THE SIMULTANEOUS EVOLUTION OF GROWTH AND INEQUALITY

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Abstract

Research on inequality and growth can be divided into two strands. One, deriving from Kuznets and Lewis, has tried to identify a mechanistic relationship between growth and inequality. The other has tried to find causal explanations of growth and inequality, treating them independently. In this paper, we draw from both strands to test whether growth and inequality are the joint outcomes of other variables and processes. We find that simultaneous examination of growth and inequality yields significantly different results, and has different consequences for policy, than previous independent studies. We also examine the determinants of growth among different income classes. We find that the disaggregated analysis clarifies and extends the results of the aggregate analysis, especially concerning the impact of globalization on the poor.

Keywords: inequality, growth, development, poverty, panel data

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I. Introduction

The empirical literature on growth and inequality can be divided into two broad strands. One, following the pioneering work of Kuznets and Lewis, is motivated by the pessimistic conjecture that the process of growth and development somehow requires increasing inequality, at least in the early stages. This has led to many efforts to identify a mechanical relationship between the level of development and inequality (see Anand and Kanbur 1993). The second strand has sought to identify the causal factors influencing the evolution of growth and inequality. This research has looked either at growth (for example, Barro and Sala-i-Martin 1995) or at inequality (Li, Squire, and Zou 1998), but has not tried to identify those factors which might simultaneously influence both growth and inequality.[1]

Thus, the literature that looks *simultaneously* at growth and inequality, descending most notably from Kuznets (1955), relates them in a mechanistic manner that ignores or minimizes the role of other causal factors including policy, while the literature that incorporates other causal factors including policies looks *independently* at growth and inequality. Neither is particularly convincing from a theoretical standpoint: the evolution of growth and inequality must surely be the outcome of similar processes. Nor is either particularly useful for the policy maker, who needs to balance the impact of policies on both growth and distribution. In this paper, we investigate whether growth and inequality are simultaneously determined by a set of common factors. While our primary focus is to demonstrate this methodological point, we hope that the

[1] The growth literature sometimes includes inequality as an independent variable -- see Persson and Tabellini (1994), Clarke (1996), and Forbes (1999) – but still focuses on the evolution of growth independently.

analysis will point to those policies that in proper combination will simultaneously increase income and reduce inequality.

According to recent studies, there is no robust, systematic correlation between the evolution of income and inequality, although both positive and negative relationships have been found in some studies, for some countries. In the most comprehensive cross-country study to date, Deininger and Squire (1998) detect no statistically significant relationship between the evolution of income and inequality in 39 out of a sample of 48 developed and developing countries. For four of the remaining countries the data suggest the presence of a U-shaped relationship, rather than the inverted U predicted by Kuznets, while five countries exhibit an inverted U.[2] Similarly, Ravallion and Chen (1997) find no correlation between growth and inequality in a sample of developing countries.

A simple comparison of India and Taiwan (China), the two developing economies for which we have the best data on inequality, provides anecdotal evidence that growth and inequality are not systematically related. Figure 1 shows the evolution of inequality and income per capita over a period of 30 years. Growth since the war has differed enormously between these two countries, while inequality has changed comparatively little. Income per capita in India roughly doubled during this period, but income in Taiwan has increased more than five-fold. In contrast, inequality in both India and Taiwan has been roughly constant since the early 1960s, with a Gini coefficient of around 30.

Thus, both the cross-country evidence and the India-Taiwan comparison suggest that the evolution of income and inequality is not bound in some immutable relationship. Or, at least, if

[2] We tried to find a Kuznets-type relationship among the 49 countries for which we have sufficient data. A generous interpretation of standard errors gives eight countries which conform to the Kuznets hypothesis, 11 which show inverse Kuznets curves, and 30 countries (more than 60 percent) in which there is no relationship.

there is some underlying relationship, other factors – institutions, policies, external events – are more important in determining the wide range of observed outcomes. Instead of trying to identify a mechanistic link between growth and inequality, a more fruitful approach asks: what factors, including policies and institutions, influence the joint evolution of inequality and income? The main focus of this paper, therefore, is to assess the importance of treating growth and inequality simultaneously in contrast to the current tendency to analyze them independently.

We are particularly interested in identifying mutually exclusive variables – variables that influence growth but not inequality or vice versa – for two reasons. First, while we know that in principle all actions will affect both growth and inequality, it would be useful to find that in practice some actions predominantly affect one or the other but not both. Governments with different objectives with respect to growth and inequality would then have the instruments to pursue preferred outcomes. Second, the empirical literature to date has largely proceeded on the assumption of mutually exclusive policies. Or, to put the same point differently, the current practice of investigating the determinants of growth and inequality independently is only justified to the extent that policies influencing growth do not affect inequality, and vice versa.

The paper is organized as follows. Section II briefly describes the data. We rely on an expanded version of the Deininger-Squire data set (Deininger and Squire 1996) for our measure of inequality, and on the Penn World Tables (Heston and Summers 1991) for growth data. In Section III we begin with independent models of growth and inequality using a limited set of explanatory variables commonly found in the literature. These base equations have little in common, but we test whether independent examination of even these limited models can be justified. If the models are not independent, the empirical results and consequences for policy arising from simultaneous consideration of the two goals may differ from previous analysis that assumed orthogonality.

The independent analysis finds that of the two variables which the basic models have in common – education and M2/GDP as an indicator of financial development – the former is probably uncorrelated with growth in the aggregate, while the latter is related positively with growth and negatively with distribution. Already, we find that the independent analysis of these two equations is not justified.

A further requirement of the separate treatment of growth and inequality is that, apart from the common variables, the other independent variables of the inequality equation have no impact on growth, and those from the growth equation have no impact on equality. These variables are mutually exclusive by assumption. We therefore expand the simple models by introducing all the (assumed) mutually exclusive variables from the inequality equation into the growth equation and vice versa. This allows us to test whether these variables are indeed mutually exclusive and whether their inclusion changes the empirical results. The results reveal that in fact the majority of these variables are jointly significant. Thus, at least for these simple models, we conclude that the independent analysis of growth and inequality produces potentially misleading, or at least incomplete, results for the policymaker.

In section IV, we shift from the analysis of aggregate growth and aggregate inequality to an analysis of growth rates disaggregated by population groups. In a way, this is a more direct means of looking at the joint evolution of growth and inequality. Equality of parameter estimates across groups implies that the variables influence aggregate growth but not aggregate equality. Similarly, parameter estimates which differ across quintiles suggest that the variables affect aggregate distribution. Thus, these regressions provide a rough robustness check on the aggregate results. More importantly, they add significantly to our understanding of the impact of policies on welfare. The Gini index, our empirical measure of aggregate inequality, is a summary statistic describing the area of a Lorenz curve, and tells us nothing about the shape of the Lorenz curve. The relative incomes of poor and rich may change without changing the aggregate Gini coefficient. Conversely, an improvement in the Gini coefficient does not necessarily imply that the poor are even relatively better off.

The results strongly indicate that policies can have different consequences for different groups within an economy. That a particular policy is correlated with growth in the aggregate says nothing about the correlation of that policy with growth among the poor or any other sub-group of the population. In addition, the results also show that the correlation of any policy with the aggregate Gini index does not necessarily imply a particular correlation of that policy with group-specific growth rates.

Section V summarizes the empirical results and draws three policy-relevant lessons. First, there is sufficient evidence of simultaneously influential variables to warrant the joint consideration of growth and inequality. Independent analysis may lead to the recommendation of policies for one outcome with unintended consequences for the other. For example, while trade liberalization is a *sine qua non* of growth-oriented policy reforms, it may have unintended negative consequences for income distribution, at least in the short run. Similarly, while trade liberalization is unambiguously correlated with aggregate growth, that is not true of growth among all classes. Second, we find evidence of variables which are uniquely correlated with each outcome. These mutually independent variables allow some room for manoeuver.[3] For example, the expanded model suggests that education influences equality independently – it appears to have no correlation with aggregate growth. Finally, the evidence indicates that growth is far more sensitive than inequality to changes in policies. This is consistent with previous evidence, that growth is far more volatile, while inequality persists.

II. The Data

Deininger and Squire (1996) have amassed the largest dataset of inequality measures. Their original file contained more than 2,600 observations, of which 682 (from 108 countries) were deemed acceptable (that is, they satisfied restrictions concerning national representation, sampling and survey methods, and so on). These data are available on the internet from the World Bank. We have now extended this acceptable set to 757 observations for 125 countries. The majority of the additions cover Europe and Central Asia, although new data points were added to every region and income class.[4] Our multivariate analysis uses a subset of these data, for two reasons. First, we exclude observations prior to 1960, to ensure some consistency within the sample, and since measurement errors were almost certainly higher in earlier periods. Second, the scarcity of independent variables describing policy and institutions reduces the dataset even further.

