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EXPLAINING INTERNATIONAL AND INTERTEMPORAL VARIATIONS IN INCOME INEQUALITY*

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This paper explores the propositions that, income inequality is relatively stable within countries; and that it varies significantly among countries. A new and expanded data set provides broad support for both propositions. Drawing on a political economy and capital market imperfection arguments to explain the intertemporal and international variation in inequality, the empirical analysis shows that the predicted variables associated with the first argument (a measure of civil liberties and the initial level of secondary schooling) and the second argument (a measure of financial depth and the initial distribution of land) are indeed important determinants of inequality.

This paper explores two propositions regarding income inequality. They are: first, income inequality is relatively stable within countries; and second, it varies significantly across countries.[1] To illustrate, note that the Gini coefficient in India remained almost constant for forty years (1951–92) with mean 32.6 and standard deviation 2.0.[2] In contrast, the variation in Gini coefficients across countries is large: 61.9 in Honduras in 1968 compared with 17.8 in Bulgaria in 1976. If substantiated, these propositions have potentially significant implications for poverty. The significance of the first is obvious – barring any fundamental socio-political change, poverty reduction will depend crucially on the rate of economic growth. Given this, the significance of the second is that in inegalitarian economies the poor will enjoy a smaller share of any national increment in income than in more egalitarian ones.

Drawing on a new and expanded data set on inequality (Deininger and Squire, 1996*a*), the first of the paper's three sections conducts standard statistical tests of the two propositions. The sample comprises 573 observations on the most common measure of inequality – the Gini coefficient – for 49 developed and developing countries covering the period 1947–94. The results broadly confirm our two propositions. Specifically, analysis of variance (ANOVA) shows that about 90% of the total variance in the Gini coefficients

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[1] We also explore a weaker, combined version of these two propositions – namely, that intertemporal shifts in inequality are modest compared with international differences.

[2] The mean Gini coefficient for India reported in Table 2 is 39.15. This is after the data have been adjusted for difference in definitions. The mean for the unadjusted data is 32.55.

can be explained by variation across countries, while only a small percentage of the total variance is due to variation over time. Similarly, regression analysis reveals significant differences across countries, and fails to detect any significant time trend in 32 countries. Moreover, in 10 of the 17 cases where the data reveal a significant trend, it is quantitatively small – an annual change of less than 1.0% of the country's average Gini coefficient. To take a typical example, Jamaica shows a statistically significant and negative time trend but the change in its Gini coefficient from its 1980 value of 49.9 would be only 0.2 points a year. At this rate, it would take Jamaica 70 years to bring its Gini index in line with the average index for all countries in our sample – 36.2. In this sense, the observed intertemporal changes are small relative to the observed differences across countries. On the other hand, seven of the countries in the sample have annual changes in excess of 1.0% indicating that in certain circumstances inequality as measured by the Gini index can change more quickly – in China the index was increasing during our sample period at a rate of 3% a year, the largest rate of change that we observed. What actually happened in these seven countries is an interesting issue for future research.

In general, our results suggest that inequality is determined by factors which differ substantially across countries but tend to be relatively stable within countries. The second section of the paper explores some possible determinants of inequality. To do so, it draws on two ideas that have recently received attention in the literature on inequality and growth. The first posits a link from income or wealth inequality to policy via a political economy argument. In its simplest form, the rich are assumed to have the resources to lobby for policies which are beneficial to them but may be harmful to the rest of the economy and to growth (see Bertola, 1993). The second idea has to do with imperfections in the market for credit. By preventing the poor from making productive investments (such as education), credit constraints arising from asymmetric information perpetuate a low and inequitable growth process (see, for example, Banerjee and Newman, 1991). Taken together, the two ideas suggest that an initial state of inequality may be expected to continue because the rich have the capacity to protect their wealth while the poor are unable to augment theirs.

We find considerable support for both these ideas. In particular, the key variables associated with the political economy argument (a measure of political freedom and initial secondary schooling) and those associated with the capital market imperfection (the initial degree of inequality in the distribution of assets as measured by the distribution of land and a measure of financial market development) are all shown to be significant determinants of current inequality. This suggests that the rich are indeed able to exercise sufficient control over economic policy at least to maintain their wealth while the non-rich encounter capital market imperfections that limit their capacity to accumulate capital, again reinforcing the tendency for unequal distributions of income to remain so. To check the robustness of our main findings, we conduct sensitivity analysis by controlling for various other factors identified in both theoretical and empirical work on inequality and growth. The results suggest that our findings are quite robust. Section 3 concludes by linking our results to previous work on the relationship between growth and inequality.

1. Testing the Two Propositions

This paper uses a new data set on Gini coefficients.[3] Starting with a total of 2,480 observations on Gini coefficients covering 112 developed and developing countries for the years 1947–94, several criteria were used to 'cleanse' the data. First, all observations had to be from national household surveys for expenditure or income; second, the coverage had to be representative of the national population; and third, all sources of income and uses of expenditure had to be accounted for, including own-consumption. In addition, for the purpose of this paper, all observations had to be from countries with at least four observations covering a reasonable part of the 47 year period. These procedures resulted in a sample of 573 observations covering 49 countries. This is the data set used in this study. Before presenting the sample descriptive statistics, we note two points.

First, the definition of what is being measured by the Gini coefficient in our sample varies across countries. Inequality can be measured by gross income, net income, or expenditure and it can be per capita or per household. The distribution of our sample by definitional differences is shown in Table 1. Because variation in definition can undermine the international and intertemporal comparability of the data, we include controls for different definitions throughout Section 1. The results indicate that differences between coefficients defined on net and gross income and between household-based and individual-based coefficients are not significant. Differences between expenditure-based and income-based coefficients, however, are significant. In subsequent analysis, therefore, we have adjusted the data following the procedure recommended by Deininger and Squire (1996*a*). Specifically, we adjust for differences between income-based and expenditure-based coefficients by systematically increasing the latter by 6.6 points, this being the average difference observed by Deininger and Squire (1996*a*).

