Bank Regulations and Income Inequality: Empirical Evidence

Manthos D. Delis
Surrey Business School
University of Surrey

Iftekhar Hasan
Fordham University and Bank of Finland

Pantelis Kazakis
Department of Economics
Ohio State University

Abstract

This paper provides cross-country evidence that variations in bank regulatory policies result in differences in income distribution. In particular, the overall liberalization of banking systems decreases income inequality significantly. However, this effect becomes insignificant for countries with low levels of economic and institutional development and for market-based economies. Among liberalization policies, credit and interest rate controls have the most significant negative effect on inequality. Privatizations and liberalization of international capital flows also decrease income inequality; the latter also increases the income share of the relatively poor. In contrast, liberalization of securities markets increases income inequality substantially. These findings show that bank regulations have a significant effect on real economic outcomes as well as financial stability.

Keywords: Bank regulations; Income inequality; Cross-country panel data; Instrumental variables

JEL classification: G28; O15; O16

1. Introduction

The financial crisis of 2007 reenergized the historic debate regarding the victims of economic downturns and the link between recessions and income inequality. Most critics argue that the free markets unambiguously failed to safeguard themselves and the economy, and that this failure might have had a larger marginal effect on the relatively poor. However, regulatory policies that aim to provide a safety net for lenders, borrowers, and depositors also failed to prevent this crisis from developing and spreading worldwide.

Though it may be too soon to determine who the primary victims of this crisis are, banking liberalization and/or reregulation can play an important role in shaping income inequality in different countries. Furthermore, the diversity in regulatory practices among countries exists despite the Basel Committee's recent initiatives to harmonize and benchmark regulatory frameworks. Thus, a study that assesses the impact of cross-country and timely variations in bank regulatory policies on income inequality is an interesting endeavor. This paper is, to our knowledge, the first to do that.

The extensive literature on the relationship between finance and the distribution of income generally agrees that improvements in financial markets, contracts, and intermediaries reduce income inequality because financial development affects the degree to which an individual’s future income is the result of talent and good investment ideas or inherited income (Demirgüç-Kunt and Levine, 2009). If financial intermediaries succeed in funding talented people with good investment ideas through advanced screening methods, the theory goes, an individual’s economic endowment should not dictate his or her future income. In contrast, in underdeveloped or constrained financial systems, individuals might face severely limited access to capital. That is, financial imperfections, such as information inconsistencies and transaction costs, as well as high levels of relationship lending, may be especially binding on the poor, who lack collateral and credit history (Beck, Demirgüç-Kunt and Levine, 2007).

Further, failing to liberalize the banking sector leads to local monopolies, a situation that hurts the poor because it creates inefficiently high lending rates and encourages relationship lending (Galor and Moav, 2004).

Important early contributors to research in this area, such as Greenwood and Jovanovic (1990), contradict the idea of a negative relationship between financial liberalization and inequality. They suggest that banks with profit-maximizing behavior lend to richer firms and households and avoid lending to individuals with low levels of collateral. Inherent in this is the assumption that poor individuals are riskier and that lending to them might conflict with banks' incentive to earn maximum yields on their risky assets.

This exploration of the finance-inequality relationship does not explicitly account for the dynamic nature of regulatory policies related to the banking sector. To put it another way, the literature does not address the specific features of banking regulations in different countries and their evolution as a source of income inequality. This is an important limitation in many respects. First, the initial wave of liberalization policies in the 1980s gave way to reregulation of the banking sector in the 1990s and 2000s in an effort to prevent systematic instability and crises. Furthermore, because different types of regulatory policies can have widely different objectives, the effect of these policies on the distribution of income might also be different and/or contradictory. Finally, regulatory policies play a distinct role in lending behavior, and this role may be different from the one that shapes financial development.

Notably, only Beck, Levine, and Levkov (2010) consider the impact of bank regulation on the distribution of income. In particular, they assess how liberalizing intrastate branching affects the distribution of income within the United States and find that deregulation significantly reduces inequality by boosting incomes in the lower part of the income distribution. They also find that deregulation has little impact on incomes above the median. This is first-hand evidence that bank regulatory policies may have a central role in shaping the distribution of income.

Based on these considerations, we focus on how banking regulations affect the distribution of income in different countries. We consider reforms in seven pillars of banking regulation: credit controls and reserve requirements, interest rate controls, banking-sector entry, capital-account transactions, bank privatizations, liberalization of securities markets, and banking-sector supervision and capital regulation. We examine the impact of these indices separately and jointly on income inequality (measured by the Gini coefficient and the Theil index) and the distribution of income (measured by the lower and higher 10% of the income distribution and by poverty rates) for 87 countries. Our identification strategy accounts for the endogeneity of banking regulations in empirical models of income inequality.

The empirical findings show that banking deregulation (higher liberalization) generally leads to lower inequality and narrower income distribution. Specifically, countries in the upper quartile of the banking liberalization index will, *ceteris paribus,* have 2.5 times less inequality (as measured by the Gini coefficient) compared to countries in the lower quartile of the banking liberalization index. This effect is highly statistically and economically significant.

We also find that abolishing credit and interest rate controls decreases income inequality substantially and that more effective banking supervision has a similar effect. Credit and interest-rate controls also narrow the income distribution by increasing the income share held by the poor. Specifically, the income share held by the poorest 10% of the income distribution in countries in the upper 25th percentile in the banking liberalization index will, *ceteris paribus*, be 25% higher compared to countries in the lower quartile of the banking liberalization index.

We also identify heterogeneous effects of liberalization policies according to the level of institutional and economic development. In particular, we find that abolishing barriers to entry and enhancing privatization laws lower income inequality primarily in developed countries with stronger institutions. Hence, we conclude that economic and institutional development is a prerequisite for regulations to have a positive effect on the real economy (Laffont, 2005). Moreover, the liberalization of international capital flows increases the income share of both the rich and the poor. Finally, and in contrast to previous liberalization policies, we find that liberalizing securities markets increases income inequality. We find that banks pass on their increased costs and higher capital requirements to the relatively lower-income population that lacks good credit and collateral.

The rest of this paper is structured as follows: Section 2 describes the data set and evaluates the impact of specific types of bank regulation on income inequality. Section 3 discusses the identification issues and the econometric methodology. Section 4 presents the empirical results. Section 5 offers some policy implications and concludes the paper.

2. Data Description

To examine the impact of different forms of regulation on income inequality in an international setting, we collect country-level data. The original sample includes data from 91 countries for which information on bank regulations is available over the period 1973-2005. However, the final sample includes 87 countries for which data on all our main variables is available. One of our two identification strategies requires information that has only been available since 1997. Therefore, our empirical analysis uses data for the period 1997-2005. Given that annual macroeconomic data are noisy (Roine, Vlachos, and Waldenström, 2009), we primarily use three-year averages but also conduct a sensitivity analysis with annual and cross-sectional data.[1]

Table I provides a collective, formal definition of the variables in the empirical analysis, and Table II offers summary statistics. We also include a number of tables in the internet appendix: Table AI provides information on the number of countries in our sample, their regional group, and the extent of liberalization of their banking systems; Table AII offers correlation coefficients among the explanatory variables.

[INSERT TABLES I AND II]

2.1 INCOME INEQUALITY

Our main variable to proxy inequality is the Gini coefficient from the Standardized World Income Inequality Database (SWIID) of Solt (2009). Alternatively, we use the Theil index from the University of Texas Inequality Project, which relies on data from Deininger and Squire (1996) and the United Nations Industrial Development Organization (UNIDO). The SWIID database is the most comprehensive database on the Gini coefficient and increases greatly the comparability of the values between countries, primarily because it standardizes consumption and wage income (see Solt, 2009). In the SWIID dataset, data on the Gini coefficient are available for 153 countries over the period 1960-2010. The Theil index is available for 156 countries over the period 1963-2002. Both panels are unbalanced, with the Theil index having many missing observations.

The Gini coefficient is derived from the Lorenz curve and ranges between 0 and 100. A low Gini indicates a more equal distribution, with 0 corresponding to perfect equality; a

[1] As in virtually all of the empirical studies of income inequality that rely on Gini-coefficient data (e.g., Chong and Gradstein, 2007), the dataset is unbalanced in the sense that data for some time periods are missing for certain countries. Following previous studies, we construct averages for the dependent variables over specific time intervals by using the available observations. This procedure also smooths out some abrupt jumps in the year-on-year values for some countries. Sensitivity analysis on the main results using country averages and running cross-sectional regressions confirms that the results are unaffected by exclusion of countries where such abnormal changes occur.

higher Gini indicates more unequal distribution, with 100 corresponding to perfect inequality. The Gini coefficient is the most widely used measure of inequality in the empirical literature (e.g., Beck, Demirgüç-Kunt and Levine, 2007; Dollar and Kraay, 2002).