We also adjust the data to control for differences in measurement across observations. There is no international standard; countries use different methods, often inconsistently.[5] For example, a measure of individual welfare may be based on the level of income or expenditure. In

[3] These mutually exclusive variables also serve to identify the system.
[4] See Appendix 1 for sources and details of the additions.
[5] One of the criteria applied by Deininger and Squire to select acceptable data is that the series contain information on the methods used.

the case of income, welfare may be measured prior to or following the individual's payment of taxes and receipt of benefits. There are also differences in the method of aggregation across the population. National welfare measures may be derived from household-level welfare by assuming equal weights for each household in the aggregation. Alternatively, the household-level measures may be weighted by household size in the aggregation.

We conduct the analysis in this paper on an adjusted Gini index; that is, the recorded Gini index adjusted to the level it would be, were it calculated on an individual-weighted, expenditure basis. The adjustment is discussed in more detail in Appendix 2. The impact of this adjustment is modest. The overall correlation between the adjusted and original Gini indices is 0.98, although there are some differences within regions, notably in South Asia. Using the adjusted Gini indices, poorer countries (as classified by the World Development Report) are significantly more unequal, with a mean Gini index of 38.7, compared to 31.6 among wealthier countries. A simple comparison of means overlooks significant variation, especially among developing countries. Countries in sub-Saharan Africa and Latin America have a mean Gini index of around 48, whereas countries in East and South Asia have a mean Gini index around 35 (comparable to Australia).

For income data we use the Penn World Tables, Mark 5.6 estimate of GDP per capita in 1985 PPP dollars, RGDPCH. We again restrict the sample to the years since 1960, although this still leaves more than 4300 observations on 147 countries. Overall, GDP per capita has increased more than 80 percent since the 1960s, with the greatest changes occurring during the first half of the period. Between the 1980s and the 1990s there was almost no change in median aggregate income.

The Deininger-Squire dataset also contains some 600 observations on income shares by decile or quintile. We calculate quintile-specific annual incomes as the product of a five-year moving average quintile share and annual GDP per capita, and compute growth as the average change in this measure over the period. Since the share data reflect international and intertemporal differences in measurement method, we adjust the quintile shares in the same manner as the Gini coefficients. On average, the poorest quintile receives 7.2 percent of total income, while the wealthiest receives 41 percent. Average growth among the poorest quintile is 2.03 percent per year (compared to an overall mean of 2.19 percent per year in this sample), and 2.14 percent among the wealthiest quintile.

We estimate our models on a relatively small sub-sample of countries. We don't think this is a source of significant bias. T-tests of mean differences between those countries in the sub-sample and those left out of the sub-sample are insignificant for almost all the variables used in this paper (results available). The only exception is the indicator of civil liberties, for which the included sub-sample of countries has significantly lower (that is, more favorable) values.

III. Combining growth and equality

To see whether past practice of investigating growth and inequality independently leads to misleading empirical results and possibly serious mistakes in inference and policy recommendations, we draw on “standard” growth and inequality models available in the literature. We will examine first the implicit consequences across outcomes in these models when they are estimated and analyzed separately. Under these circumstances, do the standard models imply affinity or conflict between growth and equality? Second, we will re-estimate the equations using both sets of variables in each equation to test the implicit assumption that growth variables are orthogonal to equality and vice versa.

A. Model Derivation

The standard models of growth and income distribution take the simple form:

$$\Delta y_{it} = S'_{it}\alpha + X'_{it}\beta + u_{it} \tag{1}$$

$$Gini_{it} = S'_{it}\omega + Z'_{it}\psi + e_{it} \tag{2}$$

where *X* is a vector of "growth" variables and *Z* is a vector of "equality" variables, as defined in the literature, and *S* is a vector of variables common to both models. The standard growth model assumes that the elements of *Z* are orthogonal to growth and that the elements of *X* are orthogonal to equality. Our first set of estimations maintains those assumptions in order to replicate the standard models.

Our second set of estimations drops those assumptions, allowing the growth variables to enter the equality equation and the equality variables to enter the growth equation. In addition, we estimate the equations as a system, in which growth explicitly depends on equality and vice-versa:

$$\Delta y_{it} = M'_{it}\beta + \gamma Gini_{it} + u_{it} \tag{3}$$

$$Gini_{it} = M'_{it}\delta + \tau\Delta y_{it} + e_{it}\,, \tag{4}$$

where $M = [S, X, Z]$ is the combined matrix of all right-hand-side variables from (1) and (2). Here we also make some assumptions concerning the error structure. The errors in these two equations contain both random (iid) and persistent country-specific effects.

Following Easterly, Kremer, Pritchett and Summers (1993), we assume that there is no correlation in growth rates across periods, but we do allow for unobserved heterogeneity so that

$u_{it} = \alpha_i + \upsilon_{it}$. On the other hand, previous research and casual examination of the data indicate that inequality is persistent. We therefore allow for serial correlation in the Gini index as well as permanent unobserved characteristics, such that $e_{it} = \mu_i + \eta_{it}$ and $\eta_{it} = \rho\eta_{i,t-1} + \varepsilon_{it}$. We can solve for the final forms

(5) $$\Delta y_{it} = M'_{it}\beta^* + u^*_{it}$$

(6) $$G_{it} = M'_{it}\delta^* + e^*_{it},$$

where $\beta^* = \dfrac{\beta + \gamma\delta}{1-\gamma\tau}$, $\delta^* = \dfrac{\delta + \tau\beta}{1-\tau\gamma}$, $u^*_{it} = \left[(\gamma\mu_i + \alpha_i) + \gamma\rho\eta_{i,t-1} + (\gamma\varepsilon_{it} + \upsilon_{it})\right]/(1-\gamma\tau)$, and $e^*_{it} = \left[(\tau\alpha_i + \mu_i) + \rho\eta_{i,t-1} + (\tau\upsilon_{it} + \varepsilon_{it})\right]/(1-\tau\gamma)$. Note that (5) and (6) are *not* the reduced forms of the structural equations (3) and (4). We solve the system for the included dependent variables, but each equation contains in addition within *M* both exogenous and endogenous variables. More importantly, examination of the errors in (5) and (6) reveals that even though the serial correlation is explicit only in the structural Gini equation, the lagged Gini error contaminates both growth and inequality through the interdependence of the two equations. Eliminating the unobserved heterogeneity from (5) and the serial correlation and unobserved heterogeneity sequentially from equation (6) will at least be inefficient.[6] Our solution is presented in section III.B below.

Easterly *et al*. (1993) (EKPS) provide our standard growth model, which is in turn based on Barro (1991). From Barro, they include initial GDP and school enrollment, assassinations and revolutions, and the share of government consumption. EKPS make a few changes to account for data availability and to test their hypotheses regarding economic shocks and growth: they add

[6] This is the method suggested by Kmenta (1986).

the black market exchange rate, inflation, the share of trade in GDP, the ratio of M2 to GDP, and changes in the terms of trade. Our standard inequality model comes from Li, Squire, and Zou (1998) (LSZ), who include education, a measure of civil liberties, the ratio of M2 to GDP, and a measure of the distribution of land. Our base case attempts to approximate the results presented in EKPS and LSZ. These are among the most parsimonious explanations of the (dependent) variables of interest.

In reestimating these equations, we make a few changes to the source models. First, we drop the assassinations and coups variables, partly because they are insignificant in both Barro's and EKPS's tests, and to maintain a larger sample. Second, EKPS include two measures of trade openness: the black market exchange rate premium, and the ratio of imports and exports to GDP, both of which they find to be negatively related to growth, although the coefficient on trade share is insignificant. We use instead the Sachs and Warner (1995) index, which includes measures of exchange rate overvaluation, tariffs, and non-tariff restrictions on trade. EKPS include both primary and secondary enrollment, which we replace with one variable (mean years of schooling in the adult population). We also make two changes to the inequality model. We use the original value of the civil liberties index (Gastil 1990) rather than an inverted index, which LSZ use. Their inversion is merely to aid interpretation, since the index is constructed so that *lower* values of the index imply *higher* levels of civil liberties. Second, we use the contemporaneous value of schooling, rather than schooling in 1960, appropriately instrumented where necessary.

The base growth and equality models have only two independent variables in common – education and the ratio of M2 to GDP as a measure of financial development. The implicit assumption is that the other independent variables in the growth equation do not affect inequality, and, similarly, the other independent variables in the inequality equation do not influence growth.

In the expanded models we test this assumption explicitly. We add the variables from the standard growth model to the inequality model, and we add the standard inequality variables to the growth model.

B. Specification and estimation issues

Prior to the publication of the Deininger and Squire database, empirical studies of growth and inequality were exclusively cross-sectional. This was also true for many independent studies of growth, in which average growth over a period is regressed against the average (or initial) values of a set of right-hand-side variables (e.g. $\overline{\Delta y_i} = \alpha_i + \beta_1 \overline{X}_{1i} + \beta_2 \overline{X}_{2i} + \varepsilon_i$, where the bars indicate country averages). This method is problematic for a variety of reasons. For example, data restrictions often force the country averages for each value to be calculated on different periods. If variables are measured only once or a few times in the sample period, the average values of X_{1i} and X_{2i} may refer to entirely different sub-periods. This may not be important if, as EKPS argue, country characteristics are persistent. If, on the other hand, country characteristics or policies change and these changes have an impact on growth or equality within the sample period, they will be missed by a purely cross-sectional model. For example, terms of trade shocks may have immediate consequences for growth. In recent work, Barro has acknowledged that the within-country dimension provides additional information (see e.g., Barro 1997).