Table 1
Distribution of Gini Coefficients by Different Definitions

	Income			
Unit of observation	Gross	Net	Expenditure	Total
Household	240	73	19	332
Individual	102	78	61	241
Total	342	151	80	573

[3] For further details see Deininger and Squire (1996*a*).

Second, the method used to calculate the Gini coefficients also varies across different sources. To minimise this problem we have recalculated the coefficients using a standard technique for as many observations as possible.[4]

With these points in mind, basic descriptive statistics for the adjusted data are reported in Table 2. Here we simply note that the overall sample mean is 36.2 and the standard deviation is 9.2. The number of observations per country is as follows: 28 countries have between 4 and 9 observations; 14 countries have between 10 and 19 observations; and 7 countries have 20 or more. In general, the developed countries have longer series and better coverage than the developing ones. As a preliminary test of our two propositions, note that the standard deviation of the means of the 49 countries (9.3) is substantially greater than any of the standard deviations of the within-country Gini coefficients for each country (see Table 2).

We begin with an analysis of the variance components of the Gini coefficients using the raw data. Table 3 reports the ANOVA results. Allowing for the fact that we have an unbalanced data set, we find that for the entire sample (Data set 1), 91.8% of the variance is cross-country variance, while only 0.85% is over-time variance. A total of 0.4% is due to the differences in definitions. Based on the F-values, only the country variation and variation due to income/expenditure definition are significant. After adjusting for the income/expenditure definition differences as described above (Data set 2), the ANOVA results show that the variance due to income/expenditure drops from 0.34% to 0.04% and is statistically no longer significant. This provides statistical evidence that the adjustment is necessary and useful.

Disaggregating the data (unadjusted) by income level according to the classification in the World Development Report, we obtain similar results. For high-income countries (Data set 3), the cross-country variance is 82.5% and the over time variance is only 1.9%. The corresponding figures for lower- and middle-income countries (Data set 4) are 93.1% and 1.4%. In this case, the variation due to income/expenditure definition is significant since most of the differences in income/expenditure occurs in this group. But, with the adjusted data, remaining variations due to definition are small and insignificant. We also repeated the ANOVA exercise for three subsamples in which we progressively increased the consistency of definitions: a subsample containing Gini coefficients based only on income (493 observations); a subsample containing only income-based and household-based Gini coefficients (313 observations); and a subsample containing coefficients defined on gross income per household (239 observations). In each case, the results (not reported) indicate that about 90% of the variation can be explained by country variation, while variation over time is small (1.1 to 1.7%).

We now turn to a least squares dummy variable regression which allows us to study individual country specific effects and perform explicit hypothesis testing

[4] The computational tool (POVCAL) we used for recalculating the Gini coefficients is discussed in detail in Datt (1992).

Table 2

Summary Statistics of Gini Coefficients (Adjusted for differences in definitions)

Code	Nob	Mean	St. dev.	Max	Min	Max–min	Coverage
AUS	9	37.88	2.91	41.72	32.02	9.70	69 ~ 90
BEL	4	27.00	0.76	28.25	26.22	2.03	79 ~ 92
BGD	10	35.83	1.55	39.00	33.34	5.66	63 ~ 92
BGR	28	23.30	3.34	34.42	17.83	16.59	63 ~ 93
BHS	11	45.77	3.91	54.09	40.64	13.45	70 ~ 93
BRA	14	57.84	2.82	61.94	53.00	8.94	60 ~ 89
CAN	23	31.27	1.64	32.97	27.41	5.56	51 ~ 91
CHL	5	51.84	5.15	57.88	45.64	12.24	68 ~ 94
CHN	12	32.68	3.62	37.80	25.70	12.10	80 ~ 92
COL	7	51.51	2.48	54.50	46.00	8.50	70 ~ 91
CRI	9	46.00	2.80	50.00	42.00	8.00	61 ~ 89
CSK	12	22.25	2.29	27.19	19.37	7.82	58 ~ 92
DEU	7	31.22	1.58	33.57	28.13	5.44	63 ~ 84
DNK	4	32.08	1.09	33.20	30.99	2.21	76 ~ 92
DOM	4	46.94	2.90	50.46	43.29	7.17	76 ~ 92
ESP	8	32.85	1.73	37.11	31.02	6.09	65 ~ 89
FIN	12	29.93	2.08	32.04	26.11	5.93	66 ~ 91
FRA	7	43.11	5.62	49.00	34.85	14.15	56 ~ 84
GBR	31	25.98	2.56	32.40	22.90	9.50	61 ~ 91
HKG	7	41.58	2.60	45.18	37.30	7.88	71 ~ 91
HND	7	54.49	3.36	61.88	50.00	11.88	68 ~ 93
HUN	9	24.65	3.36	32.24	20.97	11.27	62 ~ 93
IDN	11	40.09	2.07	45.19	37.30	7.89	64 ~ 93
IND	31	39.15	2.03	43.65	35.77	7.88	51 ~ 92
IRN	5	49.83	1.26	52.05	48.48	3.57	69 ~ 84
ITA	15	34.93	2.52	41.00	32.02	8.98	74 ~ 91
JAM	9	48.77	2.84	54.31	44.52	9.79	58 ~ 93
JPN	23	34.82	1.32	37.60	32.50	5.10	62 ~ 90
KOR	14	34.19	2.54	39.10	29.82	9.28	53 ~ 88
LKA	9	42.45	4.52	47.80	35.30	12.50	53 ~ 90
MEX	9	54.59	2.76	57.90	50.00	7.90	50 ~ 92
MYS	6	50.36	1.79	53.00	48.00	5.00	70 ~ 89
NLD	12	28.59	0.91	29.68	26.66	3.02	75 ~ 91
NOR	9	34.21	2.73	37.52	30.57	6.95	62 ~ 91
NZL	12	34.36	2.78	40.21	30.04	10.17	73 ~ 90
PAK	9	38.10	0.81	39.04	36.51	2.53	69 ~ 91
PAN	4	52.42	4.34	57.00	47.47	9.53	70 ~ 89
PHL	7	47.62	2.27	51.32	45.00	6.32	57 ~ 91
POL	17	25.69	2.44	33.06	20.88	12.18	76 ~ 93
PRT	4	36.60	0.59	37.23	35.63	1.60	73 ~ 91
SGP	6	40.12	1.65	42.00	37.00	5.00	73 ~ 89
SWE	14	31.74	1.42	33.41	27.31	6.10	67 ~ 92
THA	8	45.48	3.54	51.50	41.28	10.22	62 ~ 92
TTO	4	46.21	3.28	51.00	41.72	9.28	58 ~ 81
TUN	5	49.11	1.26	50.60	46.84	3.76	65 ~ 90
TWN	26	29.62	1.50	33.60	27.70	5.90	64 ~ 93
USA	45	35.28	1.27	38.16	33.50	4.66	47 ~ 91
VEN	9	44.42	4.02	53.84	39.42	14.42	71 ~ 90
YUG	10	32.62	0.95	34.73	31.18	3.55	63 ~ 90
Overall	573	36.23	9.15	61.94	17.83	44.11	47 ~ 94

