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THE SIMULTANEOUS EVOLUTION OF GROWTH AND INEQUALITY*

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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, or level, of income and inequality. The other has tried to find causal explanations of growth and inequality, treating each 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 from previous independent studies.

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 (Anand and Kanbur, 1993), or to determine whether faster growth requires higher or lower inequality.[1] The second strand has sought to identify the causal factors influencing the evolution of growth and inequality independently. This research has looked either at growth, for example, (Barro and Sala-i-Martin, 1995) or at inequality (Li *et al.*, 1997), but has not tried to identify those factors that might simultaneously influence both growth and inequality.[2]

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 minimises 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 approach is particularly convincing from a theoretical standpoint: the evolution of growth and inequality must surely be the outcome of similar processes. In addition, neither approach is 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 and if so whether or not they are subject to the same set of determining factors.

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[1] 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 that conform to the Kuznets hypothesis, 11 that show inverse Kuznets curves, and 30 countries (more than 60%) in which there is no relationship.

[2] 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.

The paper is organised as follows. Section 1 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 2 we set out the equations to be estimated. 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 only two variables in common. Nevertheless, we test whether independent examination of even these limited models can be justified. The separate treatment of growth and inequality, however, also requires that, apart from the common variables, the other independent variables of the inequality equation have no impact on growth and those in the growth equation have no impact on equality. These variables are mutually exclusive by assumption. We therefore expand the simple models to first test the structural hypothesis that growth depends on distribution and distribution depends on growth.[3] Finally, we estimate a quasi-reduced-form model, including 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.

Our results, presented in detail in Section 3 and summarised in Section 4, reveal that some, though not all, of the variables assumed to be independent 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 best incomplete, results for the policymaker. The two variables – the Sachs-Warner measure of exchange rate and trade policies and the Gastil measure of civil liberties – that we find to be both robustly jointly and independently significant to each outcome involve trade-offs between growth and distribution. We find additional variables – the average years of schooling for the population aged 25 or over, the annual rate of inflation and the distribution of land (in developing countries) – that are at least jointly significant, including one, the last mentioned, that is possibly complementary. This is good from the perspective of policy, since it enables the policy maker to advance both outcomes simultaneously. Without exception, the variables, which we find independently correlated with only one of the outcomes, are also jointly significant. This suggests that the overall rate of growth cannot be separately identified from the distributional structure of growth.

1. 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 extended this acceptable set to 757 observations for 125 countries. The majority of the additions

[3] Using our restricted sample, fixed-effects regressions of one variable on the other only (with no additional controls) yield an elasticity of 3.5 of growth w.r.t. the Gini coefficient, and an elasticity of 0.012 of the Gini w.r.t. growth (both with $p = 0.06$).

cover Europe and Central Asia, although new data points were added to every region and income class. Our 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 available covariates 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.[4] For example, a measure of individual welfare may be based on the level of income or expenditure. In 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 details of the adjustment are available from the authors. 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. 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 4,300 observations on 147 countries.

We estimate our models on a relatively small sub-sample of countries. We do not 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 favourable) values.

2. Estimating Equations and Econometric Procedure

To see whether past practice of investigating growth and inequality independently leads to misleading empirical results, we draw on 'standard' growth and inequality models available in the literature. We will examine first the consequences of the implicit orthogonal assumptions across outcomes in these models when they are estimated and analysed separately. Under these circumstances, do the standard models imply affinity or conflict between growth and equality? Second, we expand the estimation by adding growth to the inequality equation, and inequality to the growth equation. Finally, we expand the estimation to include both sets of variables from the base model in each equation, to test explicitly the assumption that the determinants of growth are orthogonal to equality and vice versa.

[4] One of the criteria applied by Deininger and Squire to select acceptable data is that the series contain information on the methods used.

2.1. *Model Derivation*

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

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

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

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.

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 *et al.* (1993), we assume that there is no correlation in growth rates across periods, but we do allow for unobserved heterogeneity so that $u_{it} = \alpha_i + v_{it}$. 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 $u_{it} = \mu_i + v_{it}$ and $e_{it} = \rho e_{i,t-1} + \eta_i + \varepsilon_{it}$.

Our second set of estimations drops the orthogonal assumptions, allowing growth to enter the equality equation and equality to enter the growth equation:

$$\Delta y_{it} = \mathbf{S}'_{it}\boldsymbol{\alpha} + \mathbf{X}'_{it}\boldsymbol{\beta} + \lambda Gini_{it} + u_{it} \tag{3}$$

$$Gini_{it} = \mathbf{S}'_{it}\boldsymbol{\omega} + \mathbf{Z}'_{it}\boldsymbol{\psi} + \zeta\Delta y_{it} + e_{it}. \tag{4}$$

This allows at least a preliminary assessment of the impact of growth on equality and of equality on growth.