We use the index that represents household income before taxes, as this shows inequality exclusive of fiscal policy. Solt (2009) extensively discusses how this index is constructed on the basis of previous datasets on economic inequality. The average value in our dataset for the period 1997-2005 is 44.68 and is 40.69 for the full period. Countries such as Armenia, Bahamas, Kuwait, Mongolia, and Qatar have very high values, and the Czech Republic, China, Macao, Slovenia, and Sweden exhibit low Gini coefficients.[2]

Theil's index of inequality is an entropy measure. Maximum entropy occurs once income earners are indistinguishable by resources (i.e., when there is perfect equality). In real societies, different resources (i.e., incomes) distinguish people. The more "distinguishable" they are, the lower the "actual entropy" of a system consisting of income and income earners. Thus, higher values on the Theil index reflect higher inequality.[3] In our dataset and over the period 1997-2005, the Theil index has a correlation coefficient of 0.73 with the Gini coefficient and obtains an average value of 0.07. Countries with high and low values are about the same as those reported for the Gini coefficient.

The two measures are not perfectly comparable because the Gini coefficient in this study is based on household income, but the Theil index is based on individual wage income. This remains partially true, despite the considerable work on this front by Solt (2009). The

[2] In deriving the single Gini coefficient employed here, Solt (2009) uses information from 21 different Gini coefficients constructed in previous studies. From this information, he estimates the SWIID Gini coefficients by assigning a standard error on the estimates. The largest standard errors, and thus the less reliable Gini coefficients, are in the early years of the SWIID dataset (1980 or before). Because our sample starts in 1997, this study does not cover these years. Further, the relatively high standard errors within our sample are usually related to large and abrupt changes in the Gini coefficients of SWIID. In our final sample there are 13 countries (both developing and developed) with changes higher than five points in the Gini from one three-year interval to the next, but there are no countries with changes higher than 10 points in the Gini. We provide sensitivity analysis for our main results by using multiple Monte Carlo simulations in the fashion of King, Honaker, Joseph and Scheve (2001). For further discussion, see Solt (2009).

[3] For a thorough description of the Theil index, see http://utip.gov.utexas.edu.

disadvantage of the specific Theil index is that a wage-based measure excludes nonwage income, such as pensions or income from self-employment (Deininger and Squire, 1996).

The Theil index does have one advantage over the Gini coefficient, however: it provides information on individuals, not households. The distinction is important if there are systematic differences in the size of rich and poor households. Deininger and Squire (1996), for example, show that the Gini coefficient based on income across households provides estimates of inequality slightly lower than the equivalent Gini based on income across individuals. Even though these considerations imply that we should compare the results from the two measures with some caution, they also suggest that the two measures in this study are complementary.

To answer whether banking sector liberalization disproportionately affects the poor, we employ data from the World Bank's World Development Indicators (WDI) database. Specifically, we employ three more dependent variables by using data on the share of income in the highest and lowest 10% of the income distribution, respectively, as well as data on the poverty gap at $1.25 a day. The first two variables provide more detailed information on the distribution of income, and the third is a direct proxy for poverty. We should state, however, that these indicators have a number of shortcomings.

First, advanced economies only have one or two years of this data, and thus we have to exclude these countries from our sample. This introduces selectivity bias in the results and reduces the number of available observations. For these variables, our sample includes information from 56 countries. Second, unlike our Gini index, the measures of inequality from the WDI are constructed on the basis of either income or consumption and are not standardized. Therefore, the values for these variables are not directly comparable across countries (for details, see Solt, 2009). Finally, some countries report different values for urban and rural populations. For them we have to take averages, which essentially introduce some measurement error. Yet, the findings still provide insights on what drives the relationship between banking-sector liberalization and income distribution.

2.2 BANK REGULATIONS AND THEIR IMPACT ON INCOME INEQUALITY

Abiad, Detragiache, and Tressel (2010) describe in detail the regulatory conditions that characterize the banking industries in 91 countries over the period 1973-2005. They offer seven indices of financial-sector policy that compose a single general indicator of financial reform. For each index, a country receives a score on a graded scale, with zero corresponding to “fully repressed,” one to “partially repressed,” two to “largely liberalized,” and three to “fully liberalized.” Here we drop four countries for which we have no information on other important variables in our empirical analysis.

Table AI in the internet Appendix groups the 87 countries in our final sample by region. Also, this table shows the extent to which countries across these regions liberalized their banking systems over the period 1997-2005. Most countries (55 in total) reformed their banking systems; 29 countries maintained the status quo. Of the 29, eight had a fully liberalized banking sector by 2005. Finally, note that the different types of liberalization policies cluster across countries. That is, countries that liberalize one aspect of their banking sector tend to liberalize other aspects simultaneously or quickly after. Thus, the seven subindices tend to move together on a country-specific basis.

The first of the seven subindices relates to credit controls, such as credit directed toward favored sectors or industries, ceilings on credit in other sectors, and excessively high reserve requirements. The second index documents interest rate controls, including whether the government directly controls interest rates, or whether floors, ceilings, or interest-rate bands exist. The third index considers entry barriers, such as licensing requirements, limits on the participation of foreign banks, and activity restrictions relating to bank specialization or establishing universal banks. The fourth index considers the prudential regulation and supervision of the banking sector. It encompasses regulatory policies pertaining to capital regulation, compliance with Basel guidelines, the degree of independence and legal power of the supervisory agency, and the authorities' effectiveness in imposing the legal framework. The fifth index quantifies the share of banking-sector assets that the state controls; we refer to this as the extent of privatization in the banking sector. The sixth index considers policies relating to international capital flows, such as restrictions on capital- and current-account convertibility, and the use of multiple exchange rates. The final index documents policies relating to securities markets. Included here are operational restrictions, such as restrictions on staffing, branching, and advertising. Establishing new securities markets is also in this category. We provide notations for these indices in Table I.

Even though Abiad, Detragiache, and Tressel (2010) label the general indicator a "financial reforms index," it primarily reflects policies related to the banking sector. In the empirical analysis we examine the effects each of the seven pillars have on income inequality, as well as the effect of the aggregate index, which is the sum of the seven indices and is labeled as "banking regulations." Fortunately, the database of Abiad, Detragiache and Tressel (2010) covers a large number of countries over a lengthy time period, whereas previous indices (e.g., Barth, Caprio, and Levine, 2006; Kaminsky and Schmukler, 2003; or the European Bank for Reconstruction and Development index of banking-sector reforms) are smaller in terms of years and/or countries. This large coverage makes our study possible and diminishes concerns about the number of available observations.

Theoretically, different forms of banking regulation can affect the distribution of income in different ways. For example, more stringent capital requirements and the efficient enforcement of prudential bank supervision (related to the fourth index) usually aim at reducing systemic risk and thereby buffer the economy from financial crises. If crises hurt primarily the poor, however, and if capital regulation lowers systemic risk, then capital regulation should lower income inequality. Repullo and Suarez (2008, 2009) highlight this procyclical effect of banking regulations in general and of Basel II in particular. Given that the majority of the related literature seems to agree that a negative correlation exists between the two (Barlevy and Tsiddon, 2006), then a negative correlation should also exist between banking regulations and inequality. In contrast, based on the fact that capital requirements exist in both good and bad times and that capital is expensive, more stringent capital requirements may raise banks' incentives to lend to "safer" individuals and firms rather than to relatively poor individuals, even if they are creditworthy or will generate income with the capital. This would be especially true when the financial system and the economy are "anxious" (Fostel and Geanakoplos, 2008).

In turn, the superior ability to enforce the abolition of interest rate and entry requirements (indices 2 and 3), the privatization of banks (index 5), and the liberalization and transparency of capital-account transactions (index 6) should improve financial intermediation services and project screening and monitoring. This would also give banks the liquidity to fund good investment ideas from individuals across the full spectrum of the income distribution, yielding a narrower income distribution and lower inequality (Beck, Demirgüç-Kunt and Levine, 2007).

The same outcome will prevail if enhanced privatizations, looser entry requirements, and liberalized capital-account transactions guarantee a more competitive and efficient banking sector. A bank's market power is usually associated with relationship lending, higher interest-rate margins, and entry restrictions, but these elements constitute barriers for individuals and firms with less collateral or poor credit. Therefore, we expect that abolishing interest rate controls and entry barriers, privatizing more banks, and liberalizing capital account transactions negatively relate to income inequality.

The potential impact of the liberalization of the securities markets on income inequality seems more difficult to predict. On one hand, liberalization enhances financial liquidity and increases the volume of lending. In line with the discussion of inequality, this would allow individuals at the lower end of the income distribution to have easier access to lending and capital and to fund their investment ideas more efficiently and at a lower cost. On the other hand, the recent financial crisis has shown that intense securitization leads banks to take excessive risk, which leads to a rise in the probability of bank failures. In particular, banks react to downturns by tightening their lending standards, which reduces lending to individuals with lower asset holdings and collateral. This would widen the distribution of income. In addition, liberalization could lead not to funding projects, but to investments in nontraditional activities. Thus, the overall impact of liberalized securities markets on income inequality is ambiguous.