With cross-sectional data, all unobserved cross-country variation is relegated to the error term, leading to the underestimation of standard errors and possible mistakes in statistical inference. Panel-data formations make it possible to control for the unobserved cross-country effects, through fixed- or random-effects specifications or regressions in differences. However, inequality varies much more across countries than over time, and the characteristics of this variance cannot be examined by techniques which eliminate cross-country effects and focus exclusively on the within-country relationships (i.e. fixed effects and first difference estimators). Banerjee and Duflo (1999) also argue that taking out fixed effects also exacerbates measurement error. We try to minimize measurement error by aggregating over five-year periods, and we keep the analysis reasonably comprehensive by including country effects as described below.

Our estimating equations are most likely biased, not only by endogeneity arising from correlation with the unobserved country-specific effects but also from reverse causality. For example, it has been argued that growth causes education as much as education causes growth (Bils and Klenow 1998). If there were no unobserved differences across countries and no endogeneity, the model could be estimated as a pair of seemingly unrelated regressions (SURE) on pooled data. With unobserved country effects, but strict exogeneity, we could estimate the model with a standard panel-data technique. In the presence of endogeneity, we must use instrumental variables; but it is difficult to find instruments that are truly exogenous to the unobserved country-specific component of the error term. Even if valid instruments could be found, two-stage least squares would be inefficient, since the unobserved persistent heterogeneity causes serial correlation in the errors. Finally, simple demeaning or differencing will also eliminate the time-invariant variables which might be of interest, since these methods simply purge the models of all cross-country differences.

Our solution is a variant of the method suggested by Keane and Runkle (1992) (KR). For a single equation model such as

$$y_{it} = B'x_{it} + u_{it}, \tag{7}$$

in which $E(u_{it} \mid x_{it}) \neq 0$, KR assume that there exists a set of instruments Z which is at least *weakly* exogenous – that is, they are predetermined, but not necessarily uncorrelated with the unobserved country-specific effect, so that $E(u_{it} \mid z_{is}) = 0$, $\forall s \le t$, but $E(u_{it} \mid z_{is}) \ne 0$, $\forall s > t$. In that case, KR suggest an estimator that is based on a forward-filtering transformation of the data. They apply the result from Hayashi and Sims (1983), who showed that if a time-series equation has serially correlated errors and predetermined instruments, the serial correlation can be eliminated by a transformation that makes the dependent variable at time *t* a linear function of the values of the dependent variable for time periods *t* and later. This is similar to the forward filtering transformation of Arellano and Bover (1990) (AB), but according to Keane and Runkle, the AB transformation does not eliminate all serial correlation and imposes the additional restriction of equicorrelation on the Ω matrix.

The KR estimator is found by considering the equation in stacked form

$$(8) \qquad \begin{pmatrix} y_1 \\ \vdots \\ y_N \end{pmatrix} = \begin{pmatrix} X_1 \\ \vdots \\ X_N \end{pmatrix} B + \begin{pmatrix} u_1 \\ \vdots \\ u_N \end{pmatrix},$$

where y_i, X_i, and u_i represent the $[T \times 1]$, $[T \times k]$, and $[T \times 1]$ observation matrices for the *i*th group. The KR estimator is constructed by first obtaining a consistent estimate of $\Sigma_{TS} = E(u_i u_i')$ for each group. Next, compute the *upper triangular* Cholesky decomposition of Σ_{TS}^{-1}, called $\hat{P}_{TS}$ (such that $\Sigma_{TS}^{-1} = P_{TS}' P_{TS}$). Finally, premultiply equation (7) by $\hat{Q}_{TS} = I_N \otimes \hat{P}_{TS}$, and estimate the transformed equation by 2SLS, using the original (untransformed) Z_{it} as instruments. The consistent estimate of Σ_{TS} derives from a 2SLS regression on the original untransformed equation (7): $\hat{\Sigma}_{TS} = \frac{1}{N} \sum_{i=1}^{N} (\hat{u}_{TS} \hat{u}_{TS}')$. In our multiple-equation case, we find $\hat{\Sigma}_{TS}^{\Delta y} = \frac{1}{N} \sum_{i=1}^{N} (\hat{u}_{TS}^{\Delta y_i} \hat{u}_{TS}'^{\Delta y_i})$ and

$\hat{\Sigma}_{TS}^{G} = \frac{1}{N}\sum_{i=1}^{N}\left(\hat{u}_{TS}^{G_i}\hat{u}_{TS}'^{G_i}\right)$ from 3SLS on the untransformed equations (1) and (2) for the base model and (5) and (6) for the expanded model.

C. Results of the standard models

Table 1 presents our approximations of the standard models. In the original papers, each equation is of course estimated independently. In order to test cross-equation restrictions, we estimate them simultaneously. We present three sets of results: pooled OLS (SURE), instrumental variables (3SLS), and the Keane and Runkle 3SLS estimates. Note first that in column two (the 3SLS specification), the relatively weak Sargan test indicates that the instruments are not uncorrelated with the error term in the Gini equation.

Our standard growth results provide weak confirmation of the main hypothesis of EKPS, that terms of trade shocks are positively related to growth – they are insignificant in the first two specifications and only weakly significant in the third. Also, as in the previous studies, initial GDP enters negatively, which suggests convergence; and the level of financial development, proxied by the ratio of M2 to GDP, is positively related to growth. However, where EKPS find many variables insignificant, we fail to find significance only for education. All remaining variables are significant in the Keane and Runkle specification. Greater government expenditure and greater openness (the Sachs-Warner index) are related to faster growth. Inflation is positively related to growth. We suspect that this is a misleading result, induced by the non-normal distribution of inflation (which is both highly variable and skewed to the right), and a non-linear relationship between growth and inflation. Indeed, in an alternative specification, both the level of inflation and its square are significant, indicating an inverse-U shaped relationship between inflation and growth.

Our standard inequality results agree broadly with those of LSZ, at least in sign. Differences are likely due to our use of a more appropriate error correction technique. These results show that the impact of education on equality is positive. An additional year of education, on average, lowers the Gini index by 0.8 points. Our estimates of the impact of land distribution also correspond to those of LSZ. A one-point reduction in the Gini index for land distribution (from a mean of 60.7) would on average improve equality by one half of one point. A one standard deviation (about 30 percent) improvement in land distribution would lower the Gini index by about 9 points (about 25 percent, roughly one standard deviation below the mean Gini). In line with LSZ, we also find that civil liberties are positively related to equality.[7] A one point (36 percent) reduction in the Gastil index, which implies an improvement in civil liberties, would lower the Gini index by 2.3 points. In contrast to LSZ, we find that broader financial development (M2/GDP) diminishes equality. The difference is most likely due to differences in estimation method, since the SURE results indicate a negative relationship between M2/GDP and the income Gini index. Although LSZ instrument M2/GDP with its one-period lag, there may be dynamic aspects to the relationship as well as correlation with the persistent unobserved characteristic which render the lag invalid as an instrument.

The results for the SURE and 3SLS specifications of the standard model indicate that the two common variables – education and the ratio of M2 to GDP – are significant only to equality. They are neither jointly nor independently significant to growth. This suggests that they may be valid instruments to identify the Gini equation, in addition to those variables unique to the Gini equation by assumption. However, KR-3SLS results reveal that the M2/GDP ratio is positively

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[7] Note that LSZ invert the values civil liberties index: in the original (Gastil 1990), and in this paper, *higher* values of the index refer to *lower* civil liberties.

related to growth. Thus, an increase in broad money will increase growth at the cost of higher inequality. This also indicates that the two equations in the system are identified only by assumption. The two common variables are jointly significant both to growth and to inequality.

These standard equations assume that openness, terms of trade shocks, and inflation are exclusive to growth, while civil liberties and land distribution are exclusive to equality. If these assumptions hold, the policymaker wanting to improve equality has tools other than education for influencing distribution independently of growth. In general, however, the consequences of these policies cannot be determined *a priori*. In the next section, we test whether these assumptions are supportable.

D. Results for the expanded models

The preceding analysis and conclusions assume that the variables classified as mutually exclusive truly are: for example, that openness does not affect equality, and land distribution does not affect growth. If they do, the empirical results for both growth and equality could be changed, as well as the corresponding implications for policy. To test this orthogonality assumption, we expand our standard models to include the growth variables in the equality regression and the equality variables in the growth regression. In other words, we relax the constraints on the excluded variables in equations (1) and (2), and estimate (5) and (6). We also make one change to the specification in this section. Since agriculture comprises a larger part of income among developing countries, we allow the impact of land distribution to differ across groups by adding an interaction between the land distribution Gini and a dummy for developing countries. The results are presented in table 2.[8]

[8] We present the Keane and Runkle 3SLS results only. Other results are available.

