The Inequality-Credit Nexus

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Abstract

This paper explores the inequality-credit nexus from both a theoretical and an empirical perspective. The paper develops an overlapping generation model in which the effect of income inequality on private credit depends on the countries' per capita income and on the quality of laws protecting creditor rights. The model predicts that greater inequality leads to higher levels of private credit in countries with low per capita incomes and weak legal rights, while this effect is ambiguous or negative in economies with higher aggregate income and stronger credit protection. Using a panel dataset of 155 countries over the 1982 to 2015 period, the paper shows empirical evidence that is robust and consistent with the model's predictions. The paper's major finding suggests a credit channel through which inequality may affect economic outcomes.

JEL classification: F34; G15; G21; G38.

Keywords: Financial development, Inequality, Legal rights, Private credit

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1 Introduction

How does income inequality affect private credit? In light of the ample evidence of a link between finance and economic growth, this is a fundamental question in economics. Despite a rich body of literature examining the effects of income inequality, the study of a credit channel through which inequality may retard or accelerate economic growth remains in its infancy. An early strand of the theoretical literature suggests that inequality promotes growth by channelling resources towards wealthy individuals whose marginal propensity to save is higher (Lewis, 1954; Kaldor, 1957; Bourguignon, 1981). Modern theory suggests that, in the presence of imperfect credit markets, inequality adversely affects growth due to the negative effect on investment in human capital (Galor and Zeira, 1993). Recent empirical articles have started to explore the relationship between inequality and credit. However, these studies have focused mainly on either advanced economies or the recent periods of financial distress. For example, Bordo and Meissner (2012) find no evidence of a positive relationship between increased income inequality and credit booms in a sample of 14 advanced economies from 1920 to 2008.

This paper contributes to the literature on the effects of inequality by examining, from both a theoretical and an empirical perspective, whether the relationship between inequality and private credit is shaped by countries' per capita income level and strength of creditor rights. We develop an overlapping generations model with income inequality and endogenous credit constraints to explore this relationship, using a modified version of the model in Matsuyama (2004). Firms produce a single final good using labour and physical capital. Factor markets are competitive, and, thus, both wages and physical capital price evolve endogenously according to the dynamics of capital formation. Individuals live for two periods. In the first period young agents are endowed with heterogeneous units of labour. They work and allocate their income to finance old-age consumption. At the end of the first period, they have two options: first, they can deposit the savings to earn the international interest rate and consume their savings when old; second, they can ask for a loan—if they satisfy the minimum credit requirement—and invest in a risky project that produces physical capital in period 2 that can be sold to firms to finance consumption.

Credit market imperfections arise from the fact that banks are not perfectly protected against default. There are two reasons for default: entrepreneurs do not repay their debt because they can abscond with the loan and use it for consumption (as in Burkart and Ellingsen, 2004, which we denote ex ante—or malicious—default), or because their project fails (denoted ex post default). In both cases, only a fraction of the loan is recovered by the legal system and repaid to banks. Agents with sufficient own capital need smaller loans, and are therefore less likely to default maliciously, given the fraction of the loan that is recovered. There exists an agent that has a minimum amount of capital that makes him indifferent between defaulting maliciously and using the loan to invest in the project. Income inequality determines the mass of agents that satisfies the minimum credit requirement, and hence the agents with access to private credit. An alternative explanation for the relationship between inequality and credit expansion is provided by Kumhof et al. (2015). These authors suggest that an autonomous increase in inequality leads high income households with a wealth preference to increase lending to middle and lower income households, eventually resulting in a crisis.

We study the dynamics of the model and its steady state, which are very similar to those in Matsuyama (2004). With these results we study the effect on the path of capital accumulation of changes in the institutional quality of the credit market, i.e., the ex ante and the ex post default parameters. Both the severity of credit constraints and income inequality determine the mass of agents with access to the credit market, and, thus, the level of capital formation in the next period. Improvements in the financial market institutions accelerate the path of capital accumulation by increasing the mass of firms with access to credit.

