (which becomes insignificant in the instrumented equation) appears to operate mainly through increasing investment. Both types of estimate are robust to the inclusion of other financial variables that have been used in the growth literature. The fact that the average area per person active in agriculture is insignificant and does not affect the coefficient on initial land inequality suggests that our land inequality variable does not suffer from systematic bias, for example through the fact that more land abundant countries might have a consistently more unequal distribution of land.

These results can be summarized in three points. First, initial inequality in the distribution of land (but less so of income) appears to be associated with lower subsequent growth—a one-standard deviation decrease in inequality decreases the average annual growth by about half a percentage point. Second, there is no support for a redistributive median-voter-based explanation of initial inequality's effect on growth; on the contrary, we find stronger support for an explanation based on imperfections in financial markets for credit and insurance. Third, such imperfections appear to be more relevant for investment in human capital rather than physical capital.

4. Is there a contemporaneous inequality–growth relationship?

We now turn away from the effect of *initial* inequality on growth and look at the *contemporaneous* relationship between levels of income and inequality. It has

often been asserted that changes in aggregate growth do not automatically translate into increased well-being for the poorer groups in society. Concern about harmful effects of growth on inequality and—by implication—the incomes of the poor, was one of the main motivations for the large number of studies that tested for the presence or absence of a Kuznets curve. In this section, we examine the validity of this hypothesis using longitudinal data.

4.1. The Kuznets hypothesis

The well-known Kuznets hypothesis postulates an 'inverted U' relationship between income and inequality according to which the degree of inequality would first increase and then decrease with economic growth. Before discussing the empirical support for this hypothesis, we briefly highlight existing treatments of the issue, both in the theoretical and the empirical literature.

The first theoretical justification for a systematic relationship between inequality and income was the migration-based model provided by Kuznets (1955) to explain the inverse U-shaped relationship between inequality and income he observed in historical data. Assuming a secular shift of the population from the agricultural sector, characterized by low inequality and a low mean income, towards the industrial sector, where mean income, but also inequality are higher, it can be shown that in the early stages of development, movements from agriculture to the high-wage sector would increase income and inequality, while at later stages, they would still increase aggregate income, but decrease inequality. This model of inter-sectoral migration in a two-sector economy has been further elaborated by Robinson (1976) and Anand and Kanbur (1993). A shortcoming of these models is that the levels of productivity in both sectors—and thus implicitly growth—are exogenous. There is no human capital accumulation that would affect wages, and hence a behavioral basis for individuals' migration decisions.

While a large number of theoretical models can generate a Kuznets curve, empirical confirmation of the relationship has been limited by the availability of longitudinal data on income distribution. Kuznets' original hypothesis relied on historical data for the first half of the nineteenth century from only three developed countries, the US, England and Germany, and he carefully concluded that the data appeared to ''justify a tentative impression of constancy in the relative distribution of income before taxes, followed by some narrowing of relative income inequality after the first world war—or earlier'' (Kuznets, 1955, p.5).

Although data from a number of developed countries do support a Kuznets-type relationship, it is easy to find countries that do not fit (Polak and Williamson, 1990; Kaelble and Thomas, 1991). Even where a Kuznets curve can be observed, such as in Great Britain and the US, intersectoral migration—the chief rationale according to the 'classical' view—accounts for only a minuscule part of changes in overall inequality. Rather, inter-occupational inequality was found to be the

driving force behind aggregate inequality trends, with the increase in the remuneration to skills created by exogenous technological shocks leading to increasing overall inequality and reductions of inequality emerging as ''skill accumulation finally began to catch up with the skill scarcities created during the industrialization surge in the first half of the 19th century'' (Williamson, 1991, p.74).

The evidence for the Kuznets curve from developing countries is even more ambiguous. Looking at long-term trends, it has been concluded that the 'Kuznets relationship' is all, but absent in present-day Asian countries (Oshima, 1994). This is attributed to large individisibilities in late 19th century technology (steam engine) which prevented all, but the richest part of the population from accumulating capital, thus facilitating industrialization only at the cost of growing inequality over time. By contrast, it is argued, almost perfect divisibility of current technology, together with greater international capital mobility allows a much broader part of the population to invest in the industrial sector, thus eliminating the historical link between growth and inequality.

Although Kuznets' hypothesis deals with intertemporal relationships, it has—in the absence of adequate longitudinal data on inequality, especially in developing countries—been tempting to draw general conclusions on inequality and development from cross-sectional data. Indeed, a large number of studies—Ahluwalia (1976), Papanek and Kyn (1986), Campano and Salvatore (1988), Bourguignon and Morrison (1990), Bourguignon (1994), Milanovic (1995) and Jha (1996)—have derived empirical 'support' for the Kuznets hypothesis using cross-country evidence.