The estimates in (3) and (4) are, however, biased by multicollinearity, at least to the extent that vector $\mathbf{S}_{it}$ is correlated with the outcomes. This problem is made worse by any correlation between the assumedly orthogonal vectors ($\mathbf{X}_{it}$ and $\mathbf{Z}_{it}$) with the Gini coefficient and growth, respectively. We therefore solve (3) and (4) for the final (quasi-reduced) forms

$$\Delta y_{it} = \mathbf{M}'_{it}\boldsymbol{\beta}^* + u^*_{it} \tag{5}$$

$$G_{it} = \mathbf{M}'_{it}\boldsymbol{\delta}^* + e^*_{it}, \tag{6}$$

where $\mathbf{M} = [\mathbf{S}, \mathbf{X}, \mathbf{Z}]$ is the combined matrix of all right-hand-side variables from (3) and (4),

$$\boldsymbol{\beta}^* = \frac{\boldsymbol{\beta} + \gamma\boldsymbol{\delta}}{1 - \gamma\tau}, \quad \boldsymbol{\delta}^* = \frac{\boldsymbol{\delta} + \tau\boldsymbol{\beta}}{1 - \tau\gamma},$$

$$u^*_{it} = \left[(\gamma\eta_i + \mu_i) + \gamma\rho e_{i,t-1} + (\gamma\varepsilon_{it} + v_{it})\right]/(1 - \gamma\tau),$$

and

$$e_{it}^{*} = \left[(\tau\mu_i + \eta_i) + \rho e_{i,t-1} + (\tau v_{it} + \varepsilon_{it})\right]/(1 - \tau\gamma).$$

Note that (5) and (6) are not strictly 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 (6) will at least be inefficient.[5] Our solution is presented 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 enrolment, 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 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 (1997) (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 approximates the results presented in EKPS and LSZ; we choose these only because they are among the most parsimonious explanations of the (dependent) variables of interest.

In re-estimating 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.[6] EKPS include both primary and secondary enrolment, which we replace with one variable (mean years of schooling in the adult population). We also make 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. Finally, we add an interaction term between the distribution of

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

[6] Rodriguez and Rodrik (2000) argue that this is a poor measure of trade openness, representing instead a range of domestic, non-trade-related policies. For this reason, we do not regard the Sachs-Warner index as strictly a measure of openness. In addition, we experiment with other variables that have also been called measures of openness including the share of trade in GDP and the Summers-Heston index. See our discussion in Section 4.

land and a dummy for developing countries, on the basis that the impact of land distribution is likely to differ across countries according to the share and distribution of income from agriculture.

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. The second and third specifications relax this assumption and test whether the factors determining growth also influence the distribution of the gains from growth and vice-versa.

2.2. *Specification and Estimation Issues*

Prior to the publication of the Deininger and Squire database, multi-country studies of 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. $\bar{\Delta y}_i = \alpha_i + \beta_1 \bar{X}_{1i} + \beta_2 \bar{X}_{2i} + \varepsilon_i$, 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 over 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 institutions change and these changes have an impact on growth or equality within the sample period, a purely cross-sectional model will miss them.

With cross-sectional data, all unobserved cross-country variation is relegated to an error term (ε_{it}), which (if the errors are correlated with the regressors – i.e. $E(\varepsilon_{it} \mid x_{it}) \neq 0$) leads 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 that 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 exacerbates measurement error. We try to minimise 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 potentially 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 that might be of interest, since these methods simply purge the models of all cross-country differences.

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

$$y_{it} = \mathbf{B}'\mathbf{x}_{it} + u_{it}, \tag{7}$$

in which $\mathrm{E}(u_{it} \mid x_{it}) \neq 0$, KR assume that there exists a set of instruments $\mathbf{Z}$ which is at least *weakly* exogenous – that is, they are predetermined, but not necessarily uncorrelated with the unobserved country-specific effect, so that $\mathrm{E}(u_{it} \mid x_{it}) = 0, \forall s \leq t$, but $\mathrm{E}(u_{it} \mid x_{it}) \neq 0,\ \forall s > t$. In that case, KR suggests 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 (1997) (AB), but according to KR, the AB transformation does not eliminate all serial correlation and imposes the additional restriction of equicorrelation on the $\boldsymbol{\Omega}$ matrix.

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

$$\begin{pmatrix} \mathbf{y}_1 \\ \vdots \\ \mathbf{y}_N \end{pmatrix} = \begin{pmatrix} \mathbf{X}_1 \\ \vdots \\ \mathbf{X}_N \end{pmatrix} \mathbf{B} + \begin{pmatrix} \mathbf{u}_1 \\ \vdots \\ \mathbf{u}_N \end{pmatrix}, \tag{8}$$