The main prediction of the model is that a more unequal income distribution in an economy in which credit constraints are binding (i.e., an economy with low aggregate income and/or weak creditor protection) leads to higher credit penetration and aggregate debt (and thus an upwards shift in the dynamic path of the economy) while this effect is ambiguous and eventually reversed in economies with higher aggregate income and stronger credit protection.

We empirically test the main prediction of our model by using a panel dataset that covers 155 countries from 1982 to 2015. Consistent with the model's prediction, our analysis suggests that within-country increases in income inequality lead to a higher ratio of private credit to GDP in economies with low incomes and weak legal rights, while this effect vanishes and even becomes negative in economies with higher incomes and stronger legal rights. These results are statistically significant even after controlling for the standard determinants of private credit. Moreover, they are robust to the inclusion of country and time fixed effects; to alternative measures of income inequality; to alternative sub-samples of countries; and to control for potential endogeneity problems.

The estimated magnitudes of the effects obtained from our baseline regressions are economically meaningful. For example, an increase of one standard deviation in the Gini index for economies with a GDP per capita and a strength of legal rights index in the 25th percentile increases private credit to GDP by 513 basis points. Meanwhile, in economies with GDP per capita and a strength of legal rights index in the 75th percentile, an increase of one standard deviation in the Gini index decreases private credit to GDP by 527 basis points.

The study of a causal interpretation between inequality and growth is not trivial due to endogeneity biases generally stemming from potential omitted variables and reverse causality (Demirgüç-Kunt and Levine, 2009). Although the relationship between inequality and credit can be interpreted as capturing the effect of inequality on credit, there is evidence suggesting that credit may also improve the distribution of wealth and income (Beck et al., 2007). We aim to mitigate potential endogeneity concerns by estimating panel models with country and time fixed effects, and by using instrumental variables estimations. We instrument inequality with the residual variation in inequality that is not due to private credit to GDP (Brueckner and Lederman, 2015) and with three measures of ethnic, linguistic and religious fractionalization (Alesina et al., 2003; Braun et al., 2018). Additionally, we instrument GDP per capita applying the aforementioned statistical approach used for inequality. Finally, it is important to highlight that the sign change in the effect of inequality on credit, which we test empirically, is consistent with our theoretical framework and is not only a data-driven correlation.

This study contributes to our understanding of the inequality-finance nexus in at least three ways. First, this article suggests that arguments in favor of a positive relationship between inequality and private credit are incomplete, as we find conditions under which increases in inequality may lead to ambiguous and negative effects on credit. Our analysis helps to explain the mixed evidence reported in international empirical studies that estimate the average effect of income inequality on private credit (see, e.g. Bordo and Meissner, 2012; Perugini et al., 2015). Second, this article introduces a novel channel (the credit channel) through which inequality may affect economic outcomes. While the inequality-growth and finance-growth relationships have received considerable attention in the literature, the study of the mechanisms through which inequality affects financial development and inclusion are less understood. A better understanding of the channels in the finance-growth relationship is important given the mixed evidence in the literature. Although a number of empirical studies suggest that more inequality reduces economic growth (Alesina and Rodrik, 1994; Persson and Tabellini, 1994; Perotti, 1996; Panizza, 2002; Easterly, 2007), others support a positive effect of inequality on growth (Li and Zou, 1998; Forbes, 2000). Finally, in contrast to studies that focus on the U.S. or high-income OECD economies, this article utilizes a dataset covering a broader range of countries to gain a better understanding of a more general relationship between inequality and private credit under different economic conditions.

According to Demirgüç-Kunt and Levine (2009) there are few studies on the influence of formal financial policies, such as bank regulations, on the relation between access to credit and inequality.[1] Banerjee and Newman (1993) show the importance of capital market imperfections for the relation between inequality and access to credit, and how they affect the productive structure of the economy. Similarly, Galor and Zeira (1993) show that capital market imperfections interact with income (or wealth) distributions to alter the future distribution and accumulation of human capital. Greenwood and Jovanovic (1990) is an early dynamic model analyzing the relationship between endogenous growth and financial intermediation and in which the evolution of the income distribution can be analyzed. Matsuyama (2004) studies a dynamic model in which capital market imperfections interact with initial inequality to limit the access of agents to a productive technology. In a