Note, however, the implicit assumption that a low level of institutional quality offsets increases in supervisory power (e.g., government corruption is high or bureaucratic quality is low) and/or inappropriate economic development; apparently, this is the case for developing countries that face higher absolute poverty (e.g., Laffont, 2005). Identifying which effects prevail becomes fundamentally an empirical issue.

2.3 CONTROL VARIABLES

The control variables in this study are from the extensive literature on the determinants of income inequality (e.g., Roine, Vlachos and Waldenström, 2009; Beck, Demirgüç-Kunt and Levine, 2007). In particular, we control for a number of macroeconomic, institutional, demographic, and financial variables that affect income inequality. Because we use estimators based on fixed effects, we do not control for time-invariant variables.

First, we use the log of GDP per capita to control for the level of economic development, the inflation rate to control for monetary conditions, and the log of the population size to control for the demographics in each country. Information for these variables is from the World Development Indicators (WDI).[4] In keeping with the Kuznets hypothesis on the non-linear effect of economic development on income distribution, we also experiment by including the square of the log of GDP per capita.

Second, we control for trade openness and public sector growth. Our measure for trade openness is the sum of exports and imports as a share of GDP. Data come from the Penn World Table (PWT). The literature's empirical findings concerning the impact of trade openness on inequality are rather inconclusive. For example, trade openness increases income in standard Heckscher-Ohlin trade theory, but the extent to which it reduces inequality within countries remains questionable; most studies suggest an insignificant correlation (Easterly, 2005; Roine, Vlachos, and Waldenström, 2009). In turn, to account for the activity and growth of government over the sample period, we include the ratio of central government expenditures as a share of GDP (data also taken from the PWT). Higher government spending may disproportionately help the poor if used efficiently, but when institutions are weak (primarily in developing countries), higher government spending may be wasteful.

Usually included in equations characterizing income inequality is a variable that characterizes education. Often, this variable, provided by Barro and Lee (2001), addresses years of schooling. We use this data on primary education. In further sensitivity analysis we multiply the Barro-Lee indicator by the educational quality indicator "cognitive" developed by Hanushek and Woessmann (2009). The *"*cognitive*"* indicator is constructed on the basis of

[4] We also experiment with other macroeconomic variables, such as the unemployment and the GDP growth rate, a number of interest rates, the capacity utilization rate, etc. The results on our banking-regulation variables remain practically unchanged.

student performance in internationally comparable achievements tests[5] and allows us to capture potential qualitative differences on education among different countries. We find quantitatively similar results.

Furthermore, in order to purify the relationship between bank regulations and inequality from elements pertaining to the characteristics of the financial environment, we use two relevant control variables. As a proxy for the level of liquidity, we use the ratio of bank deposits to bank credit. The higher this ratio is, the higher the *Bank liquidity* variable. A higher liquidity ratio also indicates lower dependence on the banking sector in that country, as it reflects higher financial depth; thus, we expect it to relate negatively to inequality. In addition, we use a dummy variable, *Bank crisis*, which equals 1 when a country experiences a banking crisis.[6] We expect that banking crises widen the distribution of income by having a marginally more significant effect on the poor.[7]

Finally, we control for a number of political and institutional variables. In particular, we use information on the political orientation of the government in place (left or right) to examine whether left-leaning governments yield a narrower income distribution. In addition, we include an overall index of freedom (that excludes financial freedom) to guarantee that our

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[5] Although varying across the individual assessments, testing covers math, science, and reading for three age/grade groups: primary education (age 9/10), lower secondary education (age 13 to 15), and the final year of secondary education (generally grade 12 or 13). For more information on how this variable is constructed, see the Appendix of Hanushek and Woessmann (2009).

[6] On this front, a recent line of research suggests that banking crises may in fact be endogenous to inequality (e.g., Atkinson and Morelli, 2011; Kumhof and Ranciere, 2010; Stiglitz, 2009). The hypothesis comes from Stiglitz (2009), who suggests that in the face of stagnating real incomes, households in the lower part of the distribution borrow excessively to maintain a rising standard of living. Yet, in our case, we have 14 cases of banking crises during the period 1997-2005, most of which are related to the Asian crisis of 1997 and the Russian crisis of 1998. A significant common factor behind these crises is exchange rates, which do not have much to do with excess household borrowing. Still, we consider instrumenting the *banking crisis* dummy variable with the ratio of M2 to foreign exchange reserves of the central bank or with the share of public ownership of banks, both of which are significant determinants of banking crises in previous studies (e.g., Demirguc-Kunt and Detragiache, 1998, 2005) and do not seem to affect economic inequality. The results remain practically the same and are available on request.

[7] We also experiment with variables characterizing the mobility of funds across countries (ratio of offshore bank deposits to domestic bank deposits), the capitalization of the stock market (ratio of stock market capitalization to GDP), the performance of the banking sector (ratio of profits to total assets), the concentration in the banking sector (three-bank concentration ratio), etc. Most of these variables are only marginally correlated to our inequality measures and do not affect the impact of the banking regulations.

banking-regulation variables do not capture the overall political-liberalization processes that might have occurred within a country. Further, the presence of quality institutions will probably tend to lower inequality, even though strong endogenous effects may prevail in this relationship (Chong and Gradstein, 2007). Indeed, quality institutions might enhance the impact of regulations on the distribution of income and weaker institutions may undermine such an impact. To characterize the quality of institutions, we use the *Law and Order* and the *Transparency* (the inverse of Corruption) indices of the International Country Risk Guide (ICRG).[8] For explicit definitions of all these variables, see Table I; for descriptive statistics, see Table II; for a correlation matrix, see Table AII in the internet Appendix.

3. Econometric Identification

The empirical model is of the following form:

$$y_{i,t} = a_1 y_{i,t-1} + a_2 r_{i,t} + a_3 x_{i,t} + \lambda_t + v_i + u_{i,t}, \qquad (1)$$

where y is a measure of income inequality observed in country i at time t, r is the set of variables characterizing different types of banking regulations, x is the vector of control variables explaining y, u is the stochastic term, and λ and v are time- and country-effects, respectively.[9] The model could be dynamic due to persistence in inequality.

We seek a robust method to identify how bank regulations affect income inequality. The primary identification issue is the potential endogeneity of r. The major concern here is not that income inequality influences the choice of bank regulation (reverse causality), but that factors that influence banking regulation are also correlated with changes in income

[8] We also employ variables pertaining to democratic accountability, property rights, etc. However, these tend to be highly correlated among themselves and with the *GDP per Capita* variable.

[9] We keep the basic econometric model in levels because the time dimension of the dataset is small but the cross-sectional dimension is relatively large. Thus, we mitigate concerns about potential serial correlation of the error term. Using a correction of the error terms when applying instrumental variables and GMM also mitigates these concerns.

inequality. For example, the macroeconomic environment may simultaneously determine both elements (see e.g., Evans, 1997).

Likewise, Barth, Caprio, and Levine (2008) discuss two important episodes of large changes in the indices of banking regulation. First, Mexico responded to its 1994 crisis by easing restrictions on banks, but Argentina implemented greater regulatory restrictions after its crisis. Thus, if major economic turmoil drives changes in banking regulation, we should expect changes in banking regulation to correlate with changes in income inequality. This will happen as long as economic turmoil affects inequality. Moreover, this type of endogeneity will be relevant if politics drive changes in banking regulation, because changes in political equilibrium affect both regulatory reforms and other policies that influence income distribution.

To solve this problem, one can follow two strategies. The first is to focus on a specific episode and type of banking regulation and identify its impact on inequality within a single-country study. This strategy is the essence of Beck, Levine, and Levkov (2010). The second strategy identifies a clear source of change in banking regulations that is not highly correlated with political and economic sources of income inequality. This strategy allows for examining different types of regulation within a single empirical model. However, this comes at the cost of identifying proper instruments—a very difficult problem indeed.

Identifying a proper instrumental variable means ruling out institutional characteristics because of the potential causality with political reforms and thus with income inequality. The same holds for the legal-origin variables that many studies use in growth equations. The literature also uses geographic elements to identify growth equations, but establishing geographic elements as a source of banking regulation seems rather arbitrary.

Another intuitive approach is to consider the structural elements of the banking sector as potential elements affecting banking regulations. In particular, regulators might be interested in shaping banking industry concentration, given the trade-off between efficiency and concentration or concerns regarding foreign competition for domestic banks. The same may hold for the liquidity provided in the economic system and the importance of the banking sector in providing credit to the economy. Unfortunately, these elements may affect (or be affected by) economic outcomes, and through them, inequality.

An interesting case comes from elements of regulation not in the index of Abiad, Detragiache, and Tressel (2010). Barth, Caprio, and Levine (2006) and updates of this database are an exceptionally rich source of information on this front. From the various indices they discuss, we focus on the index named *Supervisory power*. This index indicates whether supervisory authorities can prevent and correct problems in the banking sector. Therefore, more supervisory powers reflect more stringent regulation of operational banking procedures.