where $\mathbf{y}_i$, $\mathbf{X}_i$, and $\mathbf{u}_i$ represent the $[T \times 1]$, $[T \times k]$ and $[T \times 1]$ observation matrices for the ith group. The KR estimator is constructed by first obtaining a consistent estimate of $\boldsymbol{\Sigma}_{TS} = \mathrm{E}(\mathbf{u}_i\mathbf{u}_i')$ for each group. Next, compute the upper triangular Cholesky decomposition of $\boldsymbol{\Sigma}_{TS}^{-1}$, called $\hat{\mathbf{P}}_{TS}$ (such that $\boldsymbol{\Sigma}_{TS}^{-1} = \mathbf{P}'_{TS}\mathbf{P}_{TS}$). Finally, premultiply (7) by $\hat{\mathbf{Q}}_{TS} = \mathbf{I}_N \otimes \hat{\mathbf{P}}_{TS}$, and estimate the transformed equation by 2SLS, using the original (untransformed) $\mathbf{Z}_{it}$ as instruments. The consistent estimate of $\boldsymbol{\Sigma}_{TS}$ derives from a 2SLS regression on the original untransformed equation:

$$\hat{\boldsymbol{\Sigma}}_{TS} = \frac{1}{N}\sum_{i=1}^{N} \left(\hat{\mathbf{u}}_{TS}\ \hat{\mathbf{u}}'_{TS}\right).^{7} \tag{9}$$

In our multiple-equation case, we find $\hat{\boldsymbol{\Sigma}}_{TS}^{\Delta y} = \frac{1}{N}\sum_{i=1}^{N}\left(\hat{\mathbf{u}}_{TS}^{\Delta y_i}\hat{\mathbf{u}}'_{TS}\Delta y_i\right)$ and $\hat{\boldsymbol{\Sigma}}_{TS}^{G} = \frac{1}{N}\sum_{i=1}^{N}\left(\hat{\mathbf{u}}_{TS}^{G_i}\hat{\mathbf{u}}'_{TS}G_i\right)$ from 3SLS on the untransformed equations (1) and (2) for the

[7] This is the formula for a balanced panel. In an unbalanced panel, the denominator is adjusted to the number of observations in each group.

base model, (3) and (4) for the 'structural' model, and (5) and (6) for the expanded, quasi-reduced-form model.

We instrument for the policy variables to correct for potential biases due to correlation with unobserved omitted characteristics and possible reverse causality. We use as instruments the initial values (as close to 1960 as possible) of all time-varying variables, as well as a range of other controls, including population, the urban share of population, life expectancy, total fertility, initial levels of female literacy and democracy, arable area and dummy variables indicating oil and commodity exporters and source of legal institutions (e.g. whether from English common law, German or French civil law etc.). As discussed above, we believe these variables to be at least weakly exogenous to the outcomes.

For instruments to be valid, they must be exogenous to the error term and significantly correlated with the variable they purport to represent. Bound *et al.* (1993) show that weak instruments can lead to large inconsistencies in the parameter estimates even if the instruments are only weakly correlated with the error in the structural equation. They suggest that instrumental variables estimates be accompanied by the partial r^2s and F-statistics on the excluded instruments from the first-stage regressions. We present these statistics and discuss the acceptability of our instruments at the end of the following section.

3. Results

3.1. *Results of the Standard Models*

Table 1 presents our approximations of the standard models.[8] In the original papers, each equation is of course estimated independently. 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.

For this sample, our standard growth results do not confirm the main hypothesis of EKPS, that terms of trade shocks are positively related to growth – they are insignificant in the all specifications. On the other hand, we do find that the Sachs-Warner index is consistently positively related to growth. As with EKPS, we fail to find significance for all policy variables, with the exception of education. We find education to be negatively correlated with growth in the KR-3SLS specification. This is surprising at first glance but not when we remember that education is most probably endogenous to growth – that is, growth causes education as much as education causes growth. These results are also consistent with recent empirical work by Pritchett (1996) who finds that education is negatively correlated with aggregate growth, in spite of the microeconomic evidence that education raises wages. Note that we estimate a positive coefficient on initial income in the KR-3SLS specification. This is at odds with conventional analysis. It may be a characteristic of this sample of countries; however, this oddity disappears once we relax the orthogonal assumptions in the estimations below, and include additional sources of variance in the growth model.

[8] All tables present heteroscedasticity-consistent Sargan over identification tests.

