INEQUALITY AND INSTITUTIONS

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Abstract—This paper presents theory and evidence on the relationship between inequality and institutional quality. We exhibit a model in which the two may dynamically reinforce each other and set to test this relationship with a broad array of institutional measures. The double causality between institutional strength and a more equal distribution of income is empirically established using dynamic panel and linear feedback analysis.

I. Introduction

WHILE the importance of institutions for economic development has been well documented (see, for example, Acemoglu, Johnson, & Robinson, 2002; Hall & Jones, 1999; Knack & Keefer, 1995; and the more recent Rodrik, Subramanian, & Trebbi, 2002), institutional quality varies significantly across countries. Consider, for example, the most recent report by Transparency International, an organization whose studies on corruption levels are typically published in the popular press around the world. It ranks countries such as Finland, Iceland, Denmark, and New Zealand as those with the lowest levels of corruption, with a score of cleanliness of 9.5 out of 10 points. On the other hand, countries such as Bangladesh, Nigeria, and Haiti are ranked at the highest levels of corruption, with typical scores of less than 1.5 points.[1] Moreover, this ranking tends to be fairly stable across time.[2]

Countries with bad institutions seem also more likely to have high inequality, a pattern that emerges by eyeballing contemporary data. For example, the cross-country data discussed more in detail below clearly shows the close link between the two. The correlation between income share of the middle-income quintile and various measures of institutional quality are in the range of 0.30 and 0.44; and the highest correlation is with the rule-of-law measure. Similarly, the correlation between many measures of institutional quality and the Gini coefficient ranges between 0.40 and 0.44, depending on the aggregate institutional measure employed.

It is by no means clear, however, what the dynamics between these two variables are and, consequently, the resulting causal relationship between them. Some studies indicate that social polarization negatively affects institutional quality (see Easterly, 2001; and Keefer & Knack, 2002), suggesting that institutional strength is endogenous, being determined, among other things, by political and economic conditions. To get some insight about the causal relationship, we calculated correlations between lagged measures of inequality and measures of institutional quality, and then between lagged measures of the latter and inequality, in both cases obtaining significant results with expected signs. Depending on the measures used, the correlations ranged between 0.18 and 0.45.

That the interaction of political and income inequality may play a part in blocking the adoption of good institutions is illustrated by the recent episode of Russia in transition. In the aftermath of the mass privatization in the early 1990s, a small group of entrepreneurs gained access to political power and then used it to promote their own interests, constantly subverting the emergence of institutions committed to the protection of smaller shareholders (see McFaul, 2002). Likewise, in several Latin American countries the interests of ruling elites, the military, and large businesses often converged at the expense of smaller business interests, giving rise to a significant informal sector (see, for example, Kaufmann, Mastruzzi, & Zavaletta, 2003, for a discussion of the Bolivian case). This is also consistent with the recent work by Engerman and Sokoloff (1997, 2002), who contrast the colonial experiences in the Americas, arguing that the initial differences in income inequality between North and South America affected the patterns of settlement and consequently the institutional evolution.

Preliminary evidence on the reinforcing effect on income inequality and institutional quality in a cross-country setup are shown in the figures below. In order to show that this reinforcing property might hold, we use initial values of institutional quality and graph it against income inequality for subsequent years, taking care of possible outliers and controlling for initial inequality. Thus, consider figure 1 that plots Gini coefficients for 1981–1985 against a commonly used International Country Risk Guide (ICRG) institutional index for 1996–2000, conditional on the ICRG index for 1981–1985. The relationship is quite stark.[3] Likewise, figure 2 plots the ICRG index for 1981–1985 against the Gini coefficient for 1996–2000, conditional on the initial income inequality.[4] It clearly shows that high initial quality of institutions is linked with less income inequality in subse-

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[1] As illustrated by the examples above, the higher the score, the less corruption in the country. See http://www.transparency.org/cpi/2003/cpi2003.en.html.

[2] Cf., http://www.transparency.org/cpi for the data covering 1993–2003. Note, however, that country coverage in early years was quite incomplete. Other existing data sources commonly used in empirical studies also show this same relatively stable pattern—see the empirical section, below.

[3] The coefficient of the regression, controlling for the initial institutional quality, is −2.363; the standard error is 0.490; and the corresponding R-squared is 0.74.

[4] The coefficient of the regression, controlling for the initial income inequality, is −0.0109; the standard error is 0.0035; and the corresponding R-squared is 0.78.

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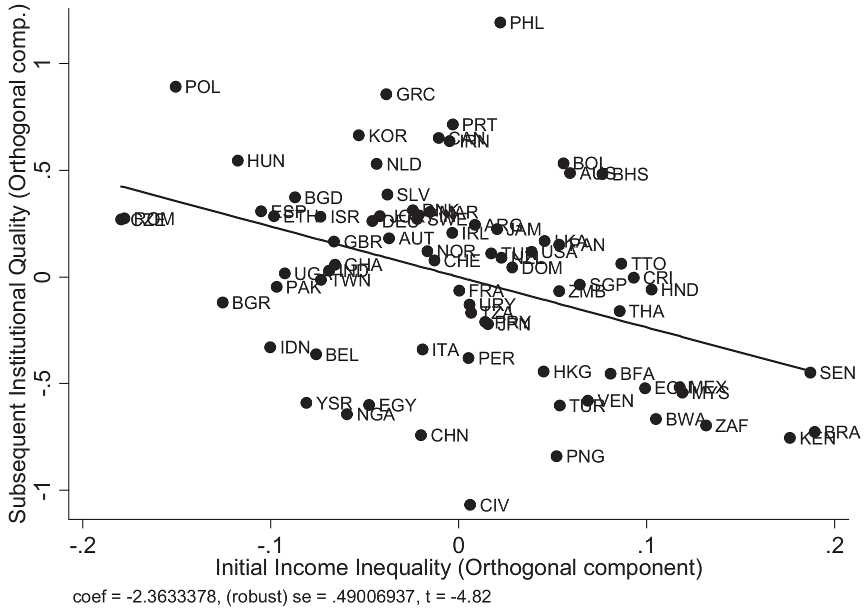
Figure 1.—Initial Income Inequality and Subsequent Institutional Quality, Controlling for Initial Institutional Quality



In this figure, income inequality is captured by the Gini coefficient for the period 1981–1985. Institutional quality is measured by the ICRG index for the period 1996–2000. Regression is conditional on ICRG 1981–1985.

quent years; taken together, these figures suggest that this pattern appears to be reinforcing.[5]

An emerging literature has generated analytical models where economic conditions affect institutional quality; see Hoff and Stiglitz (2002, 2004) and Sonin (2003).[6] In particular, this paper's framework is related to Sonin (2003), which suggests that an equal distribution is a more fertile ground for good institutions. While similar, the mechanism proposed here identifies the intensity of rent-seeking over a public asset—such as technological knowledge or natural resource—as a source of weak institutions. In the context of Russia's development cited above, one illustrative example is provided by the struggle over the control of the gas and oil industries. The particular view of institutional quality as it affects rent-seeking over public assets presented here is also related to the recent literature on the resource curse (see Torvik, 2002; and Mehlum, Moene, & Torvik, 2002).