The results for the growth variables in the expanded growth model are similar to the previous KR-3SLS results. The M2/GDP ratio, which was only weakly significant in the previous model, falls out once the other variables are included. The parameter estimates on the remaining growth variables are roughly the same, with the notable exception of the openness index. The marginal impact of greater openness to growth is nearly double that found in the previous specification. Inflation is again positively correlated with growth, but as before, the results are driven by hyperinflationary outliers from Latin America, and adding the square of inflation renders both parameter estimates insignificant.

Of the additional variables from the equality model, only land distribution is significantly related to growth. Not surprisingly, perhaps, land distribution is significantly correlated with growth only among developing countries. What is more surprising is that land distribution appears to be negatively correlated with growth – that is, more equitable distribution drives down growth. Our examination of quintile-specific growth rates in section IV below highlights and expands this result. The other unique variable from the inequality equation – civil liberties – does not appear to influence growth.

This extension of the standard growth model shows that the results do change when additional variables are considered. The changes are manifest not in omitted variables bias to the existing parameter estimates, since they change little across specifications, but rather in the omission of potentially important information for the policymaker concerning the relationship between growth and land distribution.

The results for the equality variables in the expanded equality equation are also similar to the previous results, with one exception – the civil liberties variable now falls from the model. The parameter estimate on education is unchanged, but the partial correlation between financial depth and inequality becomes stronger once other factors are controlled. The results also justify the distinction between poor and rich countries in the impact of land distribution. While the land Gini is universally positively correlated with the income Gini, the impact is far greater among developing countries. A one-point reduction in the land Gini would reduce the income Gini by 0.1 points in the average developed country, and 0.5 points in the average developing country.

Of the additional growth policy variables, two – government expenditure and openness – are significant and inversely related to equality. On average, government expenditure is regressive. We return to this point in section IV, but it suggests the need for detailed benefit incidence examination. The result for openness suggests that the process of globalization, while it is irreversible and essential for growth, may have some adverse consequences for distribution. Considering the recent debate over the consequences of globalization, this point is worth discussing in some detail. Although this result might provide some support to those who claim that the poor are harmed by globalization, this interpretation is difficult to maintain on the strength of these results alone. These regressions provide point estimates of short-run changes; they do not estimate the long-run steady-state consequences of policy changes. It is usually assumed that adjustment costs decline over time, while the gains from more efficient resource allocation continue *ad infinitum*.

A recent survey of empirical research found that following a short period of adjustment, greater openness leads to greater manufacturing employment (Matusz and Tarr, 1999). In principle, the relative costs and benefits are determined by the speed of the resource flows from the diminished (that is, previously import-substituting) sectors to those which are successful exporters. Although they rely for the most part on developed-country evidence, Matusz and Tarr suggest that adjustment costs will be even lower in developing countries, where a large share of labor is in agriculture and informal activities with very flexible labor markets. They present data which shows that employment in small and medium-size enterprises in developing countries grows during and following reforms, but there is no evidence that the increase in employment is actually caused by the reforms.

That said, these results do have some theoretical foundation. Within an economy, greater openness leads the sectors which had been protected from import competition to contract, reducing input demand. Both labor and capital must adjust, but capital is relatively more mobile. Since the poor receive a larger share of income from wages, the incomes of the poor recover more slowly than the incomes of the rich. There is some evidence that increased globalization is responsible for some of the increased wage gap between low- and high-skilled workers within industrial countries (see Rodrik 1998).

However, the total impact of globalization on global wage inequality is theoretically ambiguous. The Stolper-Samuelson theory of factor-price equalization suggests that liberalization increases the demand for goods from, and the returns to labor in, the sector which intensively uses the relatively abundant factor in a country. Developing countries are presumably rich in relatively unskilled labor, whose wage should rise both relative to skilled labor within the country and relative to unskilled labor in industrial countries, thus reducing global inequality. On the other hand, this same process leads wages for skilled labor to rise in developed countries and fall in developing countries, offsetting the fall in global inequality. It is likely that the latter effect outweighs the former, since labor market policies in developed countries are stronger than those in developing countries, so the fall in industrial-country unskilled wages following liberalization is relatively smaller than the fall in developing-country skilled wages. In addition, it has been suggested that the Stolper-Samuelson factor-price equalization theorem holds only within groups of relatively homogeneous countries, and not for the world as a whole (Davis 1999).

The welfare consequences of liberalization for rural households are also theoretically ambiguous. If liberalization is associated with the elimination of domestic food price controls and an increase in staple food prices, net producing households (those with a marketed surplus) will see an increase in income, but net purchasers of food will suffer, as will the urban poor. On the other hand, liberalization may increase farmgate prices for export crops as well as other exported goods relative to staple food prices, improving the welfare of subsistence farmers. Again, the predicted net effect for all countries is ambiguous, but according to our results, the negative effects outweigh the positive.

The results of the standard models (Table 1) suggest that education is exclusively significant to equality, and that money supply has contradictory effects on growth and equality. Assume for the moment that these are comprehensive sets of variables, including everything that is significant to either outcome. This still leaves policymakers with a wide range of options to pursue diverse development goals. For example, these simple models suggest that increased openness would yield distributionally neutral growth, as would expanded government expenditure. Conversely, more equitable land distribution would yield better income distribution, while not affecting growth.

Table 2 shows that the determinants of growth and distribution are not mutually exclusive. The policy variables which are assumed to be exclusive – government expenditure, inflation, and openness for growth, and land distribution and civil liberties for equality – are not, in general, exclusive. Three of these variables are both individually and jointly significant, and all three involve trade-offs between growth and equality. Government expenditure, openness, and the developing-country Gini for land distribution are all correlated positively with growth and negatively with equality. Thus, a policymaker who uses the results from the independent examination of growth to drive policy runs the risk of unwittingly adversely affecting distribution. The expanded results reveal that increasing government expenditure or openness for growth would lead to greater inequity.

Four of the policy variables examined do prove to be mutually exclusive – education, M2/GDP, and land distribution (in developed countries) for equality, and inflation for growth, although the non-linearity of inflation renders it ineffective as well as implausible as a policy instrument. The policymaker wishing to improve equality independently of any changes in growth therefore has a few tools at his disposal. Perhaps not surprisingly, there are no mutually beneficial variables. These results suggest that there is no single policy which advances both greater growth and greater equity simultaneously. On the other hand, it is possible to derive from these results a set of policies that will in combination achieve both goals – for example, opening borders to trade while at the same time expanding education.

These aggregate results support the second hypothesis of EKPS – that growth is highly variable and highly responsive to changes in policy. More importantly, we can also confirm a corollary hypothesis – that income distribution is both persistent and relatively unresponsive. While the uncontrolled coefficient of variation of growth (that is, across both countries and time) for our sample is 1.8, the coefficient of variation of the income Gini coefficient is 0.3. This is borne out in the correlation of growth and distribution with policy. As Table 3 shows, in no significant case is the elasticity of growth with respect to a policy smaller than the elasticity of the Gini coefficient.[9] In general, the elasticities for growth are at least double, and sometimes an order of magnitude larger than for inequality. The only significant exception is for developing-country land distribution, where the relative elasticity is a mere 50 percent larger for growth than inequality.

IV. Exploring quintile-specific growth rates

The preceding analysis suggests that the policymaker interested in both growth and distribution examines them separately at his peril. On the other hand, the results show that there is some room for manoeuver. It may be possible, through careful selection of a combination of policies, to improve both growth and distribution. But these policies may do nothing to reduce poverty. While we can speculate that improvements in growth and the income Gini index will enhance total welfare, we actually know little about how the selected policies will affect individuals or groups within the society. The Gini index can be though of as a measure of the area between the actual Lorenz curve and a 45-degree line, which represents perfectly equitable distribution. It does not contain any information about the *shape* of the underlying Lorenz curve. For any given Gini index (falling within the bounds of zero and one), the level of poverty may take on an infinite range of values. From a different perspective, a change in the Gini index may have no impact whatever on poverty, if the changes are limited to the top portions of the Lorenz curve. Conversely, policies may have an impact on poverty even without corresponding changes to the aggregate Gini index.

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[9] The calculated elasticity is larger for developed-country land distribution in the growth equation, but that parameter estimate is not significant.

In this section, we further explore the simultaneous determination of growth and inequality through changes in percentile-specific growth rates. This reveals income class-specific outcomes that may be hidden in analysis of aggregate measures of inequality such as the Gini index. It also serves as a robustness check on the earlier results – our confidence in the previous results and policy recommendations will increase if alternative specifications lead consistently to the same conclusion.