Note: Nob – Number of observations.

Table 3
Analysis of Variance of Gini Coefficients

Data set (Nob)	Source	DF	Sum of Squares	F-Value	Source	DF	Sum of Squares	F-Value	% of Total
1	Model	98	42,837.71	64.84	Country	48	42,262.78	130.61	91.81
(573)	Error	474	3,195.24		Time	47	389.10	1.23	0.85
	Total	572	46,032.95		Income	1	157.00	23.29	0.34
					Hhld.	1	27.91	4.14	0.06
					Gross	1	0.93	0.14	0.00
2	Model	98	44,716.93	67.69	Country	48	44,280.97	136.85	92.42
(573)	Error	474	3,195.24		Time	47	389.22	1.23	0.81
	Total	572	47,912.17		Income	1	17.90	2.66	0.04
					Hhld.	1	27.91	4.14	0.06
					Gross	1	0.93	0.14	0.00
3	Model	68	6,627.54	17.52	Country	19	6,449.03	61.00	82.49
(283)	Error	214	1,190.81		Time	46	149.67	0.58	1.91
	Total	282	7,818.35		Income	1	19.98	3.59	0.26
					Hhld.	1	2.99	0.54	0.04
					Gross	1	5.88	1.06	0.08
4	Model	75	33,320.84	55.98	Country	28	32,618.22	146.80	93.14
(290)	Error	214	1,698.23		Time	44	492.43	1.41	1.41
	Total	289	35,019.07		Income	1	164.73	20.76	0.47
					Hhld.	1	34.14	4.30	0.10
					Gross	1	11.32	1.43	0.03

Notes: 1: DF – Degrees of freedom
2: Descriptions for different data sets:
Data set 1: The whole sample, including all definitions.
Data set 2: The whole sample, adjusted for difference in income or expenditure definitions.
Data set 3: Subsample, high-income countries, including all definitions.
Data set 4: Subsample, low- and middle-income countries, including all definitions.

concerning the two propositions. Because we have seen that standard deviation of within-country Gini coefficients is small and because a plot of the Gini coefficients by country revealed trends for some countries, we consider a simple linear trend model:

$$g_{it} = \phi_i D_i + \theta_i t_i + \delta_1 d_1 + \delta_2 d_2 + \delta_3 d_3 + \omega_{it} \tag{1}$$

where g_{it} is the Gini coefficient, $i = 1, 2, \ldots, N$ (number of countries), $D_i = 1$ for country i and 0 otherwise, $t_i = 1, 2, \ldots, T_i$, and $\omega_{it} \sim \text{iid}(0, \sigma^2_\omega)$. The panel data are unbalanced since in general $T_i \neq T_j$ for $i \neq j$. In light of the ANOVA results, we use the adjusted data but include definitional dummies to test for any remaining effect. d_1 is the control dummy for income $(= 1)$/expenditure $(= 0)$; d_2 is the control dummy for households $(= 1)$/individual $(= 0)$; d_3 is the control dummy for gross income $(= 1)$/net income $(= 0)$.

We would like to know whether, after controlling for the differences in definitions, the difference between the country specific effects $\phi_1, \phi_2, \ldots,$ and ϕ_N is statistically significant or not and we want to test for the existence of a significant, within-country time trend, $\theta_1, \theta_2, \ldots, \theta_N$. Accordingly, we test the following two hypotheses:

(a) H_0^a: $\phi_1 = \phi_2 = \ldots = \phi_N$,
(b) H_0^b: $\theta_i = 0$, for $i = 1, 2, \ldots, N$.

Based on the F-statistic in Table 4, hypothesis (a) is rejected at a 5% significance level. This confirms our first proposition – Gini coefficients differ significantly across countries. For individual time trends, we find statistical support for our second proposition in 32 of the 49 countries or 65% of the

Table 4

LSDV Estimation Results (Unrestricted Regression)

Dep var.: GINI

Control Dummy	Income	Households	Gross
Estimate	4.00	−0.35	−0.80
t-value	1.70	−0.61	−1.13