Specifically, in this paper, we explore a possible double relationship between income inequality and institutional quality. It is suggested that while income inequality may cause subversion of institutions by the politically powerful rich elite, the reverse holds as well, namely, poor institutional quality renders a higher degree of inequality. For example, in our sample of countries discussed more in detail below, countries that experienced the largest increase in inequality in the covered period (China, Iran, Ethiopia, Niger) are also countries that rank at the very bottom of the Gastil index of civil liberties, one of the commonly used measures of institutional strength. This double-causality relationship is exhibited in a simple dynamic model and is then tested in a cross-country panel framework. The model shows that when the political bias in favor of the rich is large, income inequality and poor institutional quality may reinforce each other, indicating a double feedback between the two. Consequently, the economy may converge either to a steady state with high-quality institutions and minimal inequality, or to low institutional quality and high inequality. The empirical evidence provides tentative support for these hypothesized relationships by using a panel of countries and a broad array of institutional measures commonly employed in the literature. We use a GMM-system estimator technique that allows measuring the potential statistical impact of each variable on the other. In particular, we are able to explore the contribution of the two different causal relationships that emerge from the model.[7] One implication of our results concerns the observed persistence in income inequality levels within countries, which goes along with the institutional quality persistence.

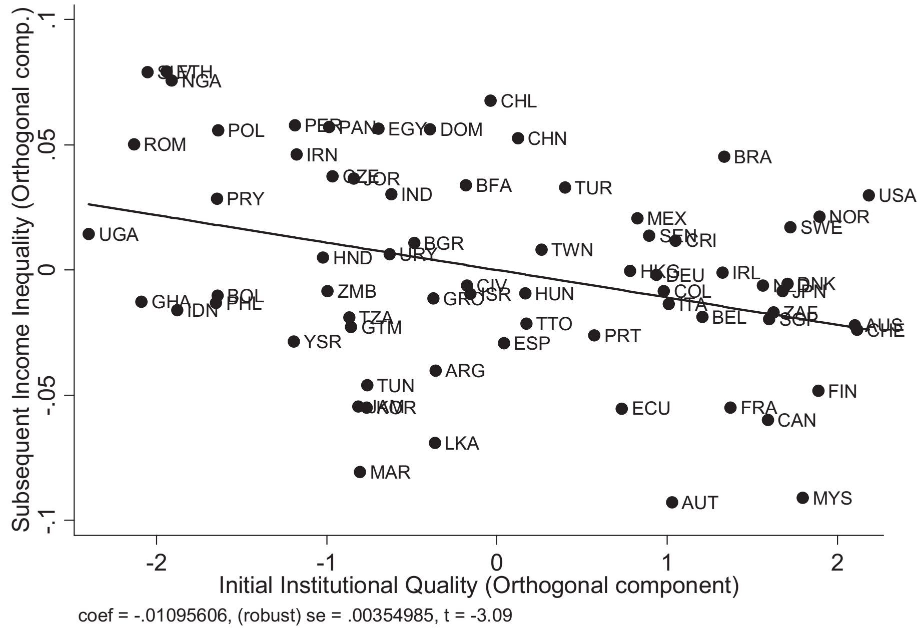
The plan of the paper is as follows. The next section presents and section III solves a simple model, which exhibits double causality between income inequality and institutional quality. Section IV discusses the methodology and the data, and section V presents the main empirical findings. Finally, section VI concludes.

[5] These results are robust with respect to inequality and institutional measures used; details are available upon request.

[6] See also Glaeser, Scheinkman, and Shleifer (2003) for a more micro-based model, and Gradstein (2007), where democracy is viewed as a commitment device to ensure high-quality institutions.

[7] In the working-paper version, Chong and Gradstein (2004), very similar results were obtained using panel VAR estimation and Granger causality tests.

FIGURE 2.—INITIAL INSTITUTIONAL QUALITY AND SUBSEQUENT INCOME INEQUALITY, CONTROLLING FOR INITIAL INCOME INEQUALITY



In this figure, institutional quality is captured by using the ICRG index for the period 1981–1985. Income inequality is measured using the Gini coefficient for the period 1996–2000. Regression is conditional on the Gini coefficient for the period 1981–1985.

II. The Model

Consider an economy populated by a measure one of households indexed by i, each consisting of a parent and child, operating in discrete time t. The initial level of household i's income is exogenously given at y_{io}, and the income level in period t, y_{it}, is determined endogenously. The initial income distribution is assumed to be log normal with the parameters μ_0 and σ_0^2, and the distributions in subsequent periods are endogenously determined. The assumptions below will imply that all future distributions are log normal with the parameters, say, μ_t and σ_t^2. Each individual is also endowed with one unit of time in each period.

In each period, a certain amount of a productive resource is available in the economy. This can be interpreted as a natural resource, or, alternatively, as appropriable technological knowledge. For simplicity, we assume that the amount of the resource is constant over time and let A denote its amount. The individuals allocate their income between consumption, c_{it}, and unproductive investment in rent-seeking, r_{it+1}, to appropriate a larger share of the resource. Normalizing the prices to 1, the budget constraint then is

$$y_{it} = c_{it} + r_{it+1} \tag{1}$$

so that the households are credit constrained. In each period, the individuals also inelastically supply one unit of labor. Rent-seeking is used to appropriate a larger share of the available resource. The extent of the appropriated share by each individual depends on the amount of rent-seeking and on institutional weakness, denoted w_{t+1}. Specifically, the amount appropriated by household i is

$$a_{it+1} = A\,\frac{r_{it+1}^{w_{t+1}}}{\displaystyle\int_0^1 r_{it+1}^{w_{i+1}}\,di}. \tag{2}$$

For simplicity, we will focus on two polar cases, of *strong* institutions ($w_{t+1} = 0$) and *weak* institutions ($w_{t+1} = w$, where w is close to 1): in the former case, the individual marginal value of rent-seeking is 0, whereas in the latter case it is maximal.[8]

Individual income is produced from the share of the appropriated resource and the individual ability; the production function is then given by

$$y_{it} = \varepsilon_{it} a_{it}, \tag{3}$$

where the ability, ε_{it}, is assumed to be distributed in each period log normally, say with the parameters 0 and γ^2, where the variance is assumed to be relatively small.

Each parent's preferences are assumed to derive from consumption, as well as from the amount of income accrued to the child. This simple specification of the "warm glow" altruistic motive implies that the parents need not take into account children's actions when making their own deci-

[8] This is without much loss of generality, as it can be shown by studying the second-order conditions governing the institutional choice that only the extreme values can be optimal—details can be obtained from the authors.

sions. Assuming for simplicity symmetric logarithmic preferences, we write the expected utility:

$$V(c_{it}, y_{it+1}) = \ln(c_{it}) + \ln(y_{it+1}). \tag{4}$$

In each period, all decisions in the economy are made by the parents. They first determine the level of institutional quality and then allocate their resources between consumption, productive investment, and rent-seeking. The determination of institutional quality is done collectively, through a political process, which may generally be biased toward the rich in a manner specified below. The equilibrium consists of such mutually consistent decisions.

III. Equilibrium Analysis

The analysis proceeds backwards. Given the level of institutional quality, households solve their budget allocation problem, and then anticipating these decisions, political choice of institutional quality is made.

A. *Individual Decisions*

Maximization of the utility function (4) subject to the budget constraints (1)–(3) leads to the following individually optimal allocation decisions:

$$r_{it+1} = w_{t+1}y_{it}/(1 + w_{t+1}),\ c_{it+1} = y_{it}/(1 + w_{t+1}) \tag{5}$$

implying that next-period income is

$$y_{it+1} = \varepsilon_{it}Ay_{it}^{w_{t+1}} \bigg/ \int_0^1 y_{it}^{w_{t+1}}\, di. \tag{6}$$

In particular, from equation (5), rent-seeking decreases and current consumption increases with the level of institutional quality.