For these regressions we continue to work with periods of five years. Annual percentile income is the product of a five-year moving average share accruing to each group and annual GDP per capita; percentile growth is the average annual percentage change in this variable over five-year periods, as before. Thus, percentile growth differs from mean growth only when the percentile's share changes. Since shares change relatively little, the percentile growth rates are fairly closely correlated with each other and with that of mean income. Inspection of quintile growth rates, however, reveals that growth in the income of the poorest quintile is significantly more volatile than that of the other quintiles – the mean within-country coefficient of variation of the income share is 0.09 among the poorest compared with 0.06, 0.04, 0.03, and 0.05 for the other quintiles, respectively, as we move up the income scale.

Table 4 presents the results of Keane and Runkle 3SLS regressions of the variables from the expanded growth and equality models, including the interaction of land distribution with a dummy for developing countries, on growth for a range of income classes. Since the differences across quintiles are fairly small, and may be driven partly by measurement error, we aggregate across income classes and examine growth rates for the bottom 40 percent, middle 60 percent, and richest 40 percent of households. This weakens inference from comparison of parameter estimates between adjacent groups (since the categories overlap), but the results are both strong and striking.

The penultimate column in Table 4 presents chi-squared tests of joint significance across all groups. With one exception (initial income), the parameter estimates are jointly significant. The final column in the table is a chi-squared test of the null hypothesis that all parameter estimates are equal. Roughly speaking, we can say that variables which prove significant and equivalent across income classes influence aggregate growth but not relative distribution; conversely, variables which appear in this analysis to influence growth differentially across groups can be thought of as influencing aggregate distribution. Surprisingly, the final column reveals that with one exception (again, initial income), these parameter estimates differ significantly across income classes.

Although it is independently significant only for the wealthiest group, the monotonic collapse of the influence of education is quite impressive. If we understand the coefficients to be jointly significant, these results would suggest that expanded education increases growth significantly among the poor, while reducing it significantly among the rich. This in turn would imply that there is an education premium which is currently captured by the wealthy. Expanding education would reduce the supply of, and thus lower the average returns to, educated human capital. For the poorest 20 percent of households, however, even diminished returns to expanded education appear to be greater than the returns they can currently demand, and thus increasing income average growth among the poor. These results are consistent with the aggregate growth results, suggesting that education will not necessarily improve aggregate growth, since it lowers growth among the rich while increasing growth among the poor. They are also consistent with the aggregate inequality results, since at a minimum education will lower growth among the wealthy while not affecting the poor or middle classes.

The results for government expenditure also confirm the aggregate results for distribution, and provide some support for one tenet of the median voter hypothesis, that government expenditure is targeted to the middle classes. That government expenditure lowers growth among the poor is more surprising, and something which requires more detailed examination of actual benefit incidence. Similarly, higher values of the indicator of financial depth (M2/GDP) appear to lower growth among the poor and increase growth among the middle classes. It could be that the poor actually suffer from the development of formal financial intermediation,[10] but it is equally likely that the ratio M2/GDP is a poor proxy for the access by the poor to financial services.

Although there is a large negative correlation between inflation and growth among the rich, the correlation appears to be positive among the middle classes. The parameter estimate on inflation in growth among the poor is also negative, but insignificant. This begs the question of whether inflation creates rents that are captured by the middle classes. However, these results again most likely reflect non-linearities in the underlying data. When we include a squared term on inflation, both the level and the square are insignificant for the poor and middle groups, and both parameters are negative and significant for the rich.

The results for civil liberties in these regressions are the only ones clearly at odds with the aggregate results. Here, all parameter estimates on civil liberties are positive. An improvement in civil liberties (that is, a reduction in the index) leads to a decline in growth among all classes,

[10] This could happen if formal financial institutions displace informal ones without fulfilling the same role or providing the same services (see Aleem 1990).

although only the estimate for the poor is independently significant. Even if we accept the joint significance of all parameters, improved civil liberties would slow growth among the poor far more than among the rich. This implies that greater civil liberties would worsen both aggregate growth *and* aggregate distribution, whereas the aggregate results indicate no correlation with either outcome.

Among the most remarkable results from this analysis are those for the Sachs-Warner openness index and terms of trade changes. In accord with the aggregate results, openness is correlated negatively with growth among the poorest 40 percent, but strongly and positively with growth among the middle 60 percent and wealthiest 40 percent. While greater openness benefits the majority, it harms the poorest. These results show, in fact, that the costs of adjusting to greater openness are borne *exclusively* by the poor, regardless of how long the adjustment takes. In addition, the consequences of terms-of-trade changes are far greater for the poor than for the middle or wealthy classes. The poor are far more vulnerable to shifts in relative international prices, and this vulnerability is magnified by the country's openness to trade.[11] Considering that terms of trade have been falling on average for the countries in this sample, this bodes ill for the poor, and suggests that much more needs to be done to compensate those who lose from liberalization.

Finally, these results examine in greater detail the impact of more equitable land distribution on growth among different classes. Among developed countries, it appears that growth is correlated with more equitable land distribution only among the middle 60 percent. At first glance, this is rather surprising, especially together with the positive correlation of the land

[11] An added interaction between terms-of-trade changes and openness is negative for the poorest group and positive for the other two. Among the poorest quintiles, greater openness eliminates the positive correlation between growth and terms-of-trade changes.

Gini with the aggregate income Gini coefficient in developed countries. This result implies that farmers in developed countries are not among the poorest. In addition, it is a reminder that the Gini coefficient is a weak indicator of poverty. The positive correlation between the income and land distribution Ginis in developed countries may reflect the fact that more equitable land distribution shrinks the area in the middle of the Lorenz curve, leaving the bottom left corner (which represents distribution among the poor) untouched.

Among developing countries, there is a large and significant correlation between more equitable land distribution and faster growth among the poor. A one percent reduction in the Gini coefficient for land distribution leads to a 0.6 percent increase in growth for the poorest 40 percent of individuals. While this seems relatively large, it is outweighed in the aggregate by the positive correlation between growth and the land distribution Gini for the wealthiest 40 percent in developing countries. The point estimates indicate that the incomes of the rich grow four percent faster for every one percent increase in the land Gini.[12] The aggregate results indicate that more equitable land distribution increases equality at the cost of slower growth. These disaggegated results reveal that so far as land distribution is concerned, faster growth among the poor may indeed be obtained at the expense of slower growth among the rich.

For the middle 60 percent of households in developing countries there appears to be no relationship between land distribution and growth. Although the parameter estimate on the land Gini interacted with the developing-country indicator is significant, we cannot reject the hypothesis that the sum of the two land distribution variables is zero. In other words, although

[12] A chi-squared test can't reject the hypothesis that the sum of the developing-country land distribution parameter estimates for the poorest and wealthiest groups is zero; that is, they are of the same size but have opposite signs.

land distribution significantly affects the middle class in developed countries, it is insignificant to the middle class in developing countries.

This inspection of income-class-specific growth has revealed information about the distributional consequences of policies not captured in the aggregate analysis. Even in cases where the impact on overall growth and equality is relatively small, the consequences for specific groups may be quite large. For example, a one percent reduction in the land distribution Gini leads to a 0.5 percent reduction in income inequality in developing countries; but as we just showed, it would also lead to a 0.6 percent increase in income growth among the poor and it would reduce growth among the wealthy by four percent per year.

V. Conclusions

In this paper, we have drawn attention to a simple methodological point. The search for a mechanistic relationship between inequality and income has not yielded convincing empirical results. More importantly, it ignores the potential role of policy to advance both outcomes. On the other hand, when researchers have investigated the impact of policy on growth or inequality, they have done so by focussing on one outcome independently, but not both. The central message of this paper is that future research on growth and inequality should focus on their joint determinants, and especially those that are amenable to policy.

In the aggregate specifications reviewed here, we find evidence of trade-offs (government expenditure, openness, and land distribution), indicating that the joint analysis of inequality and growth yields policy-relevant information that cannot be obtained from analysis of each independently. At the same time, we find evidence of mutually exclusive policies, suggesting that the appropriate combination of policies can allow policy-makers to move towards their preferred location in the growth-equality space. We found no evidence, however, of mutually beneficial policies.

The disaggregated specification – differentiating the determinants of growth by income classes – provides additional details that would be unavailable from an exclusive focus on the aggregate relationships. While the aggregate results indicate that terms-of-trade changes affect aggregate growth, and are unrelated to aggregate equality, the disaggregated results reveal enormous disparity in the sensitivity of growth to terms-of-trade changes. The poor are far more vulnerable to international price movements than the aggregate results suggest.

Our analysis has also revealed two empirical regularities. First, there is greater variation in the relationship of policies to quintile growth rates than to aggregate growth rates. This is perhaps not surprising, since averages by definition eliminate idiosyncratic variance. Second, growth is much more sensitive than inequality to policy interventions. For example, from Table 3, calculated using the results of the aggregate models, the elasticity of growth with respect to the index of openness is on the order of 2.3, whereas with respect to inequality it is 0.3. In no case does any right-hand side variable have relatively larger influence on equality than on growth. This is of course fully consistent with experience over the last 30 years. Growth varies substantially from period to period, while inequality is much more persistent.