The recent literature has been more cautious, noting the ease with which addition of other variables such as education or protection tends to eliminate the statistical significance of the income variables (Bourguignon and Morrison, 1990), and the fact that for many growth spells that have been recently observed in developing countries—including ones with very low per capita income—inequality does not appear to have increased (Fields, 1989). Anand and Kanbur (1993) find that the predictions of the Kuznets hypothesis regarding the existence and location of a 'turning point' are often rejected by the data. And, although still exhibiting relatively low levels of inequality, inequality in several industrialized countries—especially the UK and the USA—has increased recently, another piece of evidence that would argue against simplistic acceptance of a 'Kuznets-type' relationship. All this would make it desirable to test the relationship using longitudinal data.

4.2. Using panel data

The problem associated with most past attempts to identify the Kuznets curve is that an intertemporal relationship is being estimated by means of cross-country data. Observations drawn from countries at different income levels are being used to approximate the evolution of income in a single country. A more satisfactory

approach—the implementation of which depends on the availability of multiple observations per country—would be to allow for country-specific intercepts and coefficients for the income variables according to the following equation: [22]

$$\text{GINI}_{it} = A_i + B_i(Y_{it}) + C_i(1/Y_{it}) + D\text{S} + \text{error}, \quad (2)$$

where i denotes countries, t denotes time, Y is real per capita income, S is a dummy for socialist countries, and A through D are coefficients to be estimated. This allows us to test formally for three different possibilities, namely:

(i) whether, as implicitly assumed in the above analyses, a 'Kuznets curve' holds with equality of coefficients across all countries (i.e., $A_i = A$; $B_i = B$; and $C_i = C$ for all i; $B < 0$ and $C < 0$);

(ii) whether countries differ from each other by some structural parameter (captured in a country-specific intercept) but, once this has been allowed for, exhibit a 'universal Kuznets curve' that is equal for all countries (i.e., $B_i = B$; and $C_i = C$ for all i where $B < 0$ and $C < 0$); or

(iii) whether the parameters of the 'Kuznets curve' are country-specific (i.e., $B_i < 0$ and $C_i < 0$ for some i), with the possibility that for some countries, such a curve is either not significant or differs from the 'inverted U' predicted by Kuznets.

To estimate Eq. (2), we restrict ourselves to the countries where four or more observations are available. There are three main results.

First, we do indeed obtain evidence of an aggregate Kuznets relationship in the cross-section when decadal country-averages are considered (Table 6, column 1). [23] However, the result is sensitive to the addition of regional dummies as well as other robustness tests. [24] For example, addition of a dummy for Latin American observations makes the 'Kuznets curve' vanish (not reported), suggesting that the

[22] We utilize the test equation for Kuznets' hypothesis with the Gini coefficient that was developed by Anand and Kanbur (1993). Although formally this specification requires that rural and urban income distributions do not overlap, it is thought to be more appropriate than the 'traditional' formulation used in much of the literature where income and income squared are the regressors. Results for the latter formulation do not differ significantly from the ones reported here.

[23] Use of country averages provides a more accurate replication of the results obtained in the literature which was mostly constrained to one or two observations per country and avoids biasing the result by giving implicitly greater weight to countries with many observations. Use of all observations results in similar conclusions. Our data also confirm the presence of a cross-sectional Kuznets curve for decadal averages from the 70s, and 80s, but not for the 60s and the 90s.

[24] As noted in Deininger and Squire (1996), the fact that our data include Gini coefficients that are defined on either (gross or net) income or expenditure, and measured by household or individual, may introduce a systematic bias into empirical estimates. Since adjustment for this bias is difficult, they suggest, as a standard robustness test, to conduct empirical analysis using not only the 'raw' data, but also inequality indicators that are either consistently defined or adjusted to account for the mean difference observed between differently defined Gini coefficients in identical surveys. Performing this type of robustness test, we note that the Kuznets curve disappears if 'adjusted' data are used (not reported).