Because these procedures deal primarily with corporate governance issues of everyday banking, they should not have any substantial real effects on income inequality. This is substantiated by statistical analysis showing a negative and statistically significant relationship between official supervisory power and financial reforms. This index bears no effect on any of our inequality indices.

Table AII reports pairwise correlation coefficients between *Supervisory power* and the indices from the database by Abiad, Detragiache, and Tressel (2010). The equivalent correlation coefficients between *Supervisory power* and our inequality measures are very low (-0.046 with the Gini coefficient and 0.004 with the Theil index). In addition, OLS regressions between our inequality measures and *Supervisory power*, with or without country fixed and time effects, indicate no causality running from supervisory power to the inequality measures. Therefore, we use *Supervisory power* as our instrumental variable. Note that this variable is available from 1997 onward, which is why our empirical analysis based on conventional econometric procedures is restricted to the period 1997-2005. However, as we show in Table AI, this is still a period of significant banking reforms in many regions of the world.[10]

Given the persistence of inequality and the presence of the dynamic term y_{t-1} among the regressors, we have to use the generalized method of moments (GMM) for dynamic panels along with two-stage least squares (2SLS). However, the latter is also an efficient estimator given that we primarily resort to three-year averages of the data. Also, 2SLS allows using fixed effects that might improve the precision of our estimates (see e.g., Acemoglu, Johnson, Robinson and Yared, 2008). In turn, the GMM procedure is the system GMM estimator of Blundell and Bond (1998).

This method assumes that lagged values of the dependent and independent variables are valid instruments under certain restrictions. In our setting, however, including lagged values of *r* and *x* as additional instruments might not be a good idea. Political and institutional variables change slowly, and previous changes might correlate with contemporaneous levels of inequality. Thus, in the equations estimated by GMM, we only include lagged levels of *y* among the instrumental variables.[11] With these issues in mind, we turn to the estimation results.

4. Empirical Findings

In this section we present the main results of the study and conduct various sensitivity analyses to assess whether results change (i) when choosing between the Gini coefficient or the Theil index, (ii) for different levels of economic development, (iii) when including

[10] A potential drawback of this instrumental variable could be that it reflects the presence of influential socio-economic elites, which can have a direct effect on income inequality. We experiment with many potential measures of elites (see e.g., Angeles, 2007; Angeles and Neanidis, 2009). Using variables that capture the presence of an influential elite as in Barth, Caprio, and Levine (2006) and Morck, Yavuz, and Yeung (2011) we find that the correlations between these measures and the supervisory power index are very low and statistically insignificant.

[11] In a previous version of this paper we also used a panel VAR approach, in the fashion of Love and Zichinno (2006) and Holtz-Eakin, Newey, and Rosen (1988) to identify the isolated impact of bank regulations on inequality. The analysis was carried out for the full 1973-2005 period using annual data. In general, the results were similar to the ones reported below and are available on request.

alternative control variables among the regressors, (iv) when using annual, three-, and nine-year intervals for the data, (v) when employing the different estimation methods proposed earlier, and (vi) when dropping each region defined in Table AI in turn.[12]

4.1 BANK REGULATIONS AND INCOME INEQUALITY: RESULTS FOR THE EFFECT OF THE GENERAL INDICATOR OF BANK REGULATIONS

Panels with large cross-sectional and relatively small time dimensions are usually prone to considerable heteroskedasticity, and a simple likelihood ratio test shows that our panel is no exception. Therefore, in applying 2SLS or GMM we use robust standard errors. The first set of empirical results is in Table III, where the dependent variable is the Gini coefficient and the main explanatory variable is the aggregate index of banking regulations.

We start with a simple OLS regression, with country fixed effects and robust standard errors, of banking regulation on the Gini coefficient. The result poses an immediate challenge, as it shows that regulations increase income inequality. Of course, this finding is counterintuitive and probably driven by omitted-variable and endogeneity bias. In the second regression (column 2) we add simple time effects, and the impact of bank regulations on the Gini coefficient becomes insignificant.

[INSERT TABLE III]

Following our discussion in Section 3, we turn to identification through instrumental variables techniques. In columns (3) to (5) we report the results from a 2SLS regression with country fixed and time effects, as well as robust standard errors. The instrument used is the index of *Supervisory power* (defined in Table I). We report the first-stage results for these regressions in the lower part of Table III. The results show that the *Supervisory power*

[12] Another drawback to the empirical analysis is that outliers drive results. To determine whether our results are sensitive to outliers, we perform a jackknife analysis (Efron and Tibshirani, 1993). This involves estimating the initial equation by excluding in each replication one or more cross-sectional units (countries). Our results are robust to the exclusion of particular observations that yield extreme estimates; hence, outliers do not substantially affect the main implications of the paper.

variable is a negative and highly statistically significant determinant of banking regulation, which is in line with our expectations. As discussed in Section 3, this instrument is an insignificant determinant of the Gini coefficient in a simple fixed-effects regression with robust standard errors and time effects. The full array of the first-stage results is available on request.

The second-stage results in column (3) show that higher *Bank regulations* values (reflecting a more liberalized banking system) are associated with lower Gini values, and this effect is statistically significant at the 5% level. The economic effect is also substantial. According to the results of column (3), a one-point increase in *Bank regulations* lowers the Gini coefficient by 5.2%. In Lithuania, for example, the 3.6-point rise in *Bank regulations* over the sample period is equivalent to an 18.7% reduction in the Gini coefficient, *ceteris paribus*. Therefore, the slight increase in the Gini, from an average of 47.19 in the period 1997-1999 to an average 47.75 in the period 2003-2005, would have been much higher if the banking sector had not been liberalized.

In India, *Bank regulations* rises by 3.5 points. The associated reduction in Gini is 18.2%, *ceteris paribus*. Here, the Gini coefficient rises from an average of 39.5 in the period 1997-1999 to an average 43.5 in the period 2003-2005, and it would have been lower in the latter period if the only driving force were banking liberalization policies. In general, and given that the Gini coefficient trends upward in most countries, the liberalization of banking systems contains this upward trend by creating opportunities for those at the lower end of the income distribution. Furthermore, if we assume that the relationship between banking sector reforms and the Gini coefficient is stable over time, the cumulative benefit would be much higher after the banking sector reforms initiated in the 1980s. It remains to be seen whether all types of bank regulation contribute to this effect.

In columns (4) and (5) we add more explanatory variables. The statistical and economic significance of *Bank regulations* remain practically unchanged. Among the rest of the explanatory variables, the most interesting results are those showing a negative effect between (i) the log of *GDP per Capita* and Gini, (ii) between *Bank Liquidity* and Gini, and (iii) between *Education* and Gini. In the literature, the relationship between economic development and inequality is primarily negative (Bourguignon, 1996), but controlling for the impact of the economic development seems crucial because the *Bank regulations* variable may capture the positive trend in development. Yet, the coefficient on *Bank regulations* remains significant, even though *GDP per Capita* enters with a negative and significant coefficient.

Similarly, the impact of *Bank regulations* on inequality does not change when controlling for the increasing loans to deposits ratio, even though a higher ratio translates to a narrower distribution of income. Finally, we find that the higher the level of primary education, the lower the inequality. This finding has ample support in the relevant literature (e.g., Barro, 1999).

In columns (6) to (8) we report the results of re-estimating the previous three regressions, this time using GMM for dynamic panels. The model now includes the lagged dependent variable y_{t-1}, and as an additional instrument, y_{t-2}. The identification tests show no overidentifying restrictions and no serial correlation between the instruments and the disturbance. Compared to columns (3) to (5), the results remain practically unchanged.

The impact of *Bank regulations* on the Gini coefficient is negative and significant. The relevant coefficient is around 0.05, very close to the one observed when the estimation method is 2SLS. One difference in the results, however, is the positive and significant coefficient on *Trade Openness* in column (9). This finding suggests that globalization is partially responsible for the widening income distribution; however, to make such a statement one has to go much deeper. Here, we only document that the impact of bank regulations on inequality is not primarily due to some form of trade openness.

In Table IV we measure inequality via the Theil index and rerun specifications (1) to (5) of Table III. We are unable to estimate the model using GMM for dynamic panels, because only two periods of data are available for the Theil index (1997-1999 and 2000-2002). Again, and despite the reduced number of observations, *Bank regulations* enters with a negative and significant coefficient. However, the estimated elasticity is somewhat larger. For example, in columns (4) and (5), which represent regressions with controls, the relevant coefficients are -0.088 and -0.083, respectively.

These findings confirm that regulatory stringency increases income inequality, despite whether the inequality is in household income or individual income. More important, the fact that *Bank regulations* has a larger impact on the Theil index than on the Gini coefficient suggests that much of the effect of bank liberalization on income inequality comes from wage income, which the Theil index encompasses rather than the Gini. Notably, this finding is in line with Beck et al. (2010). Among the rest of the controls, the most significant ones remain the *GDP per Capita* and *Bank Liquidity* variables.