Our results for particular policies should be considered as illustrations rather than prescriptions. While we have focussed on the results that hold across specifications, our confidence in specific policy recommendations would be greatly increased were they to receive further empirical support from careful country studies. One result especially warrants further analysis. The adverse short-term consequences of openness to the welfare of the poor, and their vulnerability to terms-of-trade changes, compel us to pay greater attention to the disaggregated distributional consequences of policies in general and globalization in particular. At least in the short run, globalization appears to increase poverty and inequality, and as Rodrik (1997) has forcefully argued, the economic costs of exogenous shocks are magnified by distributional conflicts. Terms-of-trade shocks exacerbate social divisions, which in turn lead to greater volatility, which can further reinforce and widen distributional differences, and so on. Policies to minimize the adverse consequences of shocks to the poor are essential not only to enhance welfare but also to boost growth in an increasingly interdependent world economy.

Appendix 1. Sources of additions to Deininger-Squire dataset.

Malaysia 1973, 1984, 1987, 1989, 1992, 1995.
Vinod Ahuja, 1997, "Growth with Redistribution? Inequality and Poverty in Malaysia" (manuscript, World Bank). Data for 1984 through 1995 are from five successive rounds of the Household Income/Basic Amenities Survey (HIS/BA) conducted by the Economic Planning Unit of the Government of Malaysia. Data for 1973 is from the Household Income and Expenditure Survey, which was restricted to peninsular Malaysia and urban areas of Sabah and Sarawak.

Australia, 1989; Austria, 1987; Belgium, 1992; Canada, 1994; Denmark, 1992; Finland, 1991; France, 1989; Germany, 1989; Ireland, 1987; Israel, 1992; Italy, 1991; Luxembourg, 1991; Netherlands, 1991; Norway, 1991; Spain, 1990; Sweden, 1992; Switzerland, 1982; UK, 1986; USA, 1994.
LIS (Chen 1998): via Shaohua Chen.

Algeria, 1995; Brazil, 1995; Colombia, 1995; Costa Rica, 1996; El Salvador, 1995; Honduras, 1996; India, 1994; Indonesia, 1995; Mongolia , 1995; Panama, 1991; Papua New Guinea, 1996; Paraguay, 1995; Philippines, 1991; Sierra Leone, 1989; Venezuela, 1995; Yemen, 1992.
Chen 1998: all from national household surveys, via Shaohua Chen.

USA 1967-1996.
Bureau of the Census: www.census.gov.

The following data are from the LIS World Wide Web site. LIS Income Distribution Measures as computed by K.Vleminckx (Luxembourg Income Study - LIS), August 1998. Original data sources listed.

Australia 1981, 1985, 1989, 1994.
Australian Income and Housing Survey.

Belgium 1985, 1988, 1992.
Panel Survey of the Centre for Social Policy.

Canada 1971, 1975, 1981, 1987, 1991, 1994.
Survey of Consumer Finances

Czech Republic 1992.
Microcensus.

Denmark 1987, 1992.
Income Tax Survey.

Finland, 1987, 1991.

Income Distribution Survey.

France 1981.
CERC Survey of Women with Children.

France 1984.
Survey of Income from Income Tax.

France 1984, 1989.
Family Budget Survey.

Germany 1973, 1978, 1983.
Income and Consumer Survey (EVS).

Germany 1981.
The German Transfer Survey (Transferumfrage 1981).

Germany 1984, 1989, 1994.
German Social Economic Panel Study (GSOEP).

Hungary 1991, 1994.
Hungarian Household Panel.

Ireland 1987.
ESRI Survey of Income Distribution, Poverty and Usage of State Services.

Israel 1979, 1986, 1992.
Family Expenditure Survey.

Italy 1986, 1991, 1995.
The Bank of Italy Survey (Indagine Campionaria sui Bilanci Delle Famiglie).

Luxembourg 1985, 1991, 1994.
The Luxembourg Social Economic Panel Study (Liewen zu Letzebuerg).

Netherlands 1983, 1987.
Additional Enquiry on the Use of (Public) Services (AVO).

Netherlands 1991.
Socio-Economic Panel (SEP)

Norway 1979, 1986, 1991, 1995.
Income and Property Distribution Survey.

Poland 1986, 1992, 1995.

Household Budget Survey.

Taiwan 1981, 1986, 1991, 1995.
Survey of Personal Income Distribution, Taiwan Area.

Russia 1992, 1995.
Russian Longitudinal Monitoring Survey.

Slovak Republic 1992.
Slovak Microcensus.

Spain 1980, 1990.
Expenditure and Income Survey.

Sweden 1975, 1981, 1987, 1992.
Income Distribution Survey (Inkomstfördelningsundersokningen).

Switzerland 1982.
Swiss Income and Wealth Survey.

United Kingdom 1969, 1974, 1979, 1986, 1991, 1995.
The Family Expenditure Survey.

USA 1974, 1979, 1986, 1991, 1994.
March Current Population Survey.

The data below come from Branko Milanovic, 1998, *Income, Inequality and Poverty during the Transition from Planned to Market Economy*, WDC: World Bank.

Belarus, 1988.
National Statistical Office, Family Budget Survey. Not representative: sampling based on "branch of production," with workers and "typical" households over-represented; pensioners under-represented. Based on gross income, of which taxes comprise less than one percent. Includes in-kind income. Income definition incorrect: includes some items in income which are subtracted to compute income in Milanovic 1998. Decile ranks based on unadjusted income.

Belarus, 1995.
National Statistical Office, New Household Budget Survey. First quarter 1995. Includes in-kind income.

Estonia, 1988.
See Belarus 1988.

Hungary, 1987.

National Statistical Office, Household Budget Survey.

Hungary, 1993.
National Statistical Office, Household Budget Survey.

Kazakhstan, 1988.
See Belarus 1988.

Kyrgyz Rep, 1988.
See Belarus 1988.

Kyrgyz Rep, 1993.
Kyrgyzstan Multipurpose Poverty Survey. October-December 1993.

Latvia, 1988.
See Belarus 1988.

Latvia, 1995.
National Statistical Office, Household Budget Survey, fourth quarter 1995.

Lithuania, 1988.
See Belarus 1988.

Lithuania, 1993.
National Statistical Office, Household Budget Survey. No information on expenditures.

Moldova, 1988.
See Belarus 1988.

Moldova, 1993.
National Statistical Office, Household Budget Survey. No information on expenditures. Income definition incorrect: includes some revenue items (e.g. insurance compensation and sale of assets) that could not be deducted.

Poland, 1993.
National Statistical Office, Household Budget Survey, first half 1993.

Russia, 1988.
See Belarus 1988.

Russia, 1993.
Russian Longitudinal Monitoring Survey, round 3, third quarter 1993.

Slovak Rep, 1988.

National Statistical Office, *Microcensus* (Czechoslovakia). Recipients ranked by cash income only (exclusive of income-in-kind).

Slovak Rep, 1993.
National Statistical Office, Family Budget Survey. Not representative: excludes pensioner-headed households with economically active members and households headed by the unemployed. No information on the distribution of expenditures: households ranked by income.

Slovenia, 1987.
National Statistical Office, Household Budget Survey (Yugoslavia). Income definition incorrect: includes some items in income which are subtracted to compute income in Milanovic 1998. Decile ranks based on unadjusted income.

Slovenia, 1993.
National Statistical Office, Household Budget Survey. Income definition incorrect: includes some items in income which are subtracted to compute income in Milanovic 1998. Decile ranks based on unadjusted income. No information on the distribution of expenditures.

Turkmenistan, 1988.
See Belarus 1988.

Turkmenistan, 1993.
See Moldova 1993.

Ukraine, 1988.
See Belarus 1988.

Ukraine, 1995.
Ukraina 1995 Survey, June and first week of July, 1995. Personal income taxes comprise less than one percent of gross income.

Uzbekistan, 1989.
See Belarus 1988.

Uzbekistan, 1993.
See Moldova 1993.

Appendix 2: Estimation of Adjusted Gini Coefficient

In their original discussion of their inequality database, Deininger and Squire (1996) showed that the estimates of inequality differed systematically according to measurement method. Specifically, they found that the largest difference (6.6 points) occurred between the means of income-based and expenditure-based measures, and they did not find significant differences arising from the measurement of net versus gross incomes, or from household versus individual-based measures. Li, Squire and Zou (1998) used this result to adjust the Gini coefficients for their analysis, by adding 6.6 to the expenditure-based Ginis in their dataset.

Although this adjustment is a good first approximation, it exacerbates measurement error in the dependent variable to the extent that the mean difference does not accurately reflect the actual difference due to the choice of measurement method. Differences in the means are a combination of systematic measurement-related differences as well as actual cross-country differences in inequality. There may be country-specific measurement differences which are neither obvious in the mean nor captured by the categories provided, but which lead to great variation in specific estimates. For example, the definition of income or expenditure may vary across countries: expenditure surveys may include or exclude consumption away from home; income surveys may exclude in-kind income.