Taking logarithms, equation (6) can be rewritten as follows:

$$\ln(y_{it+1}) = \ln(\varepsilon_{it}) + \ln(A) + w_{t+1}\ln(y_{it}) - \ln(\mathrm{E}\, y_{it}^{w_{t+1}}) \tag{6'}$$

so that, in particular, the next-period inequality is

$$\sigma_{t+1}^2 = \gamma^2 + w_{t+1}\sigma_t^2. \tag{7}$$

When institutions are strong, $w_{t+1} = 0$, and inequality is constant and is determined by the individual ability differences, $\sigma_{t+1}^2 = \gamma^2$; in contrast, when institutions are weak, $w_{t+1} = w$, and inequality may increase over time, $\sigma_{t+1}^2 - \sigma_t^2 = \gamma^2 + (w - 1)\sigma_t^2 > 0$, especially when current inequality is in the moderate range.

To sum up,

Proposition 1. When institutions are strong, inequality remains constant over time. In contrast, when institutions are weak, an economy may experience an increase in inequality.

B. *Political Determination of Institutional Quality*

We assume that the choice of institutional quality is done via political process, which is biased toward the rich. The simplest way to capture this is to assume that the identity of the decisive voter, y_{dt}, is given by

$$\ln(y_{dt}) = \mu_t + \beta\sigma_t^2, \tag{8}$$

where β represents the extent of political bias in favor of the rich. For example, if $\beta = 0$, the median-income voter is decisive; when $\beta = 1/2$, the average-income voter is decisive; to make the analysis interesting we will assume that the political bias exists and that $\beta > 1/2$.

The individual utility functions corresponding to the two values of institutional quality respectively are

$$U_{it}^{\text{strong}} = \ln(y_{it}) + \ln(\varepsilon_{it}A) \tag{9}$$

and

$$U_{it}^{\text{weak}} = \ln(y_{it}/2) + \ln[\varepsilon_{it}Ay_{it}^{w}/E(y_{it}^{w})] \tag{9'}$$

so that the utility differential is

$$U_{it}^{\text{weak}} - U_{it}^{\text{strong}} = \ln(1/2) + \ln[y_{it}^{w}/E(y_{it}^{w})]. \tag{10}$$

As equation (10) decreases in income, the determination of institutional quality will be done by the decisive voter whose utility differential is

$$\begin{aligned} U_{dt}^{\text{weak}} - U_{dt}^{\text{strong}} &= \ln(1/2) + \ln[y_{dt}^{w}/E(y_{it}^{w})] \\ &= \ln(1/2) + w(\mu_t + \beta\sigma_t^2) - w(\mu_t + \sigma_t^2/2) \\ &= \ln(1/2) + (w\beta - 1/2)\sigma_t^2. \end{aligned} \tag{11}$$

Clearly, when $\beta \leq 1/2$, equation (11) is negative, indicating that a high level of institutional quality will emerge at equilibrium. If, however, the political bias is large as we have assumed, so that the individual with income above average is decisive, $\beta > 1/2$, then it is possible—when income inequality as measured by σ_t^2 is sufficiently large—that the minimal level of institutional quality will be chosen.

To sum up,

Proposition 2. When the political bias is large enough, the political choice of institutional strength hinges upon income inequality. If inequality is small, strong institutions will constitute the political choice; however, when it is large, then weak institutions will prevail.

C. *Intertemporal Evolution*

The analysis of the economy's intertemporal evolution hinges on the initial degree of inequality, σ_0^2. If it is small, then, from equation (11), a high level of institutional quality will be chosen. From equation (7) this then will lead to a constant level of income inequality, which is solely determined by the variance in individual abilities. Because of our assumption that this variance is small, it follows that in future periods strong institutions will also constitute a political choice.

In contrast, if income inequality is initially large, then weak institutions will prevail, $w_{t+1} = w$. Next-period income inequality in this case, is

$$\sigma_{t+1}^2 = \gamma^2 + w\sigma_t^2 \tag{12}$$

and inequality converges to $\sigma^{2*} = \gamma^2/(1 - w^2) > \gamma^2$. Substituting into equation (11), when $\ln(1/2) + (w\beta - 1/2)\,\gamma^2/(1 - w^2) > 0$—which is the case when institutional quality is weak enough, so that w is close to 1—then the political support for weak institutions is guaranteed.[9]

We thus obtain two steady states whose realization depends on initial conditions:[10]

Proposition 3. Initial conditions determine the nature of the steady state: with weak institutions and high inequality, and with high institutional quality and low income inequality.

IV. Empirical Analysis

A. Methodology

The empirical part focuses on the direction of causality between institutions and inequality and their implied contribution to the possible correlation between these variables.[11] The first step is to analyze the dynamic relationship between inequality and institutions. Consistent with the propositions of the model above, the objective is to examine whether changes in a given variable have a lasting impact on another variable.

Our empirical approach to study the direction of the link between institutions and inequality is based on the work by Arellano and Bover (1995), who developed dynamic panel data techniques to address endogeneity problems.[12] This method, the GMM-system estimator, combines in a single system the regression equation in both changes and levels, each with its specific set of instrumental variables. Because of this capability, the model is designed to handle both pooled cross-country and time series data. It is dynamic, since it allows for independent effects from the lagged independent variable. Also, this approach allows for a "weak exogeneity" assumption with respect to the explanatory variables while preserving the estimator's properties (Calderon, Chong, & Loayza, 2002).

The consistency of the GMM estimator depends on whether lagged values of the explanatory variables are valid instruments in the regression. This issue is addressed by considering two specification tests suggested by Arellano and Bond (1991) and Arellano and Bover (1995). The first is a Sargan test of overidentifying restrictions, which tests the overall validity of the instruments by analyzing the sample analog of the moment conditions used in the estimation process. Failure to reject the null hypothesis gives support to the model. The second test examines the hypothesis that the error term is not serially correlated. We test whether the differenced error term (that is, the residual of the regression in differences) is first- or second-order serially correlated. Actually, first-order serial correlation of this error term is expected even if the original error term (in levels) is uncorrelated, unless the latter follows a random walk. Second-order serial correlation of the differenced residual indicates that the original error term is serially correlated and follows a moving-average process at least of order one. If the test fails to reject the null hypothesis of absence of second-order serial correlation, we conclude that the original error term is serially uncorrelated and use the corresponding moment conditions (Calderon et al., 2002). Appendix A presents a detailed description of this approach.

As the GMM-system method does not allow calculating the corresponding contribution of each direction of causality on the total correlation of our variables of interest (Holtz-Eakin, Newey, & Rosen, 1998; Arellano & Bover, 1995), we explore this issue by applying a vector autoregression regressions (VAR) method in a panel setting (Calderon & Liu, 2003).[13] This helps measure the degree of linear dependence and feedback between the two panel series x *(institutions)* and y *(inequality).* We do this by measuring the sum of linear feedback from x *(institutions)* to y *(inequality),* linear feedback from y *(inequality* to x *(institutions),* and "instantaneous" linear feedback between x *(institutions)* and y *(inequality).* The decomposition test is based on likelihood ratios employing the system representation of parameter matrix and the variance-covariance matrix of residuals in order to test a specific set of measures

[9] It is also possible that the economy will oscillate between the two regimes, when w is relatively small.

[10] Note that the steady state here is defined in terms of time invariance of income inequality, whereas average income may well grow over time.

[11] To our knowledge there are only two studies that focus on the link between institutions and income inequality (Chong & Calderon, 2000b; Gupta, Davoodi, & Alonso-Terme, 2002). Both use pure cross-section approaches with relatively small samples.

[12] In addition, unobserved country-specific factors may be correlated with the explanatory variables.