[INSERT TABLE IV]

We conduct further sensitivity analyses on these results and report the findings in Table V. Because the results from choosing between 2SLS or GMM and between the Gini coefficient or the Theil index are not different, we only report the results from the computationally simpler 2SLS and from the Gini coefficient. In column (1) we introduce a multiplicative term between *Bank regulations* and *GDP per Capita* to examine whether the impact of *Bank regulations* differs with the level of economic development. In this specification we mean-center the variables and find the product of the demeaned terms. This allows for examining the effect of *Bank regulations* at the average *GDP per Capita* (and not for *GDP per Capita* equal to 0, which is not interesting) and reduces multicollinearity.

A series of important contributions (e.g., Estache and Wren-Lewis, 2009) view economic and institutional development as prerequisites for regulations to have a real effect on the economy. In fact, as Bourguignon (2005) states, “Today, it is increasingly recognized that, in many circumstances, the problem [in the developing countries] was that reformers disregarded the functioning of regulatory institutions, assuming implicitly they would work as in developed countries.”[13] The results show that the literature’s implications are valid: the negative impact of bank regulations on inequality weakens substantially for countries with low GDP per capita.

[INSERT TABLE V]

In column (2) we introduce the squared term of *GDP per capita* among the regressors to examine whether a nonlinear relationship between *GDP per capita* and inequality (Kuznets hypothesis) affects the results. Again, we mean-center the involved variable. We do not find evidence of nonlinearity, which is in line with Deininger and Squire (1998). A further robustness check includes the initial Gini coefficient and its squared term as in Clarke, Xu, and Zhu (2006); we also use an estimation method based on random effects. Even in this exercise, the relevant coefficients turn out to be statistically insignificant. The latter findings are available on request.

In columns (3) and (4) we introduce the institutional and political variables, respectively. This guarantees that the *Bank regulations* variable does not capture any general institutional characteristics of countries or the general political freedom or ideology of the specific time frame. Here we remove the macroeconomic variables because of high correlations between them and the institutional and political variables. The findings show that

[13] We also experiment with the product of our regulatory variables with the institutional variables (i.e., *Bureaucratic Quality* and *Law & Order*). However, as shown in Table IV, these variables are highly correlated with the *GDP per Capita* variable and inference is unaltered.

lower Gini coefficients are associated with stronger institutions and more political freedom. Left-leaning governments are associated with lower Gini coefficients, but the effect is not significant at conventional levels. The impact of *Bank regulations* is in fact stronger in these specifications, and the coefficients obtain a value of 6% and above.

In column (5) we examine whether the results differ when we estimate our basic specification only for bank-based economies. To define bank-based economies (as opposed to market-based ones) we use the ratio of private credit provided by banks to stock market capitalization (for a similar definition, see Demirgüç-Kunt, Levine and Beck, 2002). Countries with ratios below the mean value of this variable are bank-based economies.[14]

The results show that the negative impact of bank regulations on income inequality for these countries is much stronger. As the relevant coefficient indicates, a one-point increase in *Bank regulations* decreases the Gini coefficient by 10.6%. For India, which is a bank-based economy, *Bank regulations* rises by 3.5 points. The associated reduction in Gini over the sample period is 37.1%, *ceteris paribus*, which is a very large effect.

The respective effect on market-based economies is only 2.6% per point increase in *Bank regulations*, with the effect being statistically significant only at the 10% level (results are available on request). Therefore, for Chile, which is a market-based economy and saw a rise in *Bank regulations* of 2.33 points, the respective reduction in inequality is approximately 6%. This shows that bank regulations can control income inequality in both groups, but mostly in bank-based economies.

The results discussed so far are robust for the three-year time periods. Thus, we examine whether the results hold for one- and nine-year time intervals, and we report the findings in columns (6) and (7) of Table V. With very few differences, the findings are equivalent to those of previous regressions. When we use annual data, most of the coefficients

[14] Admittedly this is only one of the many measures that could define bank- versus market-based economies. However, more analyses on this front are beyond the scope of the study.

show stronger effects, which is probably due to the higher number of observations and the short-run fluctuations of variables. Similarly, the coefficients in column (7) are between those reported in column (6) and those of the baseline specifications in Table III.

In an important robustness check, we consider the impact of incorporating the standard errors assigned by Solt (2009) on the Gini coefficients (see our discussion in footnote 2). We proceed in two ways. The first involves dropping the coefficients for countries with standard errors higher than 5 (40 observations in the three-year panel). The second, which is the preferred method, involves generating Monte Carlo simulations for the Gini index and perform the analysis using the output of these simulations (see Solt, 2009). The results from both methods are similar, and we report the results from the preferred method in column (8) of Table V. Changes in the results are minimal compared with those of Table III.

Given that the results are somewhat different for bank- versus market-based economies, another robustness check examines whether countries from a particular region drive the results. For example, Table AI shows that 29% of the sample countries are advanced economies and 42% of the countries that have reformed their banking systems are transition countries. In Table AIII in the internet Appendix, we report the results of the Gini coefficient regressions, where we exclude from the full sample one region in turn. Changes in the results are minimal and still reflect that for every 1% increase in *Bank regulations* there is at least a 5% reduction in the Gini index.

In Table VI we report the results from using as dependent variables (i) the income share held by the highest 10% of the income distribution, (ii) the income share held by the lowest 10%, and (iii) the poverty gap at $1.25 a day, respectively. These regressions show whether banking sector reforms disproportionately benefit or hurt the poor and whether they affect the absolute level of poverty significantly.

Columns (1), (3), and (5) report the regression results that include only bank regulations, as well as fixed and time effects. In the regressions in columns (2), (4), and (6) we add all the other explanatory variables as in specification (5) of Table III. The results show that banking liberalization primarily increases the income share of the relatively poor. Further, the impact of the *Bank regulations* variable on the poverty gap at $1.25 a day is negative and significant at the 10% level. In column (4) the economic effect is quite large, as a 1% increase in Bank regulations increases the income share held by the lowest 10% of the income distribution by 0.5 percentage points. In other words, even a moderate increase in Bank regulations from 16 to 17.6, which we see in many countries over the sample period, will see a 0.53 point increase in the income share held by the lowest 10% of the income distribution. This implies that if the income share held by the lowest 10% initially is 2.41 (i.e., approximately equal to the mean value), the income share will increase to 2.94 within three years.

[INSERT TABLE VI]

These findings are in line with Galor and Moav (2004), who argue that failing to liberalize the banking sector hurts the poor. The impact of banking sector liberalization on the income share held by the top 10% of the income distribution is positive and significant at the 10% level, which shows that banking sector reforms do not hurt the relatively rich; in fact, such reforms marginally improve their income share.

This finding seems to follow Greenwood and Jovanovic (1990). Their theoretical model predicts that in transition economies, accelerating financial development is likely to encourage intermediaries to lend to individuals and firms with higher accumulated wealth. Strong political networks and abrupt liberalization of product markets could further enhance this mechanism, a situation in countries with relatively weak institutions (such as the ones in these regressions). However, and given our discussion, our results overall conflict with the

Greenwood and Jovanovic (1990) model in that the poor benefit more than the rich from banking liberalization.

A final element is that the *Bank crisis* dummy is significant for the first time, which shows that banking crises hurt both the highest and the lowest incomes but more significantly hurt the poor. However, interpret these results with caution because, as discussed, the values of the dependent variables are not observed in Western-type economies and are not directly comparable across countries. The first element reduces the number of available observations and introduces selectivity bias; the second element introduces some measurement error.

We repeat this exercise for the regional groups in Tables AI and AIII. We exclude the advanced economies because data for the income-distribution variables are very limited for these countries. We report in Table AIV of the internet Appendix the coefficient estimates and the associated robust standard errors on the *Bank regulations* variable. The picture is not very different from the one reported in Table VI. The impact of *Bank regulations* on the income share of the relatively rich (reported in panel A) seems to lose ground in statistical significance when we exclude the groups of countries with the larger number of observations (especially Latin America and transition countries). In contrast, the coefficients reflecting the impact of *Bank regulations* on the income share held by the lowest 10% remain positive and strongly significant when we exclude one regional group in turn (see results in Panel B). Finally, four out of five regressions in Panel C indicate that banking sector liberalization lowers the poverty gap at $1.25 a day; this effect is statistically significant at the 10% level.

Overall, this analysis highlights that a clear trade-off exists between stricter banking regulation and long-term income equality, and although a consensus seems to indicate that stricter regulatory policies promote stable banking systems, these policies still disproportionately hurt the poor. This finding is in line with Beck, Levine, and Levkov

(2010), which shows that deregulating the banking system in the United States in the 1970s and 1980s led to increased incomes, particularly for the poor.