In order to capture some of these country-specific differences (which are likely to include factors both independent of and interacting with the measurement methods), we estimated a fixed-effects regression of the measured Gini on the measurement methods for which we have information. The results are shown in Appendix Table 1. In contrast to the bivariate Deininger-Squire results, this table shows that once the unmeasured country-specific differences are controlled, all of the explicit measurement differences are significant.

The mean Gini coefficient, controlling for measurement methods, is 35.6. This is the mean Gini when all of the dummy variables are set to zero; in other words, the Gini that would obtain if all inequality were calculated from individual- and expenditure-based measures. As expected, the Gini coefficients from household-based expenditures are significantly lower, and those for incomes are significantly higher. Surprisingly, and something which deserves more attention, this regression finds that Gini coefficients based on gross income are lower than those based on net income. This requires either that countries which use gross income to measure inequality are generally more equitable than those which use net income; or that net income is really more unequally distributed than gross income. The latter implies that on average, taxes and transfers are regressive.

Yet this method is not entirely satisfying either. There is one category, “other income” which includes both those observations which are truly other (undefined), and also those few observations for which the income type is unrecorded. These results will be biased to the extent that the unrecorded income types are not randomly chosen from the remaining categories.

Appendix 3: Sources and definitions of data used in regressions

Arable area World Bank World Development Indicators database: arable area, ha/cap. 1961-1995, annual.

Assassinations Political Risk Services, International Country Risk Guide database. 1982-1995, annual.

Civil liberties Gastil, R.D., (1987, 1990). Three observations: 1972-73, 1980, 1990.

Corruption Political Risk Services, International Country Risk Guide. 0-6 indicator. 1982-1995, annual.

Democracy Gastil, R.D., (1987, 1990). Two observations: 1970, 1980.

Education Barro and Lee (1996), average years of schooling for population aged 25 or older: tyr25. 1960-1990, 5-year averages.

Fertility rate World Bank World Tables database, ln(total fertility rate). 1960-1997, annual.

GDP Penn World Tables Mark 5.6 database, ln(RGDPCH), 1985 PPP$. 1960-1992, annual.

GDP growth Penn World Tables Mark 5.6 database, GDP_t-GDP_{t-1}

Gini Measurement-adjusted Gini (see appendix 3) from augmented Deininger-Squire dataset. 1960-1992, annual.

Gov't expenditure Penn World Tables Mark 5.6 database, government share of GDP: g. 1960-1992, annual.

Inflation International Monetary Fund International Financial Statistics database: change in ln(CPI). 1960-1994, annual.

Initial levels First available observation or value in 1960.

Land Gini Mean of values from FAO database, Klaus Deininger personal dataset, Heng-fu Zou personal dataset. 1960-1995, annual.

Life expectancy World Bank World Tables database, ln(life expectancy at birth). 1960-1997, annual.

M2/GDP World Bank World Tables database, 1960-1996, annual.

Openness Sachs and Warner (1995), 0-1 indicator. An economy is deemed to be open to trade if it satisfies four tests: (1) average tariff rates below 40 percent; (2) average quota and licensing coverage of imports of less than 40 percent; (3) a black market exchange rate premium that averaged less than 20 percent during the decade of the 1970s and 1980s; and (4) no extreme controls (taxes, quotas, state monopolies) on exports. 1960-1992, annual.

Quintile growth Log change in the product of moving average quintile share from Deininger-Squire dataset and national GDP. 1960-1992, annual.

Rule of law Political Risk Services, International Country Risk Guide database. 0-6 indicator. 1982-1995, annual.

Terms of trade $[\ln(\text{export price index})/\ln(\text{import price index})]_t$-$[\ln(\text{export price index})/\ln(\text{import price index})]_{t-1}$, base year 1990, World Bank World Tables database. 1960-1997, annual.

NB: all regressions use data averaged over five-year periods (e.g. 1960-64, 1965-69,...).

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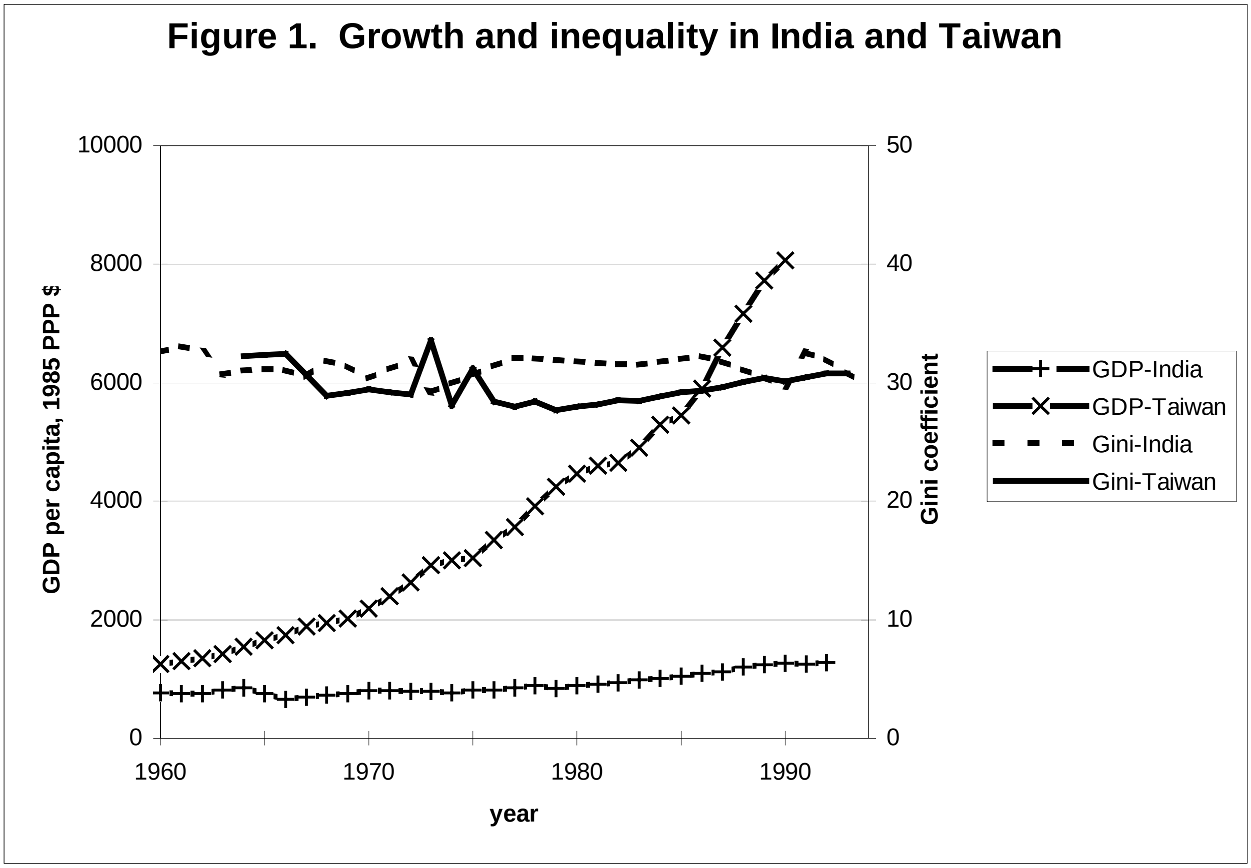
Figure 1. Growth and inequality in India and Taiwan
10000
8000
6000
4000
2000
0
50
40
30
20
10
0
1960
1970
1980
1990
year
GDP per capita, 1985 PPP $
Gini coefficient
GDP-India
GDP-Taiwan
Gini-India
Gini-Taiwan