Table 1
Base models (equations 1 and 2)†

Estimation method	OLS estimation SURE		IV estimation 3SLS		KR-3SLS	
Growth						
[1] Education‡	−0.0497	(0.32)	−0.0282	(0.16)	−0.4294	(4.36)***
[2] Government‡	−0.0345	(0.97)	−0.0537	(1.20)	0.0225	(0.64)
[3] M2/GDP‡	0.0085	(0.52)	0.0085	(0.42)	0.0074	(0.58)
[4] Inflation‡	−0.0021	(1.56)	−0.0009	(0.35)	−0.0024	(1.22)
[5] S-W openness index‡	3.2632	(5.40)***	4.1782	(4.44)***	3.4397	(4.11)***
[6] Terms-of-trade changes	17.7454	(0.81)	26.0222	(1.13)	−7.6945	(0.39)
[7] Initial income	−1.0382	(2.13)**	−1.3847	(2.53)**	0.3162	(2.67)***
[8] Dummy for 1980s	−1.6155	(3.17)***	−1.5560	(2.86)***	−1.7545	(4.53)***
[9] Dummy for 1990s	−3.3395	(5.51)***	−3.4099	(5.22)***	−3.0173	(8.14)***
[10] Intercept	10.6042	(3.06)***	13.0454	(3.30)***	0.2208	(1.03)
Gini coefficient						
[11] Education‡	−1.0008	(2.71)***	−1.2413	(2.99)***	−0.7029	(2.60)***
[12] M2/GDP‡	−0.0570	(1.21)	−0.1207	(2.08)**	0.1200	(2.51)**
[13] Civil liberties‡	0.0348	(0.06)	0.2333	(0.32)	3.1295	(9.69)***
[14] Mean land Gini‡,§	0.1685	(2.70)***	0.2903	(3.34)***	0.2543	(4.81)***
[15] Mean land Gini × LDC‡,§	0.0344	(0.82)	−0.0532	(0.93)	0.1129	(2.75)***
[16] Intercept	33.9612	(7.60)***	33.2739	(6.05)***	0.1800	(0.71)
Marginal impact of land Gini in developing country	0.2029	(5.22)***	0.2371	(5.03)***	0.3672	(17.56)***
Joint significance tests $\sim\chi^2$, (df)						
All vars in growth eqn (9)		(80.36)***		(66.61)***		(2077.30)***
All vars in Gini eqn (5)		(102.58)***		(94.74)***		(1817.77)***
All vars (14)		(188.59)***		(163.80)***		(3858.40)***
Both land Gini variables (2)		(30.63)***		(26.43)***		(420.62)***
[1] and [3] (2)		(0.37)		(0.21)		(19.10)***
[11] and [12] (2)		(10.32)***		(14.78)***		(18.75)***
[1] and [11] (2)		(7.36)**		(8.95)**		(25.03)***
[3] and [12] (2)		(1.83)		(4.57)		(6.75)**
Sargan overid tests $\sim\chi^2$ (16)						
Growth				(3.59)		(4.91)
Gini				(3.69)		(5.00)
N (observations)		119		119		119
I (countries)		38		38		38

†t-statistics in parentheses: *** significant to <0.01, ** significant to <0.05, * significant to <0.1.
‡Endogenous; instruments include initial values of all variables (except land Gini and income), population, urban share, life expectancy, fertility, initial female literacy and democracy, arable area, dummies for oil and non-oil commodity exporters, and legal origin.
§Mean value over sample period.

Our standard inequality results agree broadly with those of LSZ, at least in sign. Differences between our results and LSZ are likely due to our use of a more appropriate error correction technique. These differences are illustrated in the differences across estimation techniques in Table 1. The impact of education on equality is consistently positive and significant across methods. Using the KR-3SLS results, an additional year of education, on average, lowers the Gini index by 0.7 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 improve income equality by 0.25 points on average, and 0.37 points in an average developing country. A one standard deviation (about 30%) improvement in land distribution would lower the Gini index by about six points (about 17%, a bit more than half of one standard deviation below the mean Gini). In line with LSZ, we also find that civil liberties are positively related to equality.[9] A one point (36%) reduction in the Gastil index, which implies an improvement in civil liberties, would lower the Gini index by two 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 that render the lag invalid as an instrument.

t-tests indicate that of the two common variables – education and the ratio of M2 to GDP – the former is significant in both equations, and the latter is significant only to equality. However, joint tests indicate that education and M2/GDP are jointly significant in both equations and that the variables are themselves jointly significant across equations. Thus, even at this initial level of the analysis when attention is confined to the common variables, the results indicate that growth and equality should be analysed jointly. Indeed, the results indicate that expanding education will lead to lower inequality at the cost of slower growth. This trade-off, however, need not overly concern the policy-maker, because on the basis of these results and their underlying assumptions, there are sufficient policy instruments available for the policy maker to pursue either goal without running the risk of compromising the other. Thus, the policy-maker can increase the Sachs-Warner index to promote growth without damaging equality, and similarly the policy-maker can improve civil liberties and/or redistribute land to the benefit of equality without compromising growth. The key question, of course, is whether this desirable outcome withstands further investigation of the joint determination of growth and equality.

3.2. *Results for the Structural Models*

The preceding analysis assumes that growth and distribution are orthogonal, at least in the variables from each equation that are excluded from the other. These regressions explicitly tested only those two variables the models have in common – education and M2/GDP. Of these two, education is certainly correlated with both outcomes, and M2/GDP may be. We now expand the analysis to ask whether we may be omitting important additional sources of variation in growth that might affect inequality, and vice-versa. To do this, we simply include growth in the Gini equation and the Gini index in the growth equation. This will indicate whether, controlling for the included factors, there are omitted components of each outcome that affect the other. These results are presented in Table 2.

[9] 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.