4.2 BANK REGULATIONS AND INCOME INEQUALITY: RESULTS FOR THE EFFECT OF THE DIFFERENT REGULATORY POLICIES

Do all types of regulation have the same impact on income inequality? It seems highly unlikely, and the results in this subsection confirm this. To save space, we only report the results from the equivalent regression (5) of Table III, using each one of the seven components of the *Bank regulations* variable as independent variables. The rest of the 2SLS regressions and regressions based on the Theil index confirm these findings.

In Table VII we present the results using each one of the seven components of the bank regulations index in Abiad, Detragiache, and Tressel (2010) as the main independent variable. All individual subindices (with the exception of the liberalization of security markets) have a negative and significant effect on the Gini index. The most significant impact, both statistically and economically, comes from liberalizing interest rate controls and privatizating the banking sector. In contrast, liberalizing security markets seems to have a positive impact on income inequality.

[INSERT TABLE VII]

The impact of credit and interest rate controls on the Gini coefficient (see columns 1 and 2, respectively) is negative and statistically significant. The coefficient on *Credit controls* is -0.055, whereas the coefficient on *Interest rate controls* is even higher (-0.092). The results imply that credit controls lower liquidity and work against the poor. This seems straightforward considering that higher restrictions tend to produce less competitive markets, which tend to reduce the quality of project screening and monitoring. Under these conditions, relationship lending or lending to well-established firms with high levels of collateral and strong credit history prevail, thereby constraining access to credit for the relatively poor.

The impact of entry barriers is negative but statistically insignificant only at the 10% level, but the impact of privatization is negative and significant at the 1% level. Given that most developed countries abolished entry restrictions by the early 1990s, the developing and transition markets probably contribute to the relatively weak relationship between entry barriers and income inequality. In addition, these economies are usually characterized by inferior institutions and the partial inability to enforce the law.[15]

The relationship between the Abiad, Detragiache, and Tressel (2010) index of banking supervision and income inequality is also negative (column 4). This is a very interesting finding, given the ongoing discussion on the reregulation of the banking system. As discussed, a higher index value reflects more stringent capital regulation, more monitoring of bank activities through audits and sanctions for prudential purposes, etc. Thus, this finding supports the position that improved screening and monitoring of investment projects and more competition in banking markets drives funds to the best investment ideas and thus provides equal opportunities to the poor. Hence, financial stabilization aside, efficient supervision also seems to have a substantial, real, and positive effect in lessening income inequality, thus allowing equal opportunities in accessing credit and sustaining economic fairness.

A final notable result in Table VII is the positive, albeit relatively small, response of the Gini coefficient to a shock in securities markets (see column 7). This finding suggests that liberalizing securities markets leads to higher income inequality. A mechanism that explains this finding might be that banks gradually lend more after liberalization of credit and interest rates, but they also expand their involvement in securities markets to safeguard themselves

[15] In additional analysis, we run a panel VAR, which distinguishes between countries with GDPs per capita above the sample mean and those below it. A Gini response to a positive shock in entry barriers and privatizations is more significant in countries with GDP per capita above our sample's average. The results are available on request.

against the higher credit risk. This suggests a trade-off between liquidity going to securities markets and liquidity available to finance investment projects. That is, when banks become more involved in the securities markets, they tend to fund fewer projects (especially involving individuals with less collateral and credit history), which widens income distribution.

As a final exercise, we re-estimate equations with the Theil index, the income share held by the highest 10% of the income distribution, and the income share held by the lowest 10%. Similar to the previous table, we distinguish between the different sources of bank liberalization policies. Because this implies estimating 21 regressions, we report in Table AV in the internet Appendix only the results on the coefficients related to the regulatory indices.

The results are somewhat weaker. This probably comes from the lower number of available observations and/or the lower quality of the WDI data for the income distribution variables. Still, credit and interest rate controls, as well as privatization policies, are associated with a lower Theil index. Liberalizing securities markets tends to increase the Theil index, but this effect lightens in statistical and economic significance. Among the different liberalization policies, abolishing entry barriers and enhancing international capital flows increase the income share of the relatively rich considerably. However, this is not at the expense of the poor: none of the seven indices has a negative and significant coefficient on the regressions with the income share of the lowest 10% as the dependent variable.

In contrast, liberalizing credit and interest rate controls, as well as increasing international capital flows are the main sources of an increase in the income share of the poor (see columns 1, 2, and 6 of Panel C). Even though these findings should be treated with some caution, their interpretation is quite clear: the liberalization policies considered here do not take a toll on the income share of the poor; in fact, quite a few policies, especially those directly associated with credit availability and its pricing, increase it.

Overall, these findings imply something that the regulatory literature overlooks: bank regulation can, successfully in many circumstances, strengthen the financial system and absorb failures that lead to crises. Furthermore, most banking liberalization policies enhance the availability of credit, provide funding opportunities, and lead to a narrower income distribution. However, regulatory policies aimed at short-term financial stability (such as higher capital requirements) or short-term liquidity (such as liberalizing securities markets) have an adverse long-term effect on income equality.

5. Conclusions and Policy Considerations

This study links, for the first time, the full array of banking regulations with income inequality. We show that the banking liberalization policies contribute significantly to containing income inequality. Yet, the pattern is not similar across all regulatory policies, countries with different levels of economic and institutional development, and market-based versus bank-based economies.

In particular, abolishing credit controls decreases income inequality substantially. Interest rate controls and tighter banking supervision also decrease income inequality. In turn, abolishing entry barriers and enhancing privatization laws seem to lower income inequality, primarily in developed countries. In contrast, liberalizing securities markets increases income inequality. These results are robust to a number of estimation methods that account, *inter alia*, for the endogeneity of banking regulations.

What are the policy implications of these findings? Bank regulations and associated reforms aim at enhancing the creditworthiness of banks and at improving the stability of the financial sector. Several studies over the last decade show that regulations do matter in shaping bank risk (e.g., Laeven and Levine, 2009; Agoraki, Delis, and Pasiouras, 2011), bank efficiency (Barth, Lin, Ma, Seade, and Song, 2010), and the probability of banking crises (e.g., Barth, Caprio, and Levine, 2008). Yet, what if bank regulations have other real effects on the economy? And more important, what if these real effects counteract the intended stabilizing effects?

Answering these questions requires consideration of two issues. First, the literature on the relationship between bank regulations and financial stability is inconclusive. In fact, different types of regulation may have opposing effects on financial stability, according to the existing research. Second, even if we assume that bank regulations lower banks' risk-taking appetites and enhance stability, the empirical findings here suggest that these effects are asymmetric and that certain liberalization policies (i.e., liberalizing securities markets) or reregulation policies actually increase income inequality. That is, banks pass the increased costs of higher risks onto the lower-income population that lacks good credit and collateral. It is a trade-off between banking stability and income equality. Given the contemporary discussion surrounding the rebirth of Glass-Steagall-type reforms in securities trading and Basel III discussions to increase banks' risk-adjusted capital bases, there may be more to think about before taking those steps.

That said, three clear suggestions emerge from this paper and are consistent with Beck, Levine, and Levkov (2010). : First, liberalizing banking markets, primarily via efficient banking supervision and abolishing credit controls, helps the poor get easier access to credit. This in turn allows them to escape poverty and substantially raise their incomes. Second, appropriate prudential regulation should provide less costly incentives to banks to increase regulatory discipline without hurting the poor. Information technologies that lower the cost of transparency and more effective onsite supervision that enhance the integrity of the banking system may help achieve this goal. Finally, economies first need a certain level of economic and institutional development in order to obtain any positive effect from the abolishment of entry restrictions and privatizations on equality. Thus, though this type of deregulation had a negative impact on inequality in the United States, this may not be true for countries with weak institutions in which socioeconomic elites directly affect bank supervisors' decisions and policies.

Clearly more research is necessary. First and foremost, if data-quality concerns are dropped, researchers should study the effect of bank regulations on the incomes of individuals in the top and bottom of the income distribution. Another interesting extension relates to the potential impact of bank regulations on macroeconomic convergence (Evans and Karras, 1996) or the speed of convergence (Evans, 1997). In addition, the interplay between regulations and their actual implementation may have more to say about credit availability and income inequality. Finally, more detailed datasets from both developed and developing countries could highlight the channels that may affect the nexus of bank regulations and income inequality, with an emphasis on the impact of banking crises. We leave these ideas for future research.

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Table I. Variable definitions and sources

The table reports notations, measures, and data sources for all the variables in the empirical analysis.