Table 1. Estimation results -- base model

Estimation method	*SURE*		*3SLS*		*KR-3SLS*	
GDP growth						
[1] Education[a]	-0.05766	(0.78)	-0.05712	(0.71)	0.01167	(0.17)
[2] Govt expenditure[a]	0.00816	(0.23)	0.06184	(1.15)	0.24151	(4.92) ***
[3] M2/GDP[a]	0.00314	(0.21)	-0.00600	(0.29)	0.02258	(1.79) *
[4] Inflation[a]	-0.00080	(0.60)	0.00305	(0.93)	0.00548	(3.02) ***
[5] Openness[a]	3.16872	(5.35) ***	4.77710	(4.61) ***	4.53666	(4.19) ***
[6] Terms of trade change	14.41677	(0.66)	23.81896	(1.00)	68.77157	(1.77) *
[7] Initial ln GDP per capita[b]	-1.38430	(4.22) ***	-1.45163	(3.98) ***	-0.63853	(3.69) ***
[8] Dummy for 1980s	-1.47931	(2.81) ***	-1.73594	(2.91) ***	-2.16200	(4.74) ***
[9] Dummy for 1990s	-3.80816	(6.00) ***	-4.30083	(5.75) ***	-3.47344	(8.72) ***
[10] Intercept	12.17958	(4.84) ***	11.24762	(3.76) ***	0.19274	(1.74) *
Gini coefficient						
[11] Education[a]	-0.63902	(2.93) ***	-0.84788	(3.31) ***	-0.79236	(5.58) ***
[12] M2/GDP[a]	-0.11362	(2.77) ***	-0.05919	(1.06)	0.04516	(2.67) ***
[13] Civil liberties index[a,c]	1.06519	(2.13) **	1.97238	(2.85) ***	2.26477	(11.27) ***
[14] Ln land Gini[a,b]	0.25115	(6.33) ***	0.46375	(7.16) ***	0.53324	(34.22) ***
[15] Intercept	27.48767	(6.68) ***	10.90897	(1.76) *	0.12008	(3.27) ***
Chi2 test (df)						
[1] and [3], (2)		(0.62)		(0.63)		(5.31) *
[11] and [12], (2)		(19.86) ***		(14.42) ***		(36.59) ***
GDP Growth (9)		(84.82) ***		(70.99) ***		(1,687,376.00) ***
Gini coefficient (4)		(94.89) ***		(95.02) ***		(19,003.14) ***
Both equations (13)		(180.74) ***		(161.97) ***		(1,710,328.74) ***
Sargan test ~ F(15, 103)						
Growth				(0.63)		(0.80)
Gini coefficient				(1.83) **		(0.70)
Overidentification test ~ Chi2 (df)						
Growth (14)				(0.00)		(0.00)
Gini coefficient (11)				(0.01)		(0.00)
n (obs)		119		119		119
i (countries)		38		38		38

[a] Endogenous -- instruments include initial values of variables (except land Gini) plus population, urban population share, arable area, life expectancy at birth, total fertility, initial democracy, initial female literacy
[b] Mean value over entire sample period.
[c] Variables inverted: higher value indicates less favorable measure.
T-statistics in parentheses: * > 90, ** > 95, *** > 99 percent significance.

Table 2. Estimation results -- expanded model

	Growth			Gini coefficient			Joint significance tests Chi2 (2 df)	
Parameter estimates								
[1] Education[a]	-0.09543	(0.91)		-0.90076	(6.21)	***	(38.59)	***
[2] Govt expenditure[a]	0.23050	(3.29)	***	0.36332	(3.76)	***	(23.04)	***
[3] M2/GDP[a]	0.02105	(1.39)		0.09626	(5.26)	***	(28.78)	***
[4] Inflation[a,c]	0.00436	(2.55)	**	0.01824	(1.51)		(8.39)	**
[5] Openness[a]	7.88657	(4.41)	***	15.96615	(9.14)	***	(100.94)	***
[6] Terms of trade change	61.30023	(1.98)	**	43.58910	(0.62)		(4.06)	
[7] Initial ln GDP per capita[b]	-0.92452	(2.69)	***	0.18905	(0.30)		(7.39)	**
[8] Civil liberties index[a,c]	-0.03540	(0.10)		-0.51167	(1.00)		(0.99)	
[9] Land Gini[a,b]	-0.00291	(0.09)		0.12059	(1.70)	*	(2.93)	
[10] Land Gini X LDC dummy	0.04355	(2.19)	**	0.36746	(10.49)	***	(110.49)	***
[11] Dummy for 1980s	-1.86113	(4.78)	***	-1.93929	(1.77)	*	(23.66)	***
[12] Dummy for 1990s	-3.92063	(10.21)	***	-2.34694	(1.80)	*	(104.26)	***
[13] Intercept	0.04320	(0.40)		0.02733	(1.53)			
Joint significance tests, Chi2 (df)								
Both land Gini variables (2)		(4.80)	*		(130.04)	***		
All variables (12 df)		(2,388.56)	***		(616,444.40)	***		
Sargan test ~ F(15, 103)		(0.12)			(0.64)			
Overidentification test ~ Chi2(10 df)		(0.00)			(0.01)			
n (obs)		119			119			
i (countries)		38			38			

[a] Endogenous -- instruments include initial values of variables (except land Gini) plus population, urban population share, arable area, life expectancy at birth, total fertility, initial democracy, initial female literacy

[b] Mean value over entire sample period.

[c] Variables inverted: higher value indicates less favorable measure.

T-statistics in parentheses: * > 90, ** > 95, *** > 99 percent significance.

Table 3. Short-run elasticities of policy variables from aggregate regression results

	Growth	Gini coefficient	
Education	-0.234	-0.124	*
Govt expenditure	1.767	0.157	*
M2/GDP	0.398	0.102	*
Inflation	0.081	0.019	*
Openness	2.295	0.262	*
Terms of trade change	-0.007	0.000	
Civil liberties index	-0.047	-0.038	
Land Gini	-0.083	0.193	
Developing country land Gini	0.761	0.515	*

* Indicates jointly significant
Computed from results in Table 2.
†

Table 4. Estimation results – group-specific income growth

	Dependent variable -- growth of (average quintile share * average GDP per capita)							
	Bottom 40 %		*Middle 60 %*		*Top 40 %*		*Joint tests of parameters, Chi2(df)*	
	estimate	t-stat	estimate	t-stat	estimate	t-stat	*significance (3 df)*	*equality (2 df)*
Parameter estimates								
[1] Education[a]	0.14591	(0.80)	-0.05023	(0.65)	-0.24321	(2.37) **	(7.34) *	(5.65) *
[2] Govt expenditure[a]	-0.45601	(5.11) ***	0.16153	(3.25) ***	-0.05763	(0.76)	(39.00) ***	(38.96) ***
[3] M2/GDP[a]	-0.08944	(1.96) **	0.03819	(2.86) ***	-0.00904	(0.35)	(13.02) ***	(9.12) **
[4] Inflation[a,c]	-0.00370	(1.61)	0.00362	(1.79) *	-0.01217	(7.32) ***	(60.55) ***	(43.36) ***
[5] Openness[a]	-2.87840	(2.06) **	5.51345	(5.42) ***	9.89084	(2.90) ***	(38.55) ***	(28.53) ***
[6] Terms of trade change	250.17040	(3.12) ***	45.75256	(2.20) **	2.25016	(0.05)	(14.98) ***	(8.70) **
[7] Initial ln GDP per capita[b]	0.91370	(1.13)	-0.24779	(1.23)	-0.78018	(1.78) *	(5.87)	(4.00)
[8] Civil liberties index[a,c]	4.13267	(9.16) ***	0.10328	(0.40)	1.16411	(1.28)	(83.88) ***	(62.53) ***
[9] Land Gini[a,b]	0.04941	(0.73)	-0.06676	(3.17) ***	0.06457	(1.77) *	(17.09) ***	(13.02) ***
[10] Land Gini X LDC dummy	-0.12409	(4.05) ***	0.05350	(3.27) ***	0.14715	(2.74) ***	(34.18) ***	(33.09) ***
[11] Dummy for 1980s	3.55967	(2.93) ***	-1.45419	(4.64) ***	-2.66335	(5.81) ***	(59.71) ***	(28.74) ***
[12] Dummy for 1990s	4.14769	(2.42) **	-3.50908	(9.48) ***	-2.22664	(3.83) ***	(103.07) ***	(20.91) ***
[13] Intercept	0.00540	(0.25)	-0.08275	(0.97)	-0.08236	(3.20) ***		
Joint significance tests ~ Chi2(df)								
Both land Gini variables (2)		(17.47) ***		(13.82) ***		(14.74) ***		
All variables (12)		(6,546.49) ***		(611.21) ***		(7,511,902.00) ***		
Sargan test ~ F (15, 103)		(0.17)		(0.82)		(0.46)		
Overidentification test ~ Chi2 (7 df)		(0.00)		(0.00)		(0.00)		
n (obs)		119		119		119		
i (countries)		38		38		38		

[a] Endogenous -- instruments include initial values of variables (except land Gini) plus population, urban population share, arable area, life expectancy at birth, total fertility, initial democracy, initial female literacy.
[b] Mean value over entire sample period.
[c] Variables inverted: higher value indicates less favorable measure.
T-statistics in parentheses: * > 90, ** > 95, *** > 99 percent significance.

Appendix Table I. Fixed-effects estimation of adjusted Gini coefficients

variable	estimate	t-statistic
dummy = 1 if gross income	2.096	(1.83) *
dummy = 1 if net income	3.127	(3.06) ***
dummy = 1 if other income	5.762	(4.98) ***
dummy = 1 if household level	-3.171	(7.10) ***
intercept	35.592	(39.32) ***
F-test for fixed effects		(35.84) ***
F-test for parametric restrictions		(21.69) ***
Adjusted R^2		(0.03)

Computed from updated Deininger-Squire dataset.
*** > 99, * > 90 percent significance.
Omitted categories are individual-level and expenditure-based
Gini coefficients