Table 6
'Kuznets curve' in a cross-section of countries

Levels	Hypothesis I		Hypothesis II
	Differences		
Intercept	48.6088	0.1519	country
	(30.66)	(0.57)	specific
Socialist dummy	−15.4794		
	(8.22)		
Mean income	−0.001078	−0.000496	0.00006168
	(6.75)	(1.56)	(0.90)
Inv. mean income	−4112.00	−3112.97	863.00
	(2.53)	(1.12)	(0.80)
R2 adj	0.3227	0.0061	0.9294
No. Obs.	223	162	511
Hyp. turning point	1953.07	2505.23	3740.54

cross-sectional result may be affected by middle-income countries from Latin America that are characterized by relatively high inequality. All of these results are parallel to the ones obtained using the 'traditional' specification of the test equation for Kuznets hypothesis, and are thus directly comparable to results obtained by others (e.g., Bourguignon and Morrison, 1990; Fishlow, 1995). We conclude that any differences between the results obtained below and those reported in the literature are due to the different approach (time series instead of cross-section) rather than the data used in the analysis.

The time series character of our data allows us to estimate the model in decadal differences rather than levels and thus to eliminate possible country-specific effects. We find that this provides no support for the presence of a cross-country Kuznets curve for the data considered here (Table 6, column 2), a result also obtained by Ravallion (1995). Allowing for country-specific intercept dummies, [25] the coefficients on income and its inverse lose significance and even reverse sign (Table 6, column 3), thus suggesting the presence of a real rather than an inverted U. The hypothesis of equal intercepts across countries is rejected, thus not supporting the hypothesis of a universal cross-country Kuznets curve. This questions not only the specific interpretation of the evidence—whether or not a Kuznets curve exists—but suggests that using cross-sectional evidence to make inferences about intertemporal variation in individual countries is invalid, at least for the countries and time periods considered here.

[25] This implies that differences in definition across countries are absorbed in the country-specific intercept terms and robustness tests to adjust for expenditure- or income-based Gini coefficients are no longer required.

Failure to confirm a cross-country 'Kuznets curve' does not necessarily imply that such a relationship may not exist for individual countries. One way of testing for this is to use the specification with country-specific parameters as indicated above. We find that for the majority of countries (40 out of 49 or more than 80% of our sample), there is, at the 5% level of significance, no statistically significant U- or inverted U-shaped relationship between the level of income and inequality. For four out of the remaining nine cases, the data suggest the presence of a U-shaped relationship, rather than the inverted U predicted by Kuznets' hypothesis (Table 7). This leaves us with five countries—about 10% of the sample—where our data support the presence of an 'inverted U'.[26] While the number of observations may be insufficient to estimate a country-specific regression in some cases, this does not bias our results; in fact, focusing only on the 16 countries with 10 or more observations, we find that the relationship is confirmed for only one of them. There is no relationship in 12 cases, and in three cases, the data support a relationship contrary to the one predicted by Kuznets.

For some of the countries included, the time period covered—or the extent of income growth achieved—is certainly too short to make any meaningful conclusion about a 'Kuznets curve' which after all refers to a secular phenomenon. Still, for 40 of the countries included, the time period spanned by the available distributional variables is more than 15 years, and for 20 countries, we observe an increase in per capita income of more than US$3000 during the period. Two avenues are open to investigate the possibility of the Kuznets hypothesis in a more general way. First, one can test for country-specific linear trends separately for countries with high and low initial income (with a cut-off of US$3000). In this case, one would expect B_i to be positive for low-income countries and negative for high-income countries (resembling the ascending and descending parts of the 'Kuznets curve', respectively). However, the hypothesis is not supported by the data which confirm a linearly increasing trend for only two (China and Thailand) and a negative linear trend for two others (Egypt and India) out of 31 low-income countries. Inequality decreases in only three and increases in five of the 17 high-income countries included.

A second and weaker test for the validity of the Kuznets hypothesis as a long-run relationship would be to examine the possibility of a simple linear trend in a cross-section of high- and low-income countries. Results are not supportive of

[26] Closer inspection reveals that Hungary (eight observations) is a transitional economy where a rapid recent increase in measured inequality, coupled with a sharp decline in per capita income, creates the illusion of an 'inverted U' without conforming to the temporal sequence postulated by Kuznets. In Brazil (16 observations), the Kuznets curve can be eliminated by deleting one observation (for 1960), a fact that reduces our confidence in the result and illustrates the importance of careful data definition. This leaves Mexico (eight observations), the Philippines (six observations) and Trinidad (four observations) as the only countries where a true 'Kuznets-type' inverted U relation is, at the 5% level, supported by the data.