Notation	Measure	Data Source
A. Dependent Variables		
Gini coefficient	Measure of inequality obtained from the Lorenz curve (natural logarithm) on the basis of household income before taxes.	Standardized World Income Inequality Database by Solt (2009)
Theil index	Wage inequality measure (natural logarithm) based on wage income.	Texas Inequality Project and UTIP-UNIDO Industrial Statistics
Income share held by highest 10%	The income share held by the highest 10% of the income distribution.	World Development Indicators
Income share held by lowest 10%	The income share held by the lowest 10% of the income distribution.	World Development Indicators
Poverty gap at $1.25 a day	The mean shortfall from the poverty line (counting the non-poor as having zero shortfall), expressed as a percentage of the poverty line. This measure reflects the depth of poverty as well as its incidence.	World Development Indicators
B. Explanatory Variables		
Credit controls	Index of liberalization of credit controls and reserve requirements for banks. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
Interest rate controls	Index of liberalization of interest rates. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
Entry barriers	Index of lower entry barriers for banks. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
Banking supervision	Index of prudential regulation and supervision of the banking sector. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
Privatization	Index of the extent of privatization of banks. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
International capital flows	Index of liberalization of international capital flows. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
Securities markets	Index of liberalization of securities markets. Takes values between 0 and 3. The variable is used in logarithmic terms.	Abiad et al. (2010)
Bank regulations	This index is constructed on the basis of the sum of the values of the seven indices above, thus it takes values between 0 and 21. Higher values reflect more liberalized banking sectors. The variable is used in logarithmic terms.	Abiad et al. (2010)

Population	The natural logarithm of the population of a country.	World Development Indicators (World Bank)
GDP per capita	The natural logarithm of the gross domestic product per capita.	World Development Indicators (World Bank)
Trade openness	The ratio of the sum of exports and imports over GDP.	Penn World Tables 6.3
Government expenditure	The ratio of government expenditures over GDP.	The Heritage Foundation
Inflation	The CPI inflation rate.	World Development Indicators (World Bank)
Bank liquidity	The ratio of bank deposits to bank credit.	Beck and Demirgüç-Kunt (2009)
Bank crisis	Dummy variable equal to 1 during a period that a banking crisis occurs.	Laeven and Valencia (2008)
Education	Barro-Lee (2001) years of schooling. The variable is used in logarithmic terms.	Barro and Lee (2001)
Left power	Chief executive and largest party in congress have left-center political orientation.	Botero et al. (2004)
Law and order	Law and order are assessed separately, with each subcomponent worth 0 to 3 points. The law subcomponent is an assessment of the strength and impartiality of the legal system, while the order subcomponent is an assessment of popular observance of the law. The variable is used in logarithmic terms.	International Country Risk Guide
Transparency	Index of transparency (inverse of corruption), taking values from 0 to 6. The higher the value of the index, the more transparent is the country. The variable is used in logarithmic terms.	International Country Risk Guide
Nonfinancial freedom	Overall index of freedom minus the value on the financial freedom component. The variable is used in logarithmic terms.	The Heritage Foundation
C. Instrumental Variable		
Supervisory power	Index of the powers of the supervisor of the banking sector, reflecting whether the supervisory agency has the authority to take specific actions to prevent and correct problems in the banking sector. The variable is used in logarithmic terms.	Barth et al. (2006) and updates

Table II. Descriptive statistics and sample information

The table reports the number of available observations and summary statistics (mean, standard deviation, minimum, and maximum) for the variables in the empirical analysis over the period 1997-2005. The number of observations refers to the panel constructed by taking three-year averages. The variables are defined in Table I. Population is in thousands, and GDP per capita is in $US.

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
Gini coefficient	266	44.68	7.02	21.59	69.74
Theil index	156	0.07	0.06	0.00	0.34
Income share held by highest 10%	121	32.46	8.13	19.04	53.26
Income share held by lowest 10%	121	2.41	1.04	0.20	4.52
Poverty gap at $1.25 a day	101	7.05	10.35	0.21	52.76
Credit controls	262	2.39	0.74	0.00	3.00
Interest rate controls	262	2.78	0.54	0.00	3.00
Entry barriers	262	2.62	0.71	0.00	3.00
Banking supervision	262	1.67	0.79	0.00	3.00
Privatization	262	1.81	1.13	0.00	3.00
International capital flows	262	2.36	0.84	0.00	3.00
Securities markets	262	2.15	0.84	0.00	3.00
Bank regulations	262	15.78	3.56	5.00	21.00
Population	293	42,800	145,000	78.8	1,300,000
GDP per capita	293	10,820.12	10,988.35	299.67	66,165.13
Trade openness	292	84.97	52.23	2.01	424.00
Government expenditure	291	65.26	23.89	0.00	99.07
Inflation	293	10.56	32.17	-7.17	359.87
Bank liquidity	292	0.88	0.41	0.09	2.99
Education	276	5.02	2.92	0.04	12.25
Left power	237	0.55	0.33	0.00	1.00
Law and order	275	3.96	1.33	1.00	6.00
Transparency	275	3.30	1.43	0.29	6
Nonfinancial freedom	262	52.41	18.98	10	90
Supervisory power	286	11.01	2.22	4	14

Table III. Bank regulations and income inequality: Gini coefficient regressions

The table reports coefficient estimates and robust standard errors (in parentheses) of the estimation of Equation (1) over the period 1997-2005 using three-year averages. The dependent variable is the natural logarithm of the Gini coefficient. The explanatory variables are defined in Table I. Regressions (1) and (2) are estimated by OLS for panel data with country fixed effects and robust standard errors. Regressions (3)-(5) are estimated by 2SLS with country fixed effects, the instrumental variable being *Supervisory power* as defined in Table I. First-stage results for the instrumental variable are reported in the lower part of the table. Regressions (6)-(8) are estimated using GMM for dynamic panel data, and the second lag of Gini is added to *Supervisory power* as an instrumental variable. The overidentification test is the p-value of the Hansen J-statistic for the overidentification of the equation, and rejection of the null casts doubt on the validity of the instruments. AR2 test is the p-value of the test for second-order autocorrelation in first differences, and rejection of the null implies presence of such autocorrelation. The ***, **, and * marks denote statistical significance at the 1%, 5%, and 10% level, respectively.

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS	(5) 2SLS	(6) GMM	(7) GMM	(8) GMM
Lagged Gini						0.552**	0.408**	0.472**
						(0.235)	(0.177)	(0.184)
Bank regulations	0.037***	0.004	-0.052**	-0.050**	-0.048**	-0.050**	-0.052**	-0.050**
	(0.006)	(0.009)	(0.022)	(0.022)	(0.021)	(0.023)	(0.023)	(0.024)
Population				-0.144***	-0.202***		0.017	0.006
				(0.046)	(0.043)		(0.029)	(0.027)
GDP per capita				-0.088***	-0.097***		-0.052***	-0.047**
				(0.029)	(0.026)		(0.018)	(0.018)
Trade openness				0.009	-0.030		0.081	0.057
				(0.029)	(0.022)		(0.049)	(0.040)
Government expenditure				0.028 (0.021)	0.022 (0.029)		0.040 (0.038)	0.024 (0.038)
Inflation					0.006**			0.005**
					(0.003)			(0.002)
Bank liquidity					-0.026**			-0.036**
					(0.012)			(0.015)
Bank crisis					0.064			0.031
					(0.067)			(0.105)
Education					-0.248**			-0.264**
					(0.107)			(0.108)
Observations	260	260	260	254	241	260	254	241
Time effects	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adj. R-squared	0.145	0.264						
Overidentification test						0.340	0.332	0.325
AR2 test						0.156	0.187	0.230
First stage results								
Supervisory power			-0.471***	-0.382***	-0.380***			
			(0.105)	(0.117)	(0.117)			
R-squared			0.078	0.366	0.380			
F-statistic (p-value)			0.000	0.000	0.000			

Table IV. Bank regulations and income inequality: Theil index regressions

The table reports coefficient estimates and robust standard errors (in parentheses) of the estimation of Equation (1) over the period 1997-2002 using three-year averages. The dependent variable is the natural logarithm of the Theil index. The explanatory variables are defined in Table I. Regressions (1) and (2) are estimated by OLS for panel data with country fixed effects. Regressions (3)-(5) are estimated by 2SLS with country fixed effects and robust standard errors, the instrumental variable being *Supervisory power* as defined in Table I. First-stage results for the instrumental variable are reported in the lower part of the table. The ***, **, and * marks denote statistical significance at the 1%, 5%, and 10% level, respectively.

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS	(5) 2SLS
Bank regulations	0.009	-0.012	-0.122***	-0.088**	-0.083**
	(0.028)	(0.011)	(0.043)	(0.042)	(0.041)
Population				-0.215***	-0.205***
				(0.071)	(0.051)
GDP per capita				-0.086***	-0.082***
				(0.025)	(0.025)
Trade openness				0.016*	0.012
				(0.009)	(0.012)
Government expenditure				0.016	0.038*
				(0.022)	(0.022)
Inflation					0.002*
					(0.001)
Bank liquidity					-0.028**
					(0.011)
Bank crisis					0.049
					(0.054)
Education					-0.277**
					(0.115)
Observations	151	151	151	148	142
Time effects	No	Yes	Yes	Yes	Yes
Adj. R-squared	0.091	0.088			
First stage results					
Supervisory power			-0.274***	-0.301***	-0.299***
			(0.083)	(0.088)	(0.089)
R-squared			0.232	0.248	0.258
F-statistic (p-value)			0.000	0.000	0.000

Table V. Bank regulations and income inequality: Sensitivity analysis

The table reports coefficient estimates and robust standard errors (in parentheses) of the estimation of Equation (1) over the period 1997-2005. The dependent variable is the natural logarithm of the Gini coefficient. The explanatory variables are defined in Table I. Regressions (1)-(5) and (8) are estimated using three-year averages, regression (6) using annual data, and regression (7) using averages over the period 1997-2005 (cross-sectional analysis). All regressions are estimated using 2SLS with country fixed effects and time effects, except for regression (6), which is a cross-sectional analysis and is estimated using simple 2SLS. The instrumental variable is *Supervisory power*, as defined in Table I. First-stage results for the instrumental variable are reported in the lower part of the table. The ***, **, and * marks denote statistical significance at the 1%, 5%, and 10% level, respectively.