Country Code	Country-specific	t-value	Trend Estimate	t-value	Country Code	Country-specific	t-value	Trend Estimate	t-value
AUS	34.18	13.90	0.35	3.04*	ITA	31.48	12.96	−0.37	−3.31*
BEL	23.62	8.03	−0.10	−0.44	JAM	49.77	59.58	−0.20	−2.77*
BGD	32.71	13.21	0.01	0.12	JPN	31.49	13.07	−0.02	−0.40
BGR	20.58	8.05	0.20	4.04*	KOR	31.89	12.89	0.10	1.68
BHS	43.47	17.77	−0.30	−3.65*	LKA	39.21	18.59	−0.09	−1.22
BRA	54.76	22.20	0.05	0.61	MEX	51.60	20.11	0.01	0.25
CAN	27.89	11.66	−0.06	−1.35	MYS	46.96	18.45	−0.17	−1.19
CHL	48.65	18.51	0.51	5.04*	NLD	24.20	9.74	0.11	0.83
CHN	24.72	8.70	0.80	4.54*	NOR	29.79	12.05	−0.22	−2.73*
COL	48.40	19.03	0.00	0.00	NZL	30.14	12.36	0.49	4.15*
CRI	42.71	17.01	−0.16	−1.74	PAK	38.10	52.02	0.06	0.71
CSK	18.65	7.12	−0.04	−0.60	PAN	49.48	18.25	−0.01	−0.08
DEU	28.34	11.23	0.13	1.10	PHL	43.83	17.24	−0.11	−1.66
DNK	27.85	10.44	0.21	1.13	POL	21.12	8.14	0.31	2.95*
DOM	42.31	14.51	0.34	1.88	PRT	34.96	18.67	−0.31	−1.62
ESP	33.26	40.47	−0.13	−1.24	SGP	36.89	14.61	0.01	0.09
FIN	25.98	10.54	−0.14	−1.35	SWE	27.71	11.34	0.01	0.13
FRA	34.22	12.97	−0.58	−6.40*	THA	42.16	17.09	0.31	4.10*
GBR	23.07	8.87	0.18	4.21*	TTO	41.76	14.90	−0.15	−1.15
HKG	38.44	15.47	0.21	1.71	TUN	49.57	43.57	−0.04	−0.37
HND	54.17	20.29	−0.37	−3.50*	TWN	25.97	9.99	−0.02	−0.43
HUN	21.04	8.02	0.14	1.92	USA	32.76	13.75	0.06	2.48*
IDN	40.69	45.87	−0.04	−0.50	VEN	41.16	16.23	0.27	2.26*
IND	38.41	50.02	−0.10	−3.29*	YUG	29.60	11.53	0.06	0.61
IRN	49.96	30.19	−0.03	−0.18					

NOB	573	R^2	0.95
DF	472	F−Test	126.03
Groups	49		

Notes: 1. Standard errors of individual country specific effects: 9.89.
2. For hypothesis (a), the F-statistic is 126.03. This leads to the rejection of hypothesis (a).
3. For hypothesis (b), 7 countries have significant negative trend, 10 countries have significant positive trend. (There is a total of 17 countries with significant trends (indicated by (*).)
4. The country-specific terms are equivalent to the 1980 predicted Gini coefficients.

sample.[5] For 7 countries, however, we find significant negative trends, while 10 countries have significant positive trends when we apply the 5% t-test. But of the 17 countries with a significant trend, 10 of them have time trends that are quantitatively small – defined here as an annual change of less than 1.0% of the country's 1980 predicted Gini coefficient, the estimated country-specific term in the regression reported in Table 4. This is, of course, an arbitrary cut-off point. We note, however, that applying the mean absolute rate of change (0.6% a year) for these 10 countries to the average Gini coefficient (36.2) for our entire sample, it would take more than 20 years for the index to move 5 points. This compares with a difference between the maximum and minimum 1980 predicted values for each country of 36.1 points. Thus, whether or not one considers movements of 0.6% a year quantitatively large, it is clear that intertemporal changes are very modest compared with international differences.

For seven countries (Australia, Chile, China, France, Italy, New Zealand, and Poland), however, we observe a statistically large and quantitatively important time trend, thus establishing that countries can change the degree of inequality as measured by the Gini coefficient relatively quickly. For example, the results for New Zealand indicate an annual change of 1.6% implying that a change of 5 points in the Gini index could be achieved in only 10 years. The factors affecting changes of this magnitude in these 'non-conforming' countries present an interesting opportunity for future research. Here we simply note that four of the countries – Australia, France, Italy, and New Zealand – are OECD countries where the fiscal system in general and the welfare system in particular are well developed so that, given the political will, it should be feasible to influence inequality. And in the remaining countries, China and Poland have of course been experiencing major structural changes during the period covered by our sample.

Because we have less than 10 observations for 28 countries, the test for a significant trend may not be accurate. Reducing the sample to only those countries with 10 or more observations, however, yields broadly similar results. For the 21 countries with at least 10 observations, the time trend is insignificant for 12 countries, and in the 9 countries where the time trend is significant it is quantitatively important (an annual change of more than 1% of the mean) in only four countries – China, Italy, New Zealand, and Poland. For the group of 21 countries, the average absolute trend is 0.16, or an average absolute percentage change of 0.52% per year. For the 7 countries with at least 20 observations, the results are even stronger. Three countries have an insignificant time trend, and of the other 4 none have a quantitatively important trend.

[5] Since most of the countries do not have enough observations to allow for suitable unit root tests, we have not pursued this approach. For the United States and United Kingdom, there are observable positive time trends since late 1970's and early 1980's. Since 1970, inequality in the United States has increased at a rate of 0.62% a year, while in the United Kingdom, the increase has been 1.37% a year. For the US data which have the longest time series, the simple Augmented Dickey-Fuller test suggests the presence of a unit root. However, in Raj and Slottje (1994), they found that the US Gini is stationary around a broken trend.

In fact, the average absolute trend for this group is only 0.1, while the average absolute percentage change is only 0.32% per year. Thus, for the countries where we have the most complete and reliable data, inequality appears to be quite stable over relatively long periods of time.

Recall that the results reported in Table 4 use the adjusted data for the Gini coefficient. With these data, we see that the definitional dummies are not significant. We also obtain broadly similar results for our two propositions (not reported) for the unadjusted data and in the subsample of observations with the same definition.