Table 2

Structural models (equations 3 and 4)

	OLS estimation		IV estimation			
Estimation method	SURE		3SLS		KR-3SLS	
Growth						
[1] Gini coefficient‡	0.0401	(1.33)	−0.0182	(0.40)	0.0435	(1.73)*
[2] Education‡	0.0005	(0.00)	−0.0641	(0.33)	−0.3934	(3.91)***
[3] Government‡	−0.0281	(0.76)	−0.0565	(1.24)	0.0147	(0.45)
[4] M2/GDP‡	0.0117	(0.70)	0.0084	(0.42)	0.0000	(0.00)
[5] Inflation‡	−0.0021	(1.53)	0.0002	(0.06)	−0.0037	(1.57)
[6] S-W openness index‡	3.4789	(5.76)***	4.1663	(4.32)***	4.3520	(5.33)***
[7] Terms-of-trade changes	17.0673	(0.78)	28.5080	(1.21)	-15.1709	(0.84)
[8] Initial income	−1.0693	(2.18)**	−1.3585	(2.45)**	0.0953	(0.54)
[9] Dummy for 1980s	−1.6533	(3.24)***	−1.5750	(2.87)***	−1.4698	(4.38)***
[10] Dummy for 1990s	−3.4159	(5.52)***	−3.3872	(5.12)***	−3.0835	(8.78)***
[11] Intercept	8.7326	(2.36)**	13.7311	(3.14)***	0.0627	(0.26)
Gini coefficient						
[12] Growth‡	0.4407	(1.91)*	0.2452	(0.69)	1.4120	(5.11)***
[13] Education‡	−0.9147	(2.49)**	−1.2191	(2.96)***	−0.8350	(2.92)***
[14] M2/GDP‡	−0.0564	(1.21)	−0.1168	(2.02)**	0.0488	(1.07)
[15] Civil liberties‡	−0.0313	(0.06)	0.1357	(0.18)	2.4740	(6.09)***
[16] Mean land Gini‡,§	0.1746	(2.80)***	0.2922	(3.37)***	0.2937	(4.49)***
[18] Mean land Gini × LDC‡,§	0.0470	(1.11)	−0.0440	(0.76)	0.1203	(2.62)***
[19] Intercept	31.8956	(7.12)***	32.3019	(5.78)***	−0.1475	(0.65)
Marginal impact of land Gini in developing country	0.2216	(5.54)***	0.2483	(4.98)***	0.4140	(14.44)***
Joint significance tests $\sim\chi^2$, (df)						
All vars in growth eqn (10)	(83.21)***		(66.36)***		(289.62)***	
All vars in Gini eqn (6)	(106.69)***		(96.48)***		(4148.24)***	
All vars (16)	(188.27)***		(157.00)***		(4467.08)***	
Both land Gini variables (2)	(34.30)***		(25.32)***		(343.41)***	
[2] and [4] (2)	(0.49)		(0.30)		(15.37)***	
[13] and [14] (2)	(9.05)**		(14.29)***		(10.07)***	
[2] and [13] (2)	(6.21)**		(8.85)**		(20.97)***	
[4] and [14] (2)	(1.86)		(4.28)		(1.22)	
Sargan overid tests $\sim\chi^2$ (15)						
Growth			(3.71)		(3.03)	
Gini			(3.81)		(4.98)	
N (observations)	119		119		119	
I (countries)	38		38		38	

Notes: see Table 1.

These results show that equality possibly adversely influences growth (at the 10% significance level). The remaining variables in the growth equation do not change. Growth, on the other hand, has a statistically strong and adverse effect on inequality. Quantitatively, however, the impact is small. A full, one percentage-point increase in the growth rate (from two to three percent per year) is correlated with only a 1.4 point increase in the Gini index, corresponding to an elasticity of 0.076. The remaining parameters in the Gini equation are generally unchanged, with the exception that the money supply (M2/GDP) falls from the model. Here we are beginning to see a slightly more nuanced picture of the relationship between these two outcomes. These results suggest that some previously omitted characteristics of growth are related to inequality, and vice versa. At the very least, the base models are probably guilty of omitted variables bias.

These structural models provide a rather blunt and potentially biased test of the relationship between the rate of growth and the distribution of income. We know that something about each outcome influences the other, but it is important for the policy maker that we understand better how they are related. What is it about growth that influences distribution? More importantly, what policy-driven components of growth might also have an impact on inequality? Is it greater openness, or greater government expenditure, or perhaps some other determinant of growth? Similarly, does more equitable land distribution or broader civil liberties enhance growth?