[13] In the previous version of the paper, we provide detailed Granger causality estimates using a VAR panel analysis. As is standard in nonstructural VAR analysis, no cross-equation parameter restrictions are imposed, we allow for a free cross-equation error covariance, and we interpret each equation as a reduced-form regression to test for a dynamic relationship between institutions and inequality. We choose the optimal lag structure for the panel VARs through likelihood ratio tests. This allows us to examine whether a variable, say x *(institutions),* helps forecast the other variable in the system, say y *(inequality),* beyond what the past history of y predicts. Let us denote $z_t = (y_t, x_t)'$ the vector with information on the variables x *(institutions)* and y *(inequality),* and the VAR representation for z_t is $\Gamma_0 z_t = \Gamma_1 L z_t + \xi_t$, with $\Gamma_1 L = \sum_{i=1}^{m} \Gamma_{1i} L^i$. We obtain similar results to the GMM-system shown here (see Chong & Gradstein, 2004).

TABLE 1.—LINEAR FEEDBACK STATISTICS

Linear Feedback	Statistic	Null Hypothesis
From x to y ($F_{x\rightarrow y}$)	$\ln(\lvert\Sigma_{11}^{(1)}\rvert / \lvert\Sigma_{11}^{(2)}\rvert)$	H_0: $F_{x\rightarrow y} = 0$, i.e. That is, $\lvert\Sigma_{11}^{(1)}\rvert = \lvert\Sigma_{11}^{(2)}\rvert$
From y to x ($F_{y\rightarrow x}$)	$\ln(\lvert\Sigma_{22}^{(1)}\rvert / \lvert\Sigma_{22}^{(2)}\rvert)$	H_0: $F_{y\rightarrow x} = 0$, i.e. That is, $\lvert\Sigma_{22}^{(1)}\rvert = \lvert\Sigma_{22}^{(2)}\rvert$
Instantaneous ($F_{x \times y}$)	$\ln(\lvert\Sigma_{11}^{(2)}\rvert / \lvert\Sigma_{11}^{(3)}\rvert) = \ln(\lvert\Sigma_{22}^{(2)}\rvert / \lvert\Sigma_{22}^{(3)}\rvert)$	H_0: $F_{x \times y} = 0$, i.e. "no instantaneous causality between y and x."
Linear Dependence ($F_{x,y}$)	$(F_{x,y}) = F_{x\rightarrow y} + F_{y\rightarrow x} + F_{x \times y}$	H_0: $F_{x,y} = 0$, i.e., "no linear association between y and x."

Sources: Chong and Calderon (2000a), Calderon and Liu (2003).

of linear feedback. The proposed linear feedback statistics to be tested are shown in table 1.

In summary, the basic principle of our empirical approach is to apply a GMM-system estimator and focus on the dynamic relationship in order to test whether there is reinforcement as predicted by the model. Additionally, we apply a panel VAR method in order to decompose the contribution of each direction of causality between institutions and inequality by using a test of linear dependence and feedback.

B. Data Description

We use Gini coefficients as a proxy for income inequality from Deininger and Squire (1996). These data have several advantages. First, the observations are based on household surveys. Second, the population and income coverage are comprehensive. Furthermore, different criteria from different sources are homogenized in order to avoid problems of definition.[14] We augment these data, which cover the period 1960 to 1995, until 2000 by using household data from Milanovic (2002a, 2002b). For the sake of robustness, we also use alternative measures of income distribution such as the income share ratio of the top to the bottom quintile of the population as well as the income shares of the middle quintiles, but do not report results in all instances as they are nearly identical to the Gini findings. While the Gini coefficient ranges from 0 to 1, the income shares for the top and bottom quintiles of the population are ratios that fluctuate between 0 and 1.

Similarly, we use a broad array of governance measures. One source is the International Country Risk Guide (2005), originally used by Knack and Keefer (1995), Hall and Jones (1999), and other authors. The ICRG risk-rating system assigns a numerical value to a predetermined range of risk components for about 130 countries. We consider five of the most commonly used institutional dimensions used in the literature: (i) government stability, (ii) corruption, (iii) law and order, (iv) democratic accountability, and (v) bureaucracy quality; we also compute an ICRG index which is based on the simple average of these five dimensions for the 1960–2000 period. The scores go from 0 to 10; the higher the number the better the quality of the institution. Additionally, we use institutional indices from Freedom House (2005), in particular, an index of civil liberties, an index of political rights, and their simple average, or Gastil index. We utilize series from 1960 to 2000. While their original scores range from 1 to 7, with lower scores denoting higher degrees of freedom, we rescaled these variables from 0 to 1, with higher scores implying more freedom.[15] Finally, we use the aggregate governance indicator developed by Kaufmann, Kraay, and Mastruzzi (2003), which covers 199 countries for 1995, 1998, 2000, and 2002.[16]

In order to avoid potential country selection biases, we homogenize the number of countries to 121, which are the number of common countries in all our data sets, spanning the corresponding full time periods for each sample which are averaged over five years and ten years.[17] In other words, we use panel data of nonoverlapping five- and ten-year-period averages over the full sample period that goes from 1960 to 2000.[18] This is done under the premise that institutional change occurs relatively slowly through time and, thus, the observed variation from year to year may be rather small (Chong & Calderon, 2000a). Table 2 provides summary statistics of the institutional and inequality variables used in this research.

V. Empirical Evidence

A. Simple Correlations

As discussed above, a simple inspection of the data reveals a strong negative relationship between measures of institutional quality and income inequality. As predicted by the model, the pairwise correlations between institutional mea-

[14] Definitional problems include whether a category applies to household or individuals, whether income is measured gross or net of taxes, and whether expenditure or income is used to calculate the income share and Gini coefficient. Atkinson and Brandolini (2001) argue that the income inequality data of Deininger and Squire are far from perfect. In particular, the poor distinction between gross Gini and net Gini is a shortcoming of the data. Still overall, such data are the best we could come up with.

[15] Originally, the ICRG data go from 1985 to 2000 and Freedom House from 1970 to 2000. We take advantage of the very high correlation among institutional series and data from Bollen (1990), which is widely known in the political science literature and was used by Barro (1991) for similar purposes. We simply run bivariate regressions between (i) Bollen and ICRG and (ii) Bollen and Gastil and use the predicted values for the missing years. We do this in order to take advantage of our relatively long income inequality series and to increase the degrees of freedom in our samples, which is very data taxing when using dynamic panel methods. However, none of the findings reported in this paper change when using the unadjusted samples, although in some cases some controls have to be dropped. We would be happy to provide these regressions upon request.

[16] This variable is the average of the six dimensions of institutions. We also tested each variable and find very similar results, but we do not present them here for the sake of economy of space. These findings are available upon request.

[17] This, over the 1960–2000 span. While we also test other year averages, they give very similar results and may be provided upon request. The list of countries is shown in appendix B.

[18] The number of observations is 684 in the case of Freedom House and 430 in the case of ICRG.

TABLE 2.—SUMMARY STATISTICS

	All Countries		Industrial Countries		Developing Countries	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
I. Inequality Measures						
Gini coefficient	0.3909	0.10	0.3220	0.04	0.4121	0.10
Top to bottom	9.2360	6.12	5.9311	1.69	10.3855	6.67
Middle	0.1554	0.04	0.1777	0.02	0.1476	0.04
II. Freedom House						
Gastil index	0.5389	0.32	0.9536	0.10	0.4396	0.27
Civil liberties	0.5346	0.30	0.9372	0.11	0.4382	0.25
Political rights	0.5430	0.35	0.9701	0.10	0.4408	0.31
III. ICRG						
ICRG index	4.0283	1.21	5.6895	0.58	3.5996	0.93
Government stability	7.0606	2.00	8.2530	1.55	6.7529	1.99
Corruption	3.4412	1.33	5.1963	0.82	2.9909	1.03
Rule of law	3.7025	1.50	5.5889	0.68	3.2185	1.26
Democratic accountability	3.6942	1.51	5.6628	0.57	3.1891	1.23
Bureaucratic quality	2.2413	1.18	3.7466	0.50	1.8551	0.98
IV. World Bank						
Aggregate governance	0.2138	0.86	1.5145	0.31	−0.0797	0.65

sures and the Gini coefficient are all negative; and focusing on income shares instead we find analogous results. Further, these correlations are statistically significant at 1% or higher regardless of the measure or data sources employed.[19] Countries with bad institutions seem more likely to have high income inequality as follows from table 3, which exhibits a variety of institutional quality and income inequality measures illustrating significant correlations between them. The results of our model, however, go beyond simple correlations, and imply a mutually reinforcing link between institutions and inequality, whereby initial inequality determines subsequent institutional quality of countries, and institutional quality also determines subsequent income inequality.