We have also estimated a random-effects model (with or without a time trend and using the adjusted data) to account for the loss of degrees of freedom in the LSDV regression. We assume that the country-specific effects are drawn from a random distribution, while at the same time controlling for the definition differences. The results are presented in Table 5. The only significant explanatory variable is the constant term with an estimated value of 37.7, close to the sample average. As a result of the large variation in Gini coefficients across countries, the constant has a standard error between 9.71 (without the time trend) and 9.82 (with the time trend) that is very close to the standard error of the country-specific effects (9.89) in the LSDV regression. Thus, the constant term in the random-effects model plays the same role as the country-specific terms in the LSDV regression, lending support to our assertion that the variations in inequality arise mainly from cross-country differences. Note also that the random-effects model provides no support for a general time trend: the mean value of the time trend is not significant with a 95% confidence interval of (−0.015, 0.035).

Taken together, these results provide substantial support for our two propositions. Thus, Gini coefficients are clearly different across countries (propo-

Table 5

Random-effects Regression of Gini Coefficients

Dep var.: GINI

	Model I		Model II	
	Estimate	t-value	Estimate	t-value
Constant	37.67	17.31	37.74	17.23
Income	1.07	0.57	1.03	0.55
Households	−0.43	−0.84	−0.52	−1.00
Gross	0.77	1.20	0.78	1.21
Trend			0.01	0.80
Nob	573		573	
DF	569		568	
Groups	49		49	

Notes: Standard errors of error terms: Individual constant terms: 9.71 (Model I), 9.82 (Model II).
Standard errors of error terms: White noise error: 2.63 (Model I), 2.63 (Model II).

sition one) and there is no evidence of a time trend in 32 countries or 65% of our sample (proposition two). In addition, the absolute magnitude of the time trends for the whole sample are quantitatively small (the biggest being 0.79 and the absolute average 0.18) compared with the cross-country variation. Moreover, for 10 (20% of the sample) of the 17 countries where we do identify a significant time trend, the data still support a weaker version of our two propositions – namely, that within-country intertemporal variation in the Gini coefficient is small relative to the variation across countries. At the same time, our data provide support for larger and faster changes in inequality in seven countries (14% of the sample). Thus, for the majority of countries – 42 out of 49 – inequality changes at best very slowly, suggesting that structural factors – economic, social, political, and demographic – play a crucial role in determining the level of inequality in a country and its evolution over time.

2. The Determinants of Inequality

In this section, we draw on two ideas that have received attention in the recent literature; see Benabou (1996) for a survey and further extensions. The first embeds the determination of policy in a political economy framework. This serves two functions: it renders policies endogenous, thus allowing us to focus on structural variables; and it gives those who wield political power the capacity to protect their wealth. And the second introduces private investment and a credit market imperfection that effects the accumulation of capital. This also serves two functions: it makes investment endogenous, again allowing us to focus on structural variables; and it makes it difficult for those without access to formal credit to accumulate capital. Thus, these two ideas fit well with our initial empirical conclusions – they point to the importance of structural variables which can differ markedly across countries but change only slowly within countries, and they introduce political forces and market imperfections that would tend to preserve an existing uneven distribution of wealth.

For simplicity, we focus on the behaviour of the richest segment of society, here called the 'rich minority', and the interaction between them on one side and the rest of society – the middle class and the poor termed here the 'poor majority' – on the other. Unlike the majority-rule models in Alesina and Rodrik (1994) and Persson and Tabellini (1994), we assume that the top group (the minority) can influence economic policy, not through the voting mechanism, but through its economic power (bribes) or through direct political control (most of the political leadership will be members of this group). The ability of the top group to influence power is, however, constrained by the degree of democratisation or political liberty and by the extent of education.

This approach reflects two considerations. First, the median-voter theory has not been well supported empirically. According to this theory, the median voter's distance from the average capital endowment in the economy will increase with wealth inequality, thus leading him or her to approve a tax rate that is higher the more unequal the distribution of wealth, which in turn reduces investment and economic growth. Thus in a democracy we would expect higher inequality to be associated with lower growth. This is not substantiated by recent empirical work (Deininger and Squire, 1996*b*). On the contrary, they find that initial inequality is associated with lower growth in non-democratic countries, a conclusion that is consistent with the political economy mechanism modelled here. And second, for many of the countries and periods covered by our data the political setting cannot be characterised as democratic. Thus, we assume that, if sufficiently powerful, the rich minority can capture allocations of foreign exchange and credit, it can protect its own enterprises from foreign or domestic competition, it can influence public spending in favour of higher education and tertiary health care, and so on. While the interventions can take a multitude of forms, they basically amount to a tax on the rest of society. Such a characterisation may be appropriate for many of the developing countries included in our sample. For the more developed countries, the scope for exerting influence over policy would be much more constrained.

Both groups – the rich minority and the poor majority – can invest in capital. But, the presence of credit constraints arising from information asymmetries limits the ability of individuals to make productive investments in lumpy capital (Galor and Zeira, 1993, Banerjee and Newman, 1991, Greenwood and Jovanovic, 1990). In these circumstances, ownership of a collaterable asset can provide access to formal credit markets and cheaper credit. Since the rich are more likely to own collaterable assets, the cost of investing to the poor typically exceeds that to the rich. Thus, the poor have less incentive to invest in capital (because of the 'tax' imposed by the rich) and face higher costs of financing investment (through the credit market imperfection).

2.1. *Data and Estimation Methodology*

From the political economy argument, we experiment with two variables: a measure of civil liberties that can be expected to constrain the capacity of the rich to influence policy; and the initial level of secondary schooling on the assumption that a more educated population is able to exert a restraining influence on policy. For the poor's access to the financial market, we explore two variables – the depth of the financial sector and the initial distribution of land taken to be a collaterable asset. We expect that a more developed financial sector and a more equal distribution of land will ease the access of the poor to credit.

We use the adjusted data averaged over 5 year-periods in our empirical analysis for two reasons. First, although for most of the variables we have yearly observations, our data on Gini coefficients is more limited. Recall from Section 2 that for most of our 49 countries there are only a limited number of observations on income inequality – 28 countries have 4 to 9 observations each, while only 7 countries have more than 20 observations. By using 5-year averages we achieve a more balanced panel data set. Second, because our aggregate measures of inequality are relatively stable over time, using 5-year averages will not result in much loss of information. However, for other variables use of 5-year averages will reduce short-run fluctuations and allow us to focus on the structural relationships of most interest to us.