3.3. *Results for the Expanded Models*

To test these hypotheses, 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 (1) and (2), and estimate (3) and (4) in their final, quasi-reduced form – that is, equations (5) and (6). The results are presented in Table 3.[10]

The first thing to note is that the previous results for the growth equation are modestly robust to the expanded set of right-hand-side variables. Of the original variables in the growth equation, education falls out and no other growth variables become significant. The impact of the Sachs–Warner index remains unchanged. However, some of the additional variables from the equality model appear significantly correlated with growth. It turns out that civil liberties and land distribution are indeed related to growth. The Gastil index is positively correlated with growth: *lower* values of civil liberties (higher values of the index) lead to faster growth. In contrast (from a policy perspective), more equitable land distribution, on average, yields faster growth. Note that this is not necessarily true among developing countries, for which the marginal impact of land distribution on growth is essentially zero.[11]

The results for the equality variables in the expanded equality equation are identical in spirit to the previous results. Money supply remains insignificant, the impact of the Gastil civil liberties index is much smaller than previously estimated and the parameter estimate on the land Gini interaction is larger than that on the land Gini itself. Also, the partial derivative of inequality with respect to land distribution in developing countries is slightly smaller than previously estimated. A one-point reduction in the land Gini index would reduce the income Gini index by 0.09 points on average, and 0.25 points in the average developing country

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

[11] For developed countries, the partial derivative of growth with respect to land distribution is the parameter estimate on land distribution itself. For developing countries, the partial derivative is the sum of the parameter estimates on the land Gini and the interaction of the land Gini and a dummy for developing countries. While the parameter estimates are jointly significant, the linear combination of the two is not.

Table 3

Quasi-reduced-form models (equations 5 and 6)

Estimation method	IV estimation (KR-3SLS)				Joint significance
Dependent variable	Growth		Gini coefficient		(χ^2, 2 df)
[1] Education‡	0.0633	(0.67)	−1.0150	(2.13)**	(5.05)*
[2] Government‡	−0.0341	(0.93)	0.1027	(1.00)	(1.90)
[3] M2/GDP‡	0.0239	(1.48)	0.0474	(1.31)	(3.86)
[4] Inflation‡	−0.0042	(1.11)	0.0822	(2.30)**	(6.61)**
[5] S-W openness index‡	3.3779	(4.58)***	6.0108	(2.84)***	(28.65)***
[6] Terms-of-trade changes	22.8963	(1.39)	52.0990	(1.05)	(3.00)
[7] Civil liberties‡	0.5670	(4.24)***	1.3207	(2.59)***	(24.36)***
[8] Mean land Gini‡,§	−0.0440	(1.86)*	0.0899	(1.39)	(5.46)*
[9] Mean land Gini × LDC‡,§	0.0336	(2.15)**	0.1593	(4.69)***	(26.24)***
[10] Initial income	0.0382	(0.18)	2.5784	(3.54)***	(12.52)***
[11] Dummy for 1980s	−1.6374	(4.31)***	−1.1809	(1.33)	(20.15)***
[12] Dummy for 1990s	−3.2425	(7.99)***	−2.7460	(1.98)**	(67.25)***
[13] Intercept	0.1203	(0.58)	−0.3412	(1.23)	
Marginal impact of land Gini in developing country	−0.0104	(0.71)	0.2492	(5.75)***	
Joint significance tests $\sim\chi^2$, (df)					
Both land Gini variables (2)		(4.67)*		(89.27)***	(93.63)***
All vars, (12)		(204.66)***		(5278.86)***	(5474.72)***
Sargan overid tests $\sim\chi^2$ (13)		(4.28)		(4.86)	
N (observations)		119		119	
I (countries)		38		38	

Notes: see Table 1.

(corresponding to elasticities of 0.14 and 0.4, respectively). Education is strongly and negatively correlated with the income Gini index. Combined with the previous result that education is insignificant to aggregate growth, this suggests that broader education redistributes higher wages but does not increase productivity. Of the growth policy variables added to the equality equation, two – the Sachs–Warner index and inflation – are significant and inversely related to equality. The result for the Sachs-Warner index suggests that policies typically thought to promote growth may have some adverse consequences for distribution, at least in the short run. More happily for policy, lower inflation is correlated with greater income equality.

Taking these two equations together, there are two or possibly three policy variables that are independently correlated with both outcomes: the Sachs–Warner index, the civil liberties index, and the Gini index for land distribution among developing countries. The first two involve trade-offs between growth and distribution. Improving the policy stance captured by the Sachs–Warner index enhances growth while increasing the Gini index. The impact on growth, however, is much greater: a 10% improvement in the Sachs–Warner index increases the growth rate by 10% and worsens inequality by 1% (see Table 4). Improving civil liberties (reducing the index) will lead to a more equitable income distribution, but also to lower growth. Again, the impact on growth is much greater: a 10% improvement in civil liberties reduces the growth rate by 8% but increases equality by 1%. In contrast, improving land distribution is possibly correlated with more

Table 4

*Short-run elasticities**

Dependent variable	Growth	Gini coefficient
[1] Education	0.1698	−0.1468‡
[2] Government	−0.2735	0.0443
[3] M2/GDP	0.4719	0.0505
[4] Inflation	−0.0819	0.0856‡
[5] S-W openness index	1.0276	0.0985§
[6] Terms-of-trade changes	−0.0026	−0.0003
[7] Civil liberties	0.7900	0.0991§
[8] Mean land Gini	−1.3090	0.1442†
[9] Mean land Gini × LDC‖	−0.3083	0.3996§
[10] Initial income	0.1550	0.5637‡
[11] Dummy for 1980s	−0.3442	−0.0134†
[12] Dummy for 1990s	−0.3475	−0.0159§

*Calculated from results in Table 3.
†Significant to <0.1 for growth and jointly significant.
‡Significant to <0.1 for the Gini coefficient and jointly significant.
§Independently and jointly significant to <0.05 for both outcomes.
‖This is the elasticity with respect to land distribution in poor countries (i.e. the linear combination of the two-parameter estimates).

equitable income distribution *and* faster growth. We say possibly, since although the parameter estimate on the land distribution Gini index in developing countries is significant both to growth and to income distribution, the total partial correlation is the linear combination of the two, which is not significantly correlated with growth.