A few specific cases may help to further illustrate the link between income inequality and institutional quality. In the case of Peru, for instance, the Gini coefficient was about 0.51 at the beginning of the 1960s. This high income inequality, among the worst in Latin America at the time, was correlated with very low quality of institutions later, reflected in the fact that the corresponding ICRG index for the country was 2.43 at the end of the 1980s. Furthermore, an apparent reinforcing quality between these two variables is illustrated by the fact that this low ICRG index at the end of the 1980s is correlated with a subsequent income inequality of about 0.53 at the middle of the 1990s. Kenya represents another case in point. Whereas the Gini coefficient was about 0.52 at the beginning of the 1970s and the institutional quality of the country was low, as reflected in an ICRG index of 3.5, income inequality worsened during the 1980s to 0.57. Furthermore, while institutional quality remained during the 1980s at about 3.6, income inequality further worsened and reached about 0.61 in the middle of the 1990s. Finally, a third example is that of South Africa. The Gini coefficient in this country was about 0.53 at the beginning of the 1970s and the institutional quality of the country was somewhat low, reflected in an ICRG index of 4.7; the Gini coefficient remained stuck at 0.54 during the 1980s, and the quality of institutions worsened somewhat to 4.4 during the mid-1990s. Figure 3 generalizes these cases to a simple regression of changes in inequality and institutional quality, conditioned on the initial level of the latter.

[19] In fact, we find analogous links using alternative measures of income, such as Theil and Atkinson inequality measures.

B. Causality

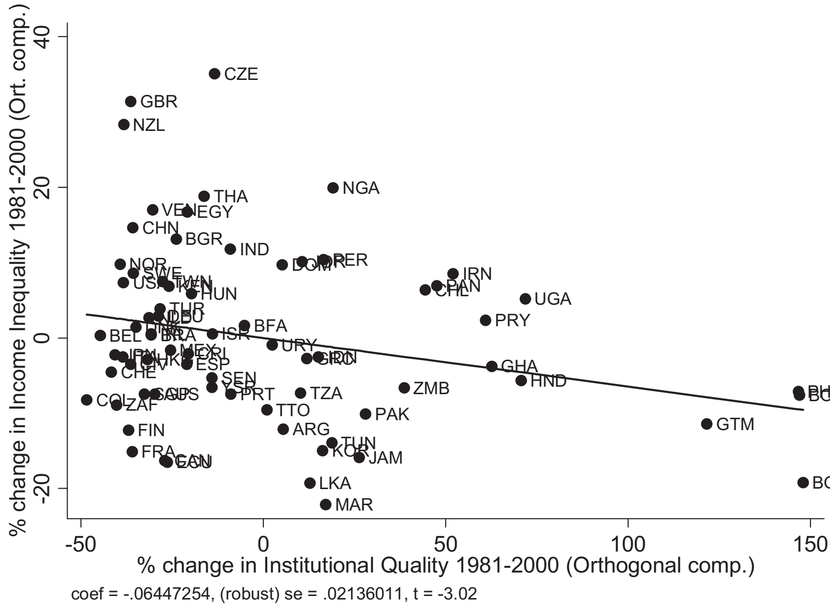
Our main finding is that institutional quality and income inequality reinforce each other, as predicted by our simple theoretical model. This appears to be true regardless of the data set, the specific measure considered, the year grouping, and the econometric methodology, as is shown in table 4, which presents findings using a GMM-system estimator

TABLE 3.—INSTITUTIONS AND INCOME INEQUALITY SIMPLE CORRELATIONS

	Gini Coefficient	Ratio of Top to Bottom Quintiles	Income Share of Middle Quintile
I. Freedom House			
Gastil index	−0.1859	−0.1177	0.1815
	(0.00)	(0.01)	(0.00)
Civil liberties	−0.1892	−0.1238	0.1845
	(0.00)	(0.01)	(0.00)
Political rights	−0.1774	−0.1089	0.1734
	(0.00)	(0.02)	(0.00)
II. ICRG			
ICRG index	−0.4393	−0.3718	0.4225
	(0.00)	(0.00)	(0.00)
Government stability	−0.2769	−0.2172	0.2380
	(0.00)	(0.00)	(0.00)
Corruption	−0.3726	−0.3383	0.3783
	(0.00)	(0.00)	(0.00)
Rule of law	−0.4336	−0.3553	0.4479
	(0.00)	(0.00)	(0.00)
Democratic accountability	−0.3634	−0.3148	0.3501
	(0.00)	(0.00)	(0.00)
Bureaucratic quality	−0.3545	−0.3195	0.3371
	(0.00)	(0.00)	(0.00)
III. World Bank			
Aggregate governance	−0.4020	−0.2986	0.4238
	(0.00)	(0.00)	(0.00)

Statistical significance is shown in parentheses.

Figure 3.—First Differences of Income Inequality and Institutional Quality



In this figure, the change in income inequality is captured by the Gini coefficient and is measured as the percentage change between 1996–2000 and 1981–1985. Change in institutional quality between 1981 and 2000 is measured by the ICRG index. Regression is conditional on ICRG 1981–1985.

approach for panel data of nonoverlapping ten-year periods.[20,21] We control for log of initial output, education, financial development, and the rate of inflation.[22] And we take advantage of the dynamic nature of the method and also include a lagged dependent variable among the explanatory variables. Thus, when the dependent variable is inequality, we include a lagged inequality variable among the explanatory variables, and when the dependent variable is the institutional measure we include a lagged institutional variable among the controls. Also since we specify the regression equation in differences, we are allowed to eliminate the country-specific effects. See appendix A.

Better institutions appear to be conducive to lower income inequality, but lower income inequality may be conducive to better institutional quality, as well. For instance, when using the Freedom House data for the sample of all countries, we observe that a one-unit increase in the Gini coefficient reduces the Gastil index by 0.28; on the other hand, a one-unit increase in the Gastil index reduces the Gini coefficient by 0.06. Consistent with this method, the panel estimates, by construction, exhibit first-order serial correlation, the reason they are not reported. One should be concerned with the presence of second-order serial correlation or higher. In fact, the specification tests applied, in particular the Sargan and second-order serial correlation tests, show that there is no such problem and guarantee the validity of the findings as well as of the instruments used in the estimation process. We also find a pattern of double causality when replicating the exercise above using the ICRG data. Thus, when using data for the sample of all countries, we observe that a one-unit increase in the Gini coefficient reduces the ICRG index by 1.15. On the other hand, we also find that a one-unit increase in the ICRG index reduces the Gini coefficient by 0.02. Again, the Sargan test and the second-order serial correlation tests applied assure the validity of these results as well as of the instruments used in the estimation process. Very similar results that further confirm bidirectionality of institutions and inequality are obtained when using all ICRG submeasures.