2.2. *Base Regression*

Based on the above considerations we test a parsimonious base regression of the form:

$$g_{it} = \alpha + \beta_1 MYSC60_i + \beta_2 CIVLIB_{it} + \beta_3 LDGINI_i + \beta_4 FNDP_{it} + u_{it} \qquad (2)$$

where $i = 1, 2, \ldots, N$ (country index), $t = 1, 2, \ldots, T$ (time index) and u_{it} is independent and identically distributed regression error. g_{it} is the Gini coefficient.[6] *MYSC*60 is the initial mean years of secondary schooling (1960 data), *CIVLIB* is the civil liberty index, *LDGINI* is the initial Gini coefficient for the distribution of land, and *FNDP* is a measure of financial development (defined as $M2/GDP$).[7] We expect *MYSC*60, *CIVLIB*, *LDGINI*, and *FNDP* to have a negative, positive, positive, and negative effect respectively on income inequality. Table 6 summarises the results. According to the OLS results, all four variables have the right sign and are significant.

We conducted a standard test of serial correlation for the base regression residuals. The DW statistic indicates serial correlation and so we reestimated the regression with an AR(1) error specification. The coefficient estimates and their significance remain largely unchanged after correction of serial correlation. The base regression may, however, suffer from an endogeneity problem. Financial depth as measured by the ratio of *M*2 to *GDP* – may be subject to policy influence and may therefore be endogenous. That is, the very factors that allow the rich to control certain economic policies may also allow them to control various policies influencing the development of the financial market. For this reason, we reestimate the base regression using the instrumental variables method (IV) with lagged variables for the instruments. The IV estimates are in general similar to the OLS or the AR(1) estimates (see Table 6).

To provide a quantitative appreciation of these results, we also show in Table 6 the standardised coefficient estimates (based on the IV method) for each of the independent variables. The key result to emerge is that the variables associated with the financial market imperfection argument have a much greater influence on inequality than those associated with the political economy argument. Thus, the joint effect of a one-standard-deviation increase in financial depth and a one-standard-deviation reduction in the inequality of land distribution results in a reduction of 5.05 points in inequality. On the

[6] We have also estimated the base regression using the ratio of the top quintile's income share to that of the bottom 80% of the population as the dependent variable with little change in the results.

[7] Data description and sources: *MYSC*60 – Human capital stocks, Nehru *et al.* (1995). *CIVLIB* – There are two indices in Gastil (various issues) measuring civil liberties and political rights. Since the two are highly correlated, we only use one of them – the civil liberty index – but interpret it broadly to capture both civil liberties and political rights. We use the five year average index for the period 1972–89 as reported in Barro and Lee (1994). The index is defined from 1 to 7, with 1 assigned to countries with the largest degree of civil liberties. *LDGINI* – World Bank. *FNDP* – Based on M2 and GDP data, IFS (IMF).

Table 6

Base Regression Estimation Results

Dep var.: GINI

	OLS		AR(1)		IV		
Variable	Estimate	t-value	Estimate	t-value	Estimate	t-value	Std. Est.
Constant	32.81	13.79	33.47	12.59	34.20	14.17	—
*MYSC*60	−4.55	−5.80	−5.29	−5.47	−4.36	−5.60	−0.32
CIVLIB	1.61	5.47	1.13	3.65	1.54	5.27	0.30
LDGINI	0.16	6.64	0.16	5.70	0.15	6.54	0.32
FNDP	−7.73	−3.17	−6.42	−2.49	−10.07	−3.90	−0.23
Nob:	166		166		166		
$\bar{R}^2$:	0.62		0.77				

Notes: The DW−statistics for OLS is 0.743.
The errors in AR(1) is specified as an AR(1) process. The estimated AR(1) coefficient is 0.65.
Instruments for *FNDP* in IV: lagged value *FNDP*.
Std. Est.: Standardised estimates. Same for Table 7 below.

other hand, the impact of a one-standard-deviation increase in civil liberties and a one-standard-deviation increase in secondary schooling yields only a reduction of 2.75 points in inequality.

2.3. *Regressions for the Poor and the Rich*

Our data set also allows us to run separate regressions for the incomes of the rich minority and the poor majority. We use the real income of the top quintile of the population to represent the income of the rich minority, and the real income of the bottom 80% to represent that of the poor majority. Using the real income variable from Summers and Heston (1995) and the share data from Deininger and Squire (1996*a*) we can easily obtain the income of the poor and the rich. We run the following regression for both the poor and the rich:

$$y_{it} = \alpha + \beta_1 MYSC60_i + \beta_2 CIVLIB_{it} + \beta_3 LDGINI_i + \beta_4 FNDP_{it} + v_{it} \quad (3)$$

where y_{it} is the income of the poor or the rich. Table 7 reports the results. As before we use the AR(1) specification for serial correlation and the instrumental variables method to deal with endogeneity.

The results for the income of the poor provide strong support for both the political economy argument and the capital market imperfection argument. The OLS results are robust to AR(1) specification or instrumental variables method. All coefficients are highly significant and have the right signs. The results for the rich show that coefficients on financial depth, the civil liberties index and secondary schooling are significant. The results are highly plausible. Development of the financial sector should benefit the rich as well as the poor, as should the degree of secondary education. The result with respect to civil liberties could be argued either way for the rich – political freedom and civil liberties could foster overall growth as well as constraining the ability of the

Table 7

Estimation Results for the Poor's and the Rich's Income Regressions

Dep var.:The Poor's Income ($000 dollars)

	OLS		AR(1)		IV		
Variable	Estimate	t-value	Estimate	t-value	Estimate	t-value	Std. Est.
Constant	5.02	6.17	4.17	5.28	4.68	5.62	—
*MYSC*60	1.44	5.73	1.37	5.10	1.41	5.68	0.32
CIVLIB	−0.77	−7.15	−0.51	−4.88	−0.75	−6.99	−0.40
LDGINI	−0.03	−3.58	−0.03	−4.05	−0.03	−3.50	−0.18
FNDP	3.59	4.16	5.09	6.81	4.13	4.50	0.27
Nob:	144		144		144		
$\overline{R}^2$:	0.66		0.84				

Notes: The DW-statistics for OLS is 0.59.
The errors in AR(1) is specified as an AR(1) process. The estimated AR(1) coefficient is 0.74.
Instruments for *FNDP* in IV: lagged value *FNDP*.