	(1) Level of development	(2) Non-linear economic development	(3) Politics	(4) Institutions	(5) Bank-based economies	(6) Annual panel	(7) Cross-sectional analysis	(8) Sensitivity to Gini standard errors
Bank regulations	-0.045**	-0.051**	-0.062**	-0.066**	-0.106***	-0.135***	-0.045**	-0.055**
	(0.019)	(0.021)	(0.025)	(0.026)	(0.033)	(0.042)	(0.021)	(0.023)
Population	-0.186***	-0.199***	-0.220***	-0.216***	-0.121*	-0.267***	-0.203***	-0.185***
	(0.051)	(0.054)	(0.056)	(0.055)	(0.063)	(0.059)	(0.063)	(0.052)
GDP per capita	-0.082***	-0.090***			-0.093***	-0.107***	-0.076***	-0.090***
	(0.029)	(0.029)			(0.028)	(0.030)	(0.027)	(0.026)
Bank regulations* GDP per capita	0.080**							
	(0.032)							
GDP per capita squared		0.002						
		(0.005)						
Trade openness	0.060*	0.062*			0.048	0.078**	0.030	0.016
	(0.036)	(0.037)			(0.047)	(0.039)	(0.046)	(0.018)
Government expenditure	0.027	0.011			0.016	0.028	0.004	0.021
	(0.022)	(0.030)			(0.027)	(0.028)	(0.049)	(0.018)
Inflation	0.009***	0.011***			0.006**	0.013***	0.006**	0.006**
	(0.003)	(0.004)			(0.002)	(0.004)	(0.003)	(0.003)
Bank liquidity	-0.036**	-0.034**			-0.041**	-0.062***	-0.033**	-0.023**
	(0.014)	(0.014)			(0.016)	(0.020)	(0.015)	(0.011)
Bank crisis	0.068	0.062			0.129*	0.148**	0.026	0.061
	(0.075)	(0.074)			(0.068)	(0.064)	(0.129)	(0.075)
Education	-0.282***	-0.285***			-0.295***	-0.269***	-0.212**	-0.250**
	(0.099)	(0.098)			(0.101)	(0.097)	(0.085)	(0.096)
Law and order				-0.042***				
				(0.010)				
Transparency				-0.050***				
				(0.010)				
Left power			-0.007					
			(0.067)					
Nonfinancial freedom			-0.048***					
			(0.015)					
Observations	241	241	220	204	123	586	89	241
<u>First stage results</u>								
Supervisory power	-0.345***	-0.351***	-0.416***	-0.383***	-0.218**	-0.491***	-0.182**	-0.374***
	(0.120)	(0.121)	(0.116)	(0.110)	(0.098)	(0.154)	(0.092)	(0.117)
R-squared	0.347	0.345	0.247	0.261	0.186	0.395	0.190	0.311
F-statistic (p-value)	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

Table VI. Bank regulations and the distribution of income

The table reports coefficient estimates and robust standard errors (in parentheses) of the estimation of Equation (1) over the period 1997-2005, using three-year averages. The sample includes information from 56 countries, excluding advanced economies for which data on the dependent variables are generally unavailable. The dependent variable is on the first line below this note. The explanatory variables are defined in Table I. All regressions are estimated using 2SLS with country fixed effects and time effects. The instrumental variable is *Supervisory power*, as defined in Table I. The ***, **, and * marks denote statistical significance at the 1%, 5%, and 10% level, respectively.

	(1) Income share held by highest 10%	(2) Income share held by highest 10%	(3) Income share held by lowest 10%	(4) Income share held by lowest 10%	(5) Poverty gap at $1.25 a day	(6) Poverty gap at $1.25 a day
Bank regulations	0.185	0.196*	0.603***	0.531***	-0.097*	-0.092*
	(0.116)	(0.115)	(0.177)	(0.191)	(0.050)	(0.054)
Population		-0.149*		-0.178**		0.139**
		(0.082)		(0.081)		(0.068)
GDP per capita		0.264***		0.110***		-0.148*
		(0.084)		(0.029)		(0.076)
Trade openness		0.099		0.128		0.096
		(0.109)		(0.102)		(0.134)
Government expenditure		0.126		0.066		0.091
		(0.133)		(0.125)		(0.086)
Inflation		-0.001		-0.019***		0.020***
		(0.005)		(0.004)		(0.005)
Bank liquidity		0.288***		0.208**		-0.303**
		(0.083)		(0.085)		(0.132)
Bank crisis		-0.102		-0.216**		0.109*
		(0.067)		(0.092)		(0.064)
Education		0.284**		0.396***		-0.450***
		(0.117)		(0.133)		(0.132)
Observations	121	101	111	93	120	100
First stage results						
Supervisory power	-0.344***	-0.421***	-0.292**	-0.302**	-0.350***	-0.340**
	(0.101)	(0.122)	(0.140)	(0.138)	(0.114)	(0.161)
R-squared	0.053	0.089	0.077	0.104	0.060	0.093
F-statistic (p-value)	0.126	0.000	0.004	0.000	0.082	0.000

Table VII. Different types of bank regulations and income inequality: Gini coefficient regressions

The table reports coefficient estimates and robust standard errors (in parentheses) of the estimation of Equation (1) over the period 1997-2005 using three-year averages. The main independent variable for each regression is on the first line below this note and defined in Table I. The explanatory variables are also defined in Table I. All regressions are estimated by 2SLS with country fixed effects and time effects. The instrumental variable is *Supervisory power*, as defined in Table I. The ***, **, and * marks denote statistical significance at the 1%, 5%, and 10% level, respectively.

	(1) Credit controls	(2) Interest rate controls	(3) Entry barriers	(4) Banking supervision	(5) Privatization	(6) International capital flows	(7) Securities markets
Bank regulations	-0.055**	-0.092***	-0.038*	-0.050**	-0.069***	-0.034*	0.038**
	(0.024)	(0.032)	(0.021)	(0.020)	(0.026)	(0.019)	(0.017)
Population	-0.118***	-0.111***	-0.109***	-0.114***	-0.100**	-0.111***	-0.115***
	(0.038)	(0.037)	(0.040)	(0.035)	(0.041)	(0.038)	(0.036)
GDP per capita	-0.070***	-0.076***	-0.075***	-0.055***	-0.074***	-0.076***	-0.065***
	(0.016)	(0.016)	(0.016)	(0.018)	(0.017)	(0.018)	(0.020)
Trade openness	-0.003	-0.004	0.003	0.001	0.010	0.004	0.001
	(0.014)	(0.015)	(0.015)	(0.002)	(0.015)	(0.013)	(0.010)
Government expenditure	0.021	0.023	0.021	0.015	0.018	0.019	0.019
	(0.038)	(0.035)	(0.035)	(0.047)	(0.047)	(0.057)	(0.056)
Inflation	0.006**	0.004*	0.006**	0.005**	0.010***	0.004*	0.005**
	(0.003)	(0.002)	(0.003)	(0.002)	(0.004)	(0.002)	(0.002)
Bank liquidity	-0.155***	-0.151***	-0.156***	-0.142***	-0.116***	-0.145***	-0.143***
	(0.032)	(0.031)	(0.032)	(0.032)	(0.044)	(0.033)	(0.033)
Bank crisis	0.067	0.063	0.066	0.070	0.058	0.064	0.077
	(0.074)	(0.071)	(0.074)	(0.076)	(0.070)	(0.070)	(0.082)
Education	-0.267**	-0.280**	-0.283***	-0.283**	-0.291***	-0.275**	-0.280**
	(0.107)	(0.109)	(0.106)	(0.109)	(0.104)	(0.118)	(0.111)
Observations	241	241	241	241	241	241	241
First stage results							
Supervisory power	-0.263***	-0.210**	-0.286***	-0.408***	-0.291**	-0.307**	-0.310***
	(0.094)	(0.101)	(0.081)	(0.125)	(0.120)	(0.150)	(0.106)
R-squared	0.285	0.263	0.271	0.300	0.225	0.209	0.268
F-statistic (p-value)	0.000	0.000	0.000	0.000	0.000	0.000	0.000