Table 4 also shows that the coefficient of persistence in income inequality is approximately 0.83–0.88, depending on the institutional measurement employed. This result is quite consistent with the findings of Bruno, Ravallion, and Squire (1998) who find a simple correlation of 0.85 for income inequality between the 1960s and the 1980s. The idea that past inequality may be an important predictor of current inequality appears to be confirmed. Indeed, the persistence appears to be unconditional to the presence of other elements in the society, as our estimates on this variable do not depend on the presence of additional regressors.[23]

[20] An advantage of this method with respect to VAR panels is that it relies on large N-asymptotics whereas in general, ordinary least squares can only be justified by large T-asymptotics when the time dimension with the available data is relatively limited; yet, the VAR analysis generates very similar results (see Chong & Gradstein, 2004).

[21] When applying GMM system to panel data grouped in five-year periods, two-year periods, and even annual data we obtain statistically significant results for the composite measures (ICRG index, Gastil index, and the World Bank's aggregate governance measure), but we do not always obtain statistically significant results that are robust enough for the individual measures. It can be argued, however, that ten-year-period breakup is more relevant in the present context because institutional quality changes slowly over time.

[22] These variables were included based on the available empirical literature (Chong & Calderon, 2000b, and references therein). The source for all the variables is World Bank (2003). Other empirical specifications—available upon request—do not yield significantly different results.

[23] This is consistent with some previous work; see Li, Squire, and Zou (1998).

TABLE 4.—DYNAMIC RELATIONSHIP BETWEEN INSTITUTIONS AND INEQUALITY

	Gini Coefficient		Top/Bottom 20%		Middle 20%	
Dependent Variable	Inequality	Institutions	Inequality	Institutions	Inequality	Institutions
I. Gastil Index						
Inequality	0.839**	−0.286**	0.719**	−0.199**	0.485**	0.817**
	(0.03)	(0.14)	(0.11)	(0.02)	(0.05)	(0.37)
Institutions	−0.059**	0.549**	−0.394**	0.539**	0.024**	0.538**
	(0.01)	(0.06)	(0.15)	(0.06)	(0.01)	(0.06)
- Sargan test	(0.48)	(0.37)	(0.56)	(0.34)	(0.53)	(0.50)
- 2nd-order correlation	(0.78)	(0.48)	(0.86)	(0.88)	(0.74)	(0.80)
II. Political Rights						
Inequality	0.844**	−0.188**	0.724**	−0.117**	0.491**	1.051*
	(0.03)	(0.06)	(0.11)	(0.02)	(0.05)	(0.58)
Institutions	−0.024**	0.536**	−0.286**	0.526**	0.020**	0.526**
	(0.01)	(0.06)	(0.13)	(0.06)	(0.01)	(0.06)
- Sargan test	(0.51)	(0.40)	(0.58)	(0.70)	(0.31)	(0.31)
- 2nd-order correlation	(0.77)	(0.53)	(0.82)	(0.85)	(0.73)	(0.78)
III. Civil Liberties						
Inequality	0.847**	−0.359**	0.726**	−0.237**	0.474**	0.716*
	(0.03)	(0.13)	(0.11)	(0.02)	(0.05)	(0.42)
Institutions	−0.107**	0.470**	−0.864**	0.445**	0.021**	0.448**
	(0.05)	(0.06)	(0.42)	(0.07)	(0.01)	(0.06)
- Sargan test	(0.38)	(0.27)	(0.38)	(0.29)	(0.30)	(0.32)
- 2nd-order correlation	(0.64)	(0.40)	(0.66)	(0.55)	(0.73)	(0.64)
IV. ICRG Total Index						
Inequality	0.870**	−11.578**	0.703**	−0.031**	0.476**	0.321**
	(0.03)	(4.06)	0.12	(0.01)	(0.05)	(0.11)
Institutions	−0.019**	0.652**	−0.021**	0.696**	0.034**	0.687**
	(0.00)	(0.04)	(0.00)	(0.04)	(0.00)	(0.04)
- Sargan test	(0.36)	(0.36)	(0.36)	(0.34)	(0.59)	(0.50)
- 2nd-order correlation	(0.51)	(0.52)	(0.64)	(0.88)	(0.65)	(0.80)
V. Government Stability						
Inequality	0.857**	−1.463**	0.688**	−0.017**	0.471**	3.907**
	(0.03)	(0.52)	(0.12)	(0.00)	(0.05)	(1.28)
Institutions	−0.026**	0.589**	−0.309*	0.632**	0.003**	0.620**
	(0.00)	(0.06)	(0.18)	(0.06)	(0.00)	(0.06)
- Sargan test	(0.34)	(0.24)	(0.23)	(0.26)	(0.36)	(0.37)
- 2nd-order correlation	(0.43)	(0.61)	(0.58)	(0.92)	(0.62)	(0.87)
VI. Corruption						
Inequality	0.867**	−1.516**	0.695**	−0.016**	0.478**	3.721**
	(0.03)	(0.46)	(0.12)	(0.01)	(0.05)	(1.48)
Institutions	−0.024**	0.714**	−0.319**	0.751**	0.003**	0.750**
	(0.01)	(0.04)	(0.11)	(0.04)	(0.00)	(0.04)
- Sargan test	(0.48)	(0.29)	(0.34)	(0.25)	(0.47)	(0.39)
- 2nd-order correlation	(0.56)	(0.43)	(0.60)	(0.48)	(0.74)	(0.52)
VII. Rule of Law						
Inequality	0.883**	−1.776**	0.699**	−0.019**	0.476**	4.516**
	(0.03)	(0.58)	(0.12)	(0.01)	(0.05)	(1.75)
Institutions	−0.023**	0.704**	−0.169**	0.733**	−0.003**	0.724**
	(0.00)	(0.04)	(0.06)	(0.04)	(0.00)	(0.04)
- Sargan test	(0.35)	(0.25)	(0.31)	(0.26)	(0.23)	(0.34)
- 2nd-order correlation	(0.55)	(0.51)	(0.49)	(0.49)	(0.39)	(0.53)
VIII. Democratic Accountability						
Inequality	0.866**	−1.453**	0.701**	−0.014**	0.480**	4.079**
	(0.03)	(0.55)	(0.12)	(0.01)	(0.05)	(1.90)
Institutions	−0.032**	0.616**	−0.296**	0.654**	0.003**	0.651**
	(0.01)	(0.04)	(0.07)	(0.05)	(0.00)	(0.05)
- Sargan test	(0.24)	(0.35)	(0.28)	(0.29)	(0.52)	(0.24)
- 2nd-order correlation	(0.53)	(0.78)	(0.63)	(0.80)	(0.70)	(0.85)
IX. Bureaucratic Quality						
Inequality	0.864**	−1.670**	0.694**	−0.017**	0.475**	3.604**
	(0.03)	(0.36)	(0.12)	(0.01)	(0.05)	(1.21)
Institutions	−0.036**	0.706**	−0.270**	0.728**	0.003**	0.733**
	(0.01)	(0.04)	(0.04)	(0.04)	(0.00)	(0.04)
- Sargan test	(0.46)	(0.42)	(0.29)	(0.26)	(0.36)	(0.36)
- 2nd-order correlation	(0.54)	(0.54)	(0.49)	(0.50)	(0.79)	(0.74)

Tests of dynamic relationship where x represents the corresponding institutional measure and y represents the inequality measure as measured by the Gini coefficient. All regressions include fixed effects. Standard errors are shown in parentheses.

* denotes significance at the 10% level.