If we accept that the joint significance tests show the probability that a pair of parameter estimates is significantly different from zero, the interrelationship between growth and distribution becomes much greater. Using this generous (but plausible) standard, five of the seven policy variables tested are significant to both growth and inequality. Of these five, three (education, inflation, and land distribution) are correlated both with faster growth and lower income inequality. The remaining two – civil liberties and the Sachs–Warner index – poses a potential conflict between the goals of faster growth and more equitable distribution.

Considering the recent debate over the consequences of globalisation, this result for the Sachs–Warner index is worth discussing further. We make two points. First, these regressions provide point estimates of short-run changes; they do not estimate the long-run steady-state consequences of policy changes. If the costs of adjustment are borne disproportionately by the lower income classes and if, as is usually assumed, adjustment costs decline over time, while the dynamic gains from better and more stable policies continue *ad infinitum*, then the short-run adverse effects on equality may eventually be offset by faster growth among the poor. Second, recall that the Sachs–Warner index has been criticised as a measure of openness (Rodriguez and Rodrik, 2000). Consequently, we prefer to present our result not as a comment on openness *per se*, but as a comment on the package of policies – exchange rate, tariffs, and non-tariff barriers – captured by the Sachs–Warner index. Our results indicate that a shift in these policies in the appropriate direction can have a substantial, positive impact on growth and a much smaller, adverse effect on equality at least in the short run.

Dollar and Kraay (2000) argue that openness – in their case measured using the volume of trade to GDP – benefits the poor and the non-poor in exactly the same proportion. By implication, the volume of trade has no impact on inequality. We re-ran our final regressions using trade volume as the measure of openness and the Summers–Heston measure of openness; we find no significant correlation either with income distribution or with growth (results available). We cannot replicate the Dollar–Kraay results with our sample, and we speculate that our treatment of potential endogeneity is one cause of the discrepancy. Dollar and Kraay dismiss endogeneity concerns, arguing that it is unlikely that the share of income accruing to the poor has any influence on policies that affect the overall growth rate. There are some reasons to worry about endogeneity, however, in addition to the purely statistical issues mentioned earlier. Persson and Tabellini (1994) have found that the position of the median voter relative to the mean of the income distribution is a good predictor of the demand for policies that have an impact on growth as well as distribution. These policies, including openness, are then correlated with the error term in both regressions. Also, the period covered by our dataset (1960s to 1990s) has seen major policy and institutional changes among poor and middle-income countries, including changes in openness and income distribution. Since these other policies and institutions are changing over time, their influence on the included variables cannot be removed simply by differencing.[12]

Finally, Appendix Table 2 presents the Bound *et al.* (1993) test statistics on the first stage regressions for the results in Table 3. (Test statistics for the other regressions are available.) These tests show that the instruments are acceptable for all endogenous variables except inflation. The excluded instruments for inflation are significant in neither the growth nor the Gini equations. Other things being equal, endogeneity will bias parameter estimates upwards (that is, overestimating the variable's importance to the outcome), and measurement error will bias the estimates towards zero. For the growth equation, the KR-3SLS estimate of the impact of inflation is larger than the OLS estimate but neither estimate is anywhere near statistically significant. We conclude from this that even if inflation is poorly instrumented, the consequences for our growth results are insignificant. For the Gini equation, the KR-3SLS estimate is an order of magnitude larger than the OLS estimate, and both are significant (though the OLS estimate is significant only at ten percent). Again, this suggests that we can be confident in the results, at least in terms of the direction of the influence of inflation on income distribution, and the insignificance of inflation to growth in this sample.

4. Conclusions

In this paper, we have tried to draw attention to a simple methodological point. On the one hand, the search for a mechanistic relationship between inequality and

[12] We thank Martin Ravallion for this observation.

income 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.

Our main result is to show that the determinants of growth and inequality are not mutually exclusive. The consequences of policy choices are then much broader than previously assumed. The policy variables which are assumed to be exclusive – government expenditure, inflation, and the Sachs–Warner index for growth, and land distribution and civil liberties for equality – are not *a priori* exclusive. Only two of these variables are clearly significant independently in both equations and jointly across equations, and they involve trade-offs between growth and equality. Thus, a policy-maker who uses the results from the independent examination of growth to drive policy runs the risk of unwittingly adversely affecting income distribution. Our results reveal that increasing the Sachs–Warner index to promote growth would lead to greater inequity. Similarly, improving income distribution through enhancing civil liberties may have deleterious consequences for growth. Of the two variables included in both base models, education and M2/GDP, the expanded analysis suggests that the former is related exclusively to income distribution and the latter with neither outcome.