Table 7
Results from estimation of the Kuznets curve with country specific dummies

	Hypothesis III							
	Coefficient on income	*t*-value	Coefficient on 1/income	*t*-value	Predicted turning point	No. obs.	GDP difference	Years difference
Countries with a significant 'Kuznets curve' (inverted U-shaped relationship)								
Brazil	−5.76E−03	(1.96)	−5.59E+04	(2.29)	3117	15	2533	29
Hungary	−1.97E−02	(2.01)	−4.23E+05	(2.06)	4628	8	1925	29
Mexico	−5.19E−03	(2.13)	−6.39E+04	(2.01)	3511	8	3368	39
Philippines	−6.60E−02	(2.30)	−1.10E+05	(2.18)	1292	6	629	31
Trinidad	−4.14E−03	(2.83)	−1.97E+05	(2.37)	6905	4	6798	23
Countries with a significant U-shaped relationship contrary to Kuznets' prediction								
Costa Rica	2.59E−02	(2.45)	2.07E+05	(2.60)	2822	7	1534	28
India	1.75E−02	(1.94)	1.83E+04	(2.49)	1022	31	674	41
United States	1.79E−03	(3.10)	2.39E+05	(2.63)	11558	45	9323	44
United Kingdom	5.01E−03	(5.03)	3.79E+05	(4.08)	8696	31	6270	30
Countries with no statistically significant association between inequality and income								
Australia	−4.93E−03	(1.39)	−9.86E+05	(1.83)	14143	9	4767	22
Belgium	−4.34E−03	(0.31)	−7.20E+05	(0.35)	12882	4	2805	13
Bangladesh	1.99E−02	(0.58)	1.54E+04	(0.49)	879	8	460	22
Bulgaria	−1.12E−03	(0.27)	8.97E+02	(0.01)		27	2942	31
Canada	−5.69E−04	(1.23)	−5.06E+04	(0.89)	9432	23	11013	40
Chile	1.00E−03	(0.11)	3.69E+04	(0.26)	6070	5	1690	24
China	7.14E−02	(2.43)	7.40E+04	(1.83)	1018	12	530	12
Cote d'Ivoire	−7.14E−01	(0.72)	−1.62E+06	(0.74)	1506	4	139	3
Colombia	8.15E−03	(0.52)	6.00E+04	(0.53)	2713	7	1157	21
Czechoslovakia	−6.56E−04	(0.17)	2.35E+03	(0.08)		10	2310	30
Germany	−3.17E−03	(0.38)	−4.21E+05	(0.45)	11532	6	3328	15
Denmark	−5.28E−03	(0.26)	−1.01E+06	(0.33)	13813	4	3170	16
Egypt	−1.40E−02	(1.19)	−2.45E+03	(0.19)	418	4	1133	32

Spain	−3.29E−03	(1.35)	−1.42E+05	(1.42)	6582	8	4658	24
Finland	−1.62E−03	(0.28)	−2.19E+05	(0.32)	11640	10	3343	14
France	−3.25E−03	(2.53)	−9.82E+04	(1.20)	5499	7	6951	28
Hong Kong	8.35E−04	(1.31)	3.19E+04	(0.68)	6176	7	10757	20
Indonesia	−5.66E−03	(0.50)	−2.13E+03	(0.10)	614	7	996	14
Iran	−1.96E−02	(0.65)	−4.21E+05	(0.68)	4635	5	1153	15
Italy	3.98E−03	(1.24)	5.71E+05	(1.71)	11975	15	4320	17
Jamaica	2.12E−02	(1.59)	1.17E+05	(1.88)	2347	8	1230	35
Japan	5.62E−04	(1.53)	3.38E+04	(1.85)	7756	23	10777	28
Korea	−1.45E−03	(1.43)	−1.22E+04	(2.15)	2898	11	4549	23
Sri Lanka	9.29E−03	(0.85)	2.34E+04	(0.91)	1588	9	991	37
Malaysia	−6.99E−03	(1.68)	−6.00E+04	(1.38)	2930	6	2520	19
Netherlands	1.57E−02	(1.74)	1.88E+06	(1.53)	10949	12	2941	16
Norway	−7.53E−04	(0.87)	−6.84E+03	(0.08)	3014	9	8969	29
New Zealand	3.50E−02	(1.45)	3.82E+06	(1.32)	10437	12	1717	17
Taiwan	8.15E−04	(1.67)	1.41E+04	(2.21)	4165	26	6521	29
Pakistan	−2.73E−02	(0.56)	−3.88E+04	(0.60)	1192	9	448	22
Panama	3.31E−03	(0.08)	1.47E+05	(0.41)	6667	4	808	19
Poland	2.57E−02	(1.02)	5.09E+05	(1.11)	4453	9	1076	16
Puerto Rico	1.01E−02	(1.07)	3.84E+05	(1.18)	6177	4	3019	17
Singapore	1.53E−04	(0.11)	5.94E+03	(0.08)	6233	6	6302	16
Soviet Union	9.07E−03	(0.34)	3.52E+05	(0.28)	6225	5	1622	13
Sweden	2.47E−03	(0.21)	4.38E+05	(0.21)	13324	13	2741	16
Thailand	4.40E−03	(2.38)	4.17E+03	(0.55)	974	8	2947	30
Tunisia	−6.92E−03	(1.00)	−2.34E+04	(0.97)	1841	5	1674	25
Venezuela	−1.68E−02	(1.74)	−7.60E+05	(1.61)	6732	9	2350	19
Yugoslavia	−3.49E−02	(1.90)	−4.11E+05	(1.97)	3429	4	3069	26