** denotes significance at the 5% level.

TABLE 5.—LINEAR FEEDBACK BETWEEN INSTITUTIONS AND INEQUALITY

	Granger Panel			
	$x \rightarrow y$	$y \rightarrow x$	$y \times x$	y, x
I. Freedom House				
Gastil index	36.5	59.0	4.5	100.0
	(0.06)	(0.02)	(0.96)	(0.03)
Civil liberties	25.5	66.0	8.4	100.0
	(0.04)	(0.03)	(0.61)	(0.03)
Political rights	24.5	73.7	1.8	100.0
	(0.09)	(0.00)	(0.84)	(0.02)
II. ICRG				
ICRG index	33.2	63.5	3.4	100.0
	(0.04)	(0.00)	(0.67)	(0.01)
Government stability	35.0	60.7	4.3	100.0
	(0.03)	(0.00)	(0.53)	(0.01)
Corruption	32.2	66.2	1.7	100.0
	(0.02)	(0.00)	(0.78)	(0.00)
Rule of law	31.5	65.3	3.2	100.0
	(0.01)	(0.00)	(0.71)	(0.00)
Democratic accountability	32.9	64.5	2.5	100.0
	(0.01)	(0.00)	(0.52)	(0.01)
Bureaucratic quality	36.5	60.6	2.9	100.0
	(0.02)	(0.00)	(0.84)	(0.00)
III. World Bank				
Aggregate governance	33.4	55.0	11.5	100.0
	(0.05)	(0.02)	(0.91)	(0.02)

The variable x represents the measure of institutional quality, whereas the variable y represents the measure of income inequality, as measured by the Gini coefficient. All feedback measures are expressed as a percentage of the total correlation or linear dependence between institutions and inequality ($F\,x, y$). Hence, the causality from *institutions* to *inequality* is represented by $x \rightarrow y$. Similarly, the causality from *inequality* to *institutions* is represented by $y \rightarrow x$. Instantaneous causality is represented by $y \times x$, which is statistically insignificant in all cases. The statistical significance of each feedback measure is shown in parentheses (p-values for χ^2 tests).

C. Feedback

The previous section has focused on the signs of the coefficients and their statistical significance in order to assess whether a double causality between institutions and inequality exists. In this section, we quantify the extent of the contribution of each direction of causality possible between these variables in the observed overall correlation. As explained above, the assumptions of the GMM system (Arellano & Bover, 1995; Blundell & Bond, 1998) do not allow decomposition of the contribution of each direction of causality to overall correlation. We explore this question by applying a panel VAR methodology and decompositions tests to help measure the degree of linear dependence and feedback between our variables of interest and by measuring the sum of linear feedback from *institutions* to *inequality* and vice versa, as shown in table 1.[24]

In table 5 we show that when using the Gastil index data and the Gini coefficient as a measure for inequality, we find that the contribution of the *institutions to inequality* causality to the total linear dependence between these two variables in the entire sample is approximately 37%, and the contribution of the *inequality to institutions* causality to the total linear dependence between these two variables is approximately 59%.[25] We obtain very similar results when using ICRG and World Bank data, as well. For instance, in the case of the Gini coefficient, and using the full sample, we find that the contribution of the *institutions to inequality* causality to the total linear dependence between these two variables is around 33%, and the contribution of the *inequality to institutions* causality to the total linear dependence between these two variables is approximately 64%. Very similar results are obtained when using the World Bank's aggregate governance measure (33% and 55%, respectively). In general, we find strong evidence of bidirectional causality with the institutions to inequality direction being always statistically significant but having a smaller share in the total linear dependence relationship.[26] In fact, the causal direction from income inequality to institutional quality dominates the linear relationship between these variables regardless of the institutional indicators, the sample of countries, and the income distribution variable used.[27]

VI. Conclusions

The starting point of this paper is the observation that there is a significant correlation between income inequality and weakness of institutions. In theory, it stands to reason that weak institutions may be conducive to income inequality. Where the poor are not given the protection by an independent judicial system, for example, their ability to extract rents is inferior to that of the rich. It has also been suggested that high income inequality allows the rich to wield stronger political influence, thereby subverting institutions.

This double-causality relationship is first exhibited in a simple formal model here and then tested empirically employing a comprehensive cross-country panel data set. The adopted approach enables us to directly establish causality links using a dynamic panel GMM-system methodology (Arellano & Bover, 1995; Blundell & Bond, 1998). Our findings indicate that, consistent with the theory, institutions cause inequality as well as inequality causes institutions, which provides strong

[24] Our methodology consists of estimating and testing vector autoregressions (VAR) in a panel setting that have the following form:

$$y_{i,t} = A(L)y_{i,t} + B(L)x_{i,t} + \eta_t + \mu_i + \varepsilon_{i,t} \tag{i}$$

$$x_{i,t} = C(L)y_{i,t} + D(L)x_{i,t} + \phi_t + \psi_i + \upsilon_{i,t} \tag{ii}$$

where y and x represent the two variables of interest, inequality and institutions; L is the lag operator; A, B, C, and D are vectors of coefficients; η_t and ϕ_t are unobserved time effects; μ_i and ψ_i are unobserved country effects; and $\varepsilon_{i,t}$ and $\nu_{i,t}$ are regression residuals. Note that we also control for other determinants, Z, in particular the log of output, education, financial development, and the rate of inflation. The subscripts i and t denote country and time, respectively.

[25] This finding also holds in the subsamples of countries. For example, in the case of the Gastil index the contribution of the institutions to inequality causality to the total linear dependence for the subsample of industrial countries is approximately 25% and the contribution of the inequality to institutions causality to the total linear dependence is approximately 64%; for developing countries, the contribution of the institutions to inequality causality to the total linear dependence is 33% and the contribution of the inequality to institutions causality to the total linear dependence is 58%. These findings are reported in further detail in Chong and Gradstein (2004).

[26] The instantaneous causality between these two variables, as defined in table 1, is never statistically significant and is not reported.

[27] Furthermore, robustness checks regarding the other measures of income inequality (that is, income share ratio of top to bottom quintiles, share of the middle-income quintile, Atkinson and Theil indexes) generate very similar results.

empirical support to the mutually reinforcing mechanism between these variables. Unlike typical causality studies, we use a simple VAR panel methodology in order to decompose the contribution of each type of causality onto the observed total linear dependence between variables. In fact, the direction of causality from inequality to institutions appears to dominate the reverse causality. These findings hold for various institutional measures, as well as for different year groupings, sample sizes, inequality measures, and changes in specification. While our findings do not dispute the premise that better institutions may lead to a more equal distribution of income, the established reverse causality may help explain why countries with full awareness of the need to pursue dramatic institutional reforms have failed to do so. Institutional reform may be an instrument to reduce inequality; political factors, however, may prevent its implementation.