This still gives policy makers ample tools to use to pursue both goals of faster growth and greater equality. One can derive from these results a set of policies that will (*ceteris paribus*) in combination achieve almost any desired outcome in the growth-distribution space. For example, expanded education and more equitable land distribution will at least improve income distribution, and may also enhance growth. These policies could be used in combination with an increased value of the Sachs–Warner index to alleviate the distributional costs associated with 'openness'. Thus, growth could be improved without worsening income distribution by increasing the Sachs–Warner index (that is, increasing the proportion of time in a given period in which the country satisfies the Sachs–Warner criteria) by one standard deviation, and simultaneously improving the Gini index of land distribution by one standard deviation. This combination would nearly double the mean growth rate, and would also improve income distribution by four percent. These are of course very large policy shifts, but more modest policy changes could also yield improvements in both growth and distribution.

Finally, if we interpret the standard errors a bit more generously, we find that most of the policy variables tested in our final, reduced-form estimation are jointly significant to both outcomes. The few that are not jointly significant (government expenditure and money supply) are also not significant to either outcome independently. In the specifications tested above, we find no variables that robustly uniquely identify the determinants of growth or of distribution. This implies that conventional analysis, which looks at each outcome independently, fails in two respects. First, it ignores the evidence that policies designed to improve one outcome will probably also influence the other; and second, to the extent that an independent model is under identified, it can't even be entirely certain what it is estimating.

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Appendix: 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, (1987, 1990). Three observations: 1972–1973, 1980, and 1990.

Commodity exporting dummy World Bank World Tables database, = 1 if at least half of export revenue from primary commodities.

Corruption Political Risk Services, International Country Risk Guide. 0–6 indicators. 1982–1995, annual.

Democracy Gastil, (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.

Government expenditure Penn World Tables Mark 5.6 database, government share of GDP: g. 1960–1992, annual.

Inflation International Monetary Fund International Financial Statistics database: December-to-December 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.

Legal origin La Porta *et al.* (1998), source of legal institutions and traditions.

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.

Oil exporting dummy World Bank World Tables database, = 1 if at least half of export revenue from oil.

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.

Rule of law Political Risk Services, International Country Risk Guide database. 0–6 indicators. 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. 1960–1964, 1965–1969...).

Appendix

Table A1

Descriptive Statistics

Variable	Mean	Std. Dev.
Growth	2.039	3.012
Gini coefficient	37.842	9.411
Education	5.471	2.888
Government expenditure/GDP	16.338	7.143
M2/GDP	40.267	20.323
Inflation	39.411	171.843
Sachs-Warner openness index	0.620	0.470
Terms of trade change	0.000	0.010
Civil liberties	2.840	1.594
Mean land Gini	60.673	17.522
Initial GDP per capita	8.273	1.000
Dummy for 1980s	0.429	0.497
Dummy for 1990s	0.218	0.415
Initial education	4.095	2.610
Initial gov't expenditure	15.157	4.343
Initial M2/GDP	31.984	18.754
Initial inflation	9.743	20.907
Initial terms of trade change	0.003	0.015
Initial S-W openness	0.264	0.429
Initial civil liberties	3.092	1.809
Lagged terms of trade change	−0.001	0.018
Population (million)	76.200	155.000
Urban share of population	53.460	23.510
Ln life expectancy at birth	4.183	0.148
Ln total fertility rate	1.080	0.493
Initial female literacy rate	78.076	27.106
Initial democracy	2.966	2.046
Mean arable area (ha/cap)	0.346	0.407
Dummy for oil exporter	0.067	0.251
Dummy for commodity exporter	0.084	0.279
British legal origin	0.387	0.489
French legal origin	0.370	0.485
German legal origin	0.092	0.291
Scandinavian legal origin	0.126	0.333
Scandinavian legal origin	0.025	0.157

Table A2

*Significance of Instrumenting Equations**

	Growth		Gini coefficient	
	F-statistic	r^2	F-statistic	r^2
[1] Education	6.360***	0.258	82.874***	0.694
[2] Government	6.858***	0.267	21.762***	0.443
[3] M2/GDP	5.199***	0.235	61.712***	0.640
[4] Inflation	0.505	0.075	0.717	0.089
[5] S-W openness index	3.819***	0.203	9.929***	0.317
[7] Civil liberties	6.723***	0.265	16.771***	0.398
[8] Mean land Gini	6.291***	0.257	44.028***	0.575
[9] Mean land Gini × LDC	6.505***	0.261	88.501***	0.706

*F-tests and adjusted r^2 statistics of excluded instruments from the first-stage regressions for the models in Table 3. All are significant to 0.0001, with the exception of inflation.

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