Dep var.: The Rich's Income ($000)

	OLS		AR(1)		IV		
Variable	Estimate	t-value	Estimate	t-value	Estimate	t-value	Std. Est.
Constant	2.48	4.52	1.82	3.38	2.39	4.27	—
*MYSC*60	0.85	4.99	0.84	4.59	0.84	5.01	0.32
CIVLIB	−0.40	−5.47	−0.19	−2.74	−0.39	−5.44	−0.36
LDGINI	0.002	−0.40	−0.08	−1.44	0.002	−0.36	−0.02
FNDP	2.12	3.66	3.47	6.91	2.26	3.66	0.26
Nob:	144		144		144		
$\overline{R}^2$:	0.54		0.78				

Notes: The DW−statistics for OLS is 0.61.
The errors in AR(1) is specified as an AR(1) process. The estimated AR(1) coefficient is 0.75.
Instruments for *FNDP* in IV: lagged *FNDP*.

rich to influence policy in their favour. Of interest is that the coefficient on the distribution of land is not significant for the rich. Because all of the rich would be expected to have access to credit markets, it would be reasonable to suppose that holdings of collateral assets would not be a significant determinant of the incomes of the rich.

Finally, we note that this disaggregation allows us a clearer understanding of the results that were obtained with inequality as the dependent variable. Taken together, the results for the base regression and the current regression suggest that a more egalitarian distribution of land benefits the poor but not the rich thus leading to improvements in inequality. And the results show that expansion of political liberties and secondary education and greater financial depth affect income growth for both the rich and the poor in the same direction, but in a way that also reduces inequality. Thus, from Table 7, we note that the joint impact of a one-standard-deviation decrease in the civil liberty index and a one standard-deviation increase in both initial secondary schooling and financial depth results in an increase of \$3,000 dollars in the incomes of the poor but only an increase of \$1,600 dollars in the incomes of the rich.

Consistent with our overall approach, the independent variables behave like structural variables in the sense that they are different across countries but relatively stable within countries. For example, the Gini coefficient for the distribution of land in Thailand is 45.4 compared with 92.4 in Venezuela. Increasing Thailand's coefficient to that of Venezuela would increase the Gini coefficient from its fitted value of 41.8 to 49.6, an increase of about 7.8. On the other hand, substituting the final-period value of the variable for financial depth for its initial-period value, the Gini coefficient in Korea falls by only 0.2 (from 38.9 to 38.7), in Sri Lanka by 0.5 (from 42.6 to 42.1), and in Indonesia by 1.0 (from 47.2 to 46.2).

2.4. *Sensitivity Analysis of the Base Regression*

To complete our tests of robustness, we undertake a sensitivity analysis. We conduct a stepwise regression analysis by adding other variables discussed in the literature to the base regression. The results are summarised in Table 8.[8] We consider the following variables: initial real per capita GDP (*INIGDP*), gross domestic investment ratio (*INVSHR*), urbanisation ratio (*URB*), black market premium (*BMP*), terms of trade shocks (*TOTSK*), openness (*XGDP*) and per capita arable area (*AAREA*).[9] Two results emerge. First, throughout the sensitivity test, our four key variables keep the right sign, remain significant, and have values for the estimated coefficient similar to those in the base regression. Our results appear to be quite robust. Second, initial per capita income is the only additional variable which turns out to be significant in all the sensitivity regressions. The results indicate that higher-income countries tend to have more equal distributions of income. Given the relative constancy of inequality over time despite in some cases significant increases in income, this result suggests that initial income is proxying for a range of cross-country structural differences at the start of the sample period. All other variables are consistently insignificant.

The stepwise regressions may be subject to change according to the order in which the variables are added. We therefore conduct a sensitivity analysis similar to that of Levine and Renelt (1992). We select three variables from the pool of seven variables each time, add these three variables to our base regression, and see whether the parameters in our base regression are stable or not. This procedure gives a total of 36 regressions. The summary results are presented in Table 9.

[8] Since the AR(1) results and the IV results are quite similar to the OLS results, they are therefore not presented to conserve space.

[9] Data description and sources: *INIGDP* – Initial real per capita GDP, Summers and Heston (1994). *INVSHR* – Gross domestic investment to GDP ratio, National Accounts (World Bank). *URB* – Urbanisation ratio, World Tables (World Bank). *BMP* – Black market premium, Barro and Lee, (1994). *TOTSK* – Terms of trade shocks, defined as $\Delta \ln$ (Export price) – $\Delta \ln$ (Import price), IFS (IMF) and Trade Statistics (World Bank). *XGDP* – Export to GDP ratio, IFS (IMF). *AAREA* – Per capita arable area, World Bank.