DW = 1.875

Adj. R2 = 0.9481

No information on income for the Bahamas was available.
Differences refer to the difference between the first and last observation.

the hypothesis either. For a wide variety of cut-off points (from a per capita GDP of US$1000 to US$10,000), the coefficient on income for low-income countries is significant at 5% and positive only in two cases (in the sample below a per capita income of US$3000 and US$4000, respectively), but even then, it disappears if a dummy for Latin America is added. Similarly, for high-income countries, the coefficient on income is generally negative and significant, but evaporates as soon as a Latin American dummy is introduced. Together, these results offer virtually no support for an increase of inequality at low levels of income and a decrease at higher income levels as suggested by Kuznets' inverted-U relationship. This leads us to conclude that, based on the available evidence, the Kuznets hypothesis is either too flat to be noticeable in the data (and thus unlikely to be of relevance for policy-makers) or is not relevant for developing countries.

5. Determinants of quintile groups' income growth

The main result from Section 3 is that inequality in the initial distribution of land is associated with lower long-run growth. Section 4 has rejected the hypothesis of a systematic contemporaneous link between inequality and income levels. Both results have implications for the poor, but neither on its own nor both together facilitate predictions regarding the evolution of the income distribution. Looking directly at the growth rate in income for different income groups including the poor allows us to examine initial inequality and the evolution of inequality over time in a single empirical framework. In particular, we look at factors—including initial inequality—that influence the growth in income for the bottom quintile, the bottom 40%, the 'middle class'—the third and fourth quintile —and the top quintile.

To make inferences along these lines, we derive each quintile's income by multiplying the income share by average per capita GDP. Ideally, we would like to use 30-year growth periods as we did in Section 3. Unfortunately, the availability of share data limits the length of time period that can be investigated. Accordingly, we analyze the effect of initial inequality on subsequent growth for growth episodes where the available distributional data span at least 10 years. The fact that the main conclusions are similar to the ones emerging from the analysis of 30-year growth spells (Table 8) and relatively robust [27] suggests that it is legitimate to use 10-year growth episodes.

[27] Availability of data on the initial land distribution limits the sample to 57 observations (41 from developing countries). The result does not change significantly if averages of all available values of the land Gini, rather than the initial distribution are used, a procedure that increases the sample to 88 observations (57 developing ones). In both cases, the coefficient on the initial land distribution remains statistically significant at the 5% level for developing countries (and 10% for developed countries) if regional dummies are introduced.

Table 8
Quintile-specific growth regressions (10-year episodes)