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APPENDIX A

GMM-System Estimator[28]

This methodology formulates a set of moment conditions that can be estimated using GMM techniques in order to generate consistent and efficient estimates. Specifying the regression equation in differences allows elimination of the country-specific effect. First-differencing yields

$$y_{i,t} - y_{i,t-1} = \beta_1(y_{i,t-1} - y_{i,t-2}) + \beta_2(X_{i,t} - X_{i,t-1}) + (\varepsilon_{i,t} - \varepsilon_{i,t-1}). \quad \text{(A1)}$$

The use of instruments is required to deal with two issues: first, the likely endogeneity of the explanatory variables, X, which is reflected in the correlation between these variables and the error term; and second, the correlation of the new error term, $(\varepsilon_{i,t} - \varepsilon_{i,t-1})$, by construction with the differenced lagged dependent variable, $(y_{i,t-1} - y_{i,t-2})$. We adopt a flexible assumption of weak exogeneity, according to which current explanatory variables may be affected by past and current realizations of the dependent variable but not by its future innovations. Under the assumptions that the error term, ε, is not serially correlated and the explanatory variables are weakly exogenous, the following moment conditions apply:

$$E[y_{i,t-s} \times (\varepsilon_{i,t} - \varepsilon_{i,t-1})] = 0 \quad \text{for } s \geq 2;\ t = 3, \ldots, T \quad \text{(A2)}$$

$$E[X_{i,t-s} \times (\varepsilon_{i,t} - \varepsilon_{i,t-1})] = 0 \quad \text{for } s \geq 2;\ t = 3, \ldots, T \quad \text{(A3)}$$

The GMM estimator simply based on the moment conditions in equations (A2) and (A3) is known as the differences estimator. Although asymptotically consistent, this estimator has low asymptotic precision and large biases in small samples, which leads to the need to complement it with the regression equation in levels.[29] For the regression in levels, the country-specific effect is not directly eliminated but must be controlled for by the use of instrumental variables. The appropriate instruments for the regression in levels are the lagged differences of the corresponding variables if the following assumption holds; although there may be correlation between the levels of the right-side variables and the country-specific effect, there is no correlation between the differences of these variables and the country-specific effect. This assumption results from the following stationarity property:

$$E\lfloor y_{i,t+p} \times \eta_i \rfloor = E\lfloor y_{i,t+q} \times \eta_i \rfloor \text{ and } E\lfloor X_{i,t+p} \times \eta_i \rfloor = E\lfloor X_{i,t+q} \times \eta_i \rfloor \quad \text{for all } p \text{ and } q. \quad \text{(A4)}$$

Therefore, the additional moment conditions for the second part of the system (the regression in levels) are given by the following equations:

$$E[(y_{i,t-s} - y_{i,t-s-1}) \times (\eta_i + \varepsilon_{i,t})] = 0 \quad \text{for } s = 1 \quad \text{(A5)}$$

$$E[(X_{i,t-s} - X_{i,t-s-1}) \times (\eta_i + \varepsilon_{i,t})] = 0 \quad \text{for } s = 1 \quad \text{(A6)}$$

Using the moment conditions presented in equations (A2), (A3), (A4), and (A5), and following Arellano and Bond (1991) and Arellano and Bover (1995), we employ a generalized method of moments (GMM) procedure to generate consistent estimates of the parameters of interest. The weighting matrix for GMM estimation can be any symmetric, positive-definite matrix, and we obtain the most efficient GMM estimator if we use the weighting matrix corresponding to the variance-covariance of the moment conditions. Since this variance-covariance is unknown, Arellano and Bond (1991) and Arellano and Bover (1995) suggest a two-step procedure. First, assume that the residuals, $\varepsilon_{i,t}$, are independent and homoskedastic both across countries and over time. This assumption corresponds to a specific weighting matrix that is used to produce first-step coefficient estimates. We construct a consistent estimate of the variance-covariance matrix of the moment conditions with the residuals obtained in the first step, and we use this matrix to reestimate our parameters of interest (that is, second-step estimates). Asymptotically, the second-step estimates are superior to the first-step ones insofar as efficiency is concerned. The moment conditions are applied such that each of them corresponds to all available periods, as opposed to each moment condition corresponding to a particular time period. In the former case, the number of moment conditions is independent of the number of time periods, whereas in the latter case, it increases more than proportionally with the number of time periods. Most of the literature dealing with GMM estimators applied to dynamic models of panel data treats the moment conditions as applying to a particular time period. This approach is advocated on the grounds that it allows for a more flexible variance-covariance structure of the moment conditions. Such flexibility is achieved without placing a serious limitation on the degrees of freedom required for estimation of the variance-covariance matrix because the panels commonly used in the literature have both a large number of cross-sectional units and a small number of time series periods.

[28] This section draws heavily from Calderon et al. (2002).

[29] Alonso-Borrego and Arellano (1999) and Blundell and Bond (1998) show that when the lagged dependent and the explanatory variables are persistent over time, lagged levels of these variables are weak instruments for the regression equation in differences. This weakness has repercussions for both the asymptotic and small-sample performance of the differences estimator. As persistence increases, the asymptotic variance of the coefficients obtained with the differences estimator rises.

APPENDIX B.—LIST OF COUNTRIES

1	ARE	United Arab Emirates	61	LBY	Libya
2	ARG	Argentina	62	LKA	Sri Lanka
3	AUS	Australia	63	LSO	Lesotho
4	AUT	Austria	64	LTU	Lithuania
5	BEL	Belgium	65	LUX	Luxembourg
6	BFA	Burkina Faso	66	LVA	Latvia
7	BGD	Bangladesh	67	MAR	Morocco
8	BGR	Bulgaria	68	MDG	Madagascar
9	BHR	Bahrain	69	MEX	Mexico
10	BHS	Bahamas	70	MLI	Mali
11	BLR	Belarus	71	MLT	Malta
12	BOL	Bolivia	72	MNG	Mongolia
13	BRA	Brazil	73	MRT	Mauritania
14	BWA	Botswana	74	MUS	Mauritius
15	CAN	Canada	75	MYS	Malaysia
16	CHE	Switzerland	76	NER	Niger
17	CHL	Chile	77	NGA	Nigeria
18	CHN	China	78	NIC	Nicaragua
19	CIV	Côte d'Ivoire	79	NLD	Netherlands
20	COL	Colombia	80	NOR	Norway
21	CRI	Costa Rica	81	NPL	Nepal
22	CYP	Cyprus	82	NZL	New Zealand
23	CZE	Czech Republic	83	OMN	Oman
24	DEU	Germany	84	PAK	Pakistan
25	DNK	Denmark	85	PAN	Panama
26	DOM	Dominican Republic	86	PER	Peru
27	DZA	Algeria	87	PHL	Philippines
28	ECU	Ecuador	88	PNG	Papua New Guinea
29	EGY	Egypt	89	POL	Poland
30	ESP	Spain	90	PRT	Portugal
31	EST	Estonia	91	PRY	Paraguay
32	ETH	Ethiopia	92	QAT	Qatar
33	FIN	Finland	93	ROM	Romania
34	FRA	France	94	RUS	Russia
35	GBR	United Kingdom	95	RWA	Rwanda
36	GHA	Ghana	96	SAU	Saudi Arabia
37	GIN	Guinea	97	SEN	Senegal
38	GNB	Guinea-Bissau	98	SGP	Singapore
39	GRC	Greece	99	SLE	Sierra Leone
40	GTM	Guatemala	100	SLV	El Salvador
41	HKG	Hong Kong	101	SVK	Slovak Rep.
42	HND	Honduras	102	SVN	Slovenia
43	HRV	Croatia	103	SWE	Sweden
44	HUN	Hungary	104	SYR	Syria
45	IDN	Indonesia	105	THA	Thailand
46	IND	India	106	TTO	Trinidad and Tobago
47	IRL	Ireland	107	TUN	Tunisia
48	IRN	Iran	108	TUR	Turkey
49	IRQ	Iraq	109	TWN	Taiwan
50	ISR	Israel	110	TZA	Tanzania
51	ITA	Italy	111	UGA	Uganda
52	JAM	Jamaica	112	UKR	Ukraine
53	JOR	Jordan	113	URY	Uruguay
54	JPN	Japan	114	USA	United States
55	KAZ	Kazakhstan	115	VEN	Venezuela
56	KEN	Kenya	116	VNM	Vietnam
57	KGZ	Kyrgyz Rep.	117	YEM	Yemen
58	KOR	Korea, Rep.	118	YSR	Yugoslavia
59	KWT	Kuwait	119	ZAF	South Africa
60	LBN	Lebanon	120	ZMB	Zambia
			121	ZWE	Zimbabwe