Table 8

Sensitivity Analysis(I)

Dep var.: GINI

	Base Reg.	1	2	3	4	5	6	7
Constant	32.81	35.62	33.54	33.54	33.55	33.62	35.84	34.72
	(13.787)	(13.318)	(10.303)	(10.187)	(10.030)	(10.026)	(9.541)	(9.251)
*MYSC*60	−4.55	−3.51	−3.48	−3.48	−3.46	−3.51	−3.98	−3.96
	(−5.803)	(−3.874)	(−3.835)	(−3.820)	(−3.791)	(−3.823)	(−4.058)	(−4.087)
CIVLIB	1.62	1.15	1.14	1.14	1.21	1.20	1.15	1.11
	(5.472)	(3.236)	(3.208)	(3.197)	(3.321)	(3.304)	(3.168)	(3.073)
LDGINI	0.16	0.16	0.17	0.17	0.17	0.17	0.15	0.15
	(6.636)	(6.543)	(6.488)	(6.266)	(6.262)	(6.263)	(5.158)	(4.993)
FNDP	−7.74	−7.64	−8.61	−8.61	−8.64	−8.67	−7.56	−8.45
	(−3.171)	(−3.126)	(−3.322)	(−3.306)	(−3.283)	(−3.287)	(−2.826)	(−3.099)
INIGDP		− 0.54	−0.60	−0.60	0.62	−0.62	−0.80	−0.82
		(−2.225)	(−2.418)	(−1.967)	(−2.010)	(−2.008)	(−2.531)	(−2.344)
INVSHR			0.10	0.10	0.08	0.08	0.05	0.04
			(1.118)	(1.082)	(0.902)	(0.837)	(0.583)	(0.374)
URB				0.00	0.01	0.01	0.03	0.05
				(0.008)	(0.148)	(0.181)	(0.921)	(1.444)
BMP					−1.09	−1.15	−1.19	−1.39
					(−0.851)	(−0.896)	(−0.939)	(−1.106)
TOTSK						−1.33	−1.10	−1.22
						(−0.629)	(−0.522)	(−0.583)
XGDP							−0.06	−0.01
							(−1.551)	(−0.249)
AAREA								−0.02
								(−0.278)
Nob	166	163	163	163	162	162	159	155
$\overline{R}^2$:	0.62	0.61	0.61	0.61	0.61	0.60	0.62	0.62

Note: The t-statistics are reported in parentheses.

Table 9

Sensitivity Analysis (II)

Dep var.: GINI

		Estimate	t−value	$\overline{R}^2$	Extra variables		
Constant	Minimum	31.76	9.736	0.60	*INIGDP*	*BMP*	*TOTSK*
	Base Reg.	32.81	13.787	0.61			
	Maximum	37.54	11.963	0.62	*INIGDP*	*URB*	*BMP*
*MYSC*60	Minimum	−4.97	−5.756	0.60	*INIGDP*	*BMP*	*AAREA*
	Base Reg.	−4.55	−5.803	0.61			
	Maximum	−3.46	−3.800	0.61	*INIGDP*	*INVSHR*	*URB*
CIVLIB	Minimum	1.05	2.962	0.62	*INIGDP*	*INVSHR*	*AAREA*
	Base Reg.	1.62	5.472	0.61			
	Maximum	1.64	5.241	0.60	*INIGDP*	*BMP*	*TOTSK*
LDGINI	Minimum	0.14	5.336	0.62	*INIGDP*	*INVSHR*	*XGDP*
	Base Reg.	0.16	6.636	0.61			
	Maximum	0.17	6.457	0.60	*INIGDP*	*URB*	*XGDP*
FNDP	Minimum	−9.23	−3.445	0.61	*INIGDP*	*BMP*	*XDGP*
	Base Reg.	−7.74	−3.171	0.61			
	Maximum	−6.70	−2.693	0.62	*INIGDP*	*URB*	*TOTSK*

Table 9 shows the maximum and minimum values of the estimates (based on OLS), together with the base regression estimates. Again, all the coefficients for the base regression variables are consistent with those in the base regression and statistically significant. For secondary schooling, the civil liberty index, land distribution and financial depth the ranges are (−4.97, −3.46), (1.05, 1.64), (0.14, 0.17) and (−9.32, −6.7), respectively. In short, the coefficient estimates for our four variables are fairly stable and insensitive to various extra regressors. Note that the range for the coefficients of initial income is (−0.81, −0.47) and significant in most of the regressions.

3. Conclusion

To conclude, we briefly link our results to other work on inequality. The first result of this paper – the relative stability of inequality over time – runs counter to one of the most famous conjectures in economics – Kuznets' hypothesis asserting an inverted U-shaped relationship between inequality and income. Although the analysis performed here does not constitute a direct test of the Kuznets hypothesis, the very fact that inequality has been shown to be relatively stable while incomes have almost certainly increased significantly during the 40-year period under study suggests that there is unlikely to be much support in the data for the systematic relationship between inequality and income suggested by Kuznets. Indeed, a more formal test of the Kuznets Hypothesis using the data employed in this paper finds little support for the conjecture (Deininger and Squire, 1996*b*). Rather than treating the inverted-U as a 'stylised fact' of development (Adelman and Robinson, 1989), the data presented here suggest instead that relative stability through time is a much more accurate representation.

Given this, our other result – that inequality differs significantly across countries – assumes importance in the light of recent research suggesting a relationship between initial inequality and subsequent growth (Alesina and Rodrik, 1994, Persson and Tabellini, 1994, Clarke, 1995). If the negative relationship between initial inequality and subsequent growth is indeed supported by the data, then inequitable economies will be condemned to lower growth rates far into the future because inequality is relatively stable. To illustrate, Clarke (1995) notes that a reduction in inequality from one standard deviation above the mean to one standard deviation below the mean would increase the long-term growth rate by approximately 1.3% per annum. This is not an insignificant increase in the growth rate. Even so, our results suggest that on average such a shift in inequality could take almost 70 or even 150 years when evaluated at the average absolute trends based on countries with at least 10 or 20 observations, respectively. Thus, while differences in inequality may explain a part of the difference in growth rates across countries, reducing inequality does not emerge as a simple remedy for increasing growth.

That said, the paper does identify the variables that influence inequality. Our results suggest that inequality is largely determined by factors that change only slowly within countries but are quite different across countries. The two channels identified in the recent literature – the political economy argument and the capital market imperfection channel – received strong support from the empirical analysis with the latter appearing to have the greater influence.

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