Dependent variable	Overall GDP growth		Growth of income received by							
			Bottom 20%		Bottom 40%		Middle class		Top 20%	
Intercept	4.349	3.312	7.184	8.891	6.005	7.113	4.215	5.786	3.275	−2.829
	(3.41)	(1.83)	(2.45)	(2.12)	(2.72)	(2.25)	(2.05)	(1.97)	(1.74)	(1.11)
Investment	0.116	0.114	0.224	0.218	0.165	0.159	0.141	0.133	0.089	0.122
	(4.44)	(4.02)	(3.44)	(3.15)	(3.36)	(3.06)	(3.09)	(2.76)	(2.12)	(2.89)
Land Gini	−0.036	−0.028	−0.053	−0.062	−0.047	−0.052	−0.035	−0.043	−0.027	0.011
	(3.07)	(1.92)	(1.99)	(1.92)	(2.36)	(2.16)	(1.88)	(1.91)	(1.59)	(0.55)
Income Gini	−0.017	0.004	−0.095	−0.145	−0.048	−0.084	−0.010	−0.048	0.007	0.065
	(0.73)	(0.12)	(1.82)	(2.10)	(1.23)	(1.63)	(0.27)	(0.99)	(0.20)	(1.55)
Init. GDP	−0.198	−0.202	−0.438	−0.392	−0.374	−0.335	−0.311	−0.285	−0.170	−0.077
	(4.19)	(3.39)	(4.31)	(2.88)	(4.87)	(3.28)	(4.36)	(3.00)	(2.61)	(0.93)
Latin Dummy		−0.766		2.219		1.623		1.635		−1.876
		(0.89)		(1.11)		(1.08)		(1.17)		(1.54)
Africa Dummy		−0.750		−0.663		−0.998		−0.881		0.512
		(0.82)		(0.24)		(0.48)		(0.45)		(0.30)
Asia Dummy		0.112		0.733		0.665		0.406		1.963
		(0.17)		(0.48)		(0.58)		(0.38)		(2.10)
R2 adj	0.225	0.217	0.251	0.240	0.275	0.268	0.214	0.207	0.091	0.169
No. Obs.	133	133	98	98	98	98	98	98	98	98

Results from conducting an analysis similar to the one performed in Section 3 for each quintile using 10-year growth episodes (Table 8) facilitate three conclusions. First, we find that initial land inequality is important for the poor—where the coefficient is consistently significant—but not for the rich whose income growth is not significantly affected by this variable. This, together with the decline in the magnitude of the coefficient as one proceeds from the bottom towards the top of the income distribution, would be consistent with a collateral-based explanation according to which highly unequal distribution of assets excludes only credit-constrained individuals from making profitable indivisible investments.

Second, we find that investment is significant for all individual quintile groups. The size of the coefficient decreases as one progresses from the bottom towards the top, suggesting that, in relative terms and accounting for initial conditions, the poor are likely to benefit disproportionately from aggregate investment. This does not support the hypothesis that policies that increase investment and foster economic growth would, at least in the medium term, hurt the poor. [28]

Third, we are unable to ascertain a robust and consistent effect of either schooling or other variables (e.g., the black market premium, other measures of financial development, or initial conditions) on income growth for specific quintiles in the population. We conclude from this that, over and above their effect on aggregate investment, major policy variables do not have an independent effect on the poor. Growth-enhancing policies are therefore, at least in the medium term, not inconsistent with the goal of poverty alleviation.

To illustrate these results, Table 9 shows the impact of a one-standard deviation change in aggregate investment and land inequality on quintile-specific growth rates. Results suggest that the former (about 9.4 percentage points) would increase average annual growth by about 1.5 percentage points, two and a half times the predicted effect of a one-standard deviation change in the land Gini coefficient. Across quintiles we observe that, as noted earlier, investment and initial land inequality are, both in absolute and in relative terms, particularly important for the bottom 20%, especially in view of the fact that land inequality does not have a significant effect on income growth for the top quintile. The joint effect of an increase in investment and a reduction in land inequality amounts to an increase of almost three percentage points for the incomes of the poorest quintile, compared to just above one percentage point for the top quintile.

A policy conclusion that emerges directly from this discussion is that accumulation of new assets is likely to be a more effective way of reducing poverty than efforts to redistribute existing assets. The two options are not of course mutually exclusive. While a redistribution of assets that is associated with increased investment can increase overall growth and provide significant benefits to the

[28] As we are working with 10-year averages, we cannot make any claims regarding short term effects of these policies.

Table 9
Effect of a one-standard deviation change in investment and the land Gini on quintile income growth

	Change in income growth of				
	Overall	Bottom 20%	Bottom 40%	Middle class	Top 20%
Investment	1.54	1.84	1.24	1.40	1.13
Land Gini	0.46	1.05	0.92	0.63	

poor, pursuit of a redistributive strategy that comes at the expense of aggregate investment may well have negative overall effects on the poor.

While our data do not suggest any specific effect of government policy on income growth for different groups in the population, the finding that such policy variables do have a strong impact on the poor as well as the rich via their effects on aggregate investment by itself is of some interest. We believe that investigation of the determinants of income growth for various groups of income receivers may add an interesting and relevant aspect to existing cross-country studies concerned with the relationship between inequality and growth. In addition to extracting the discussion on inequality and growth from the impasse of aggregate cross-country regressions, often with quite limited policy implications, such an approach seems also to point towards a number of options that could increase welfare of the poor without being adverse to aggregate growth. Exploring these in more detail may be a worthwhile effort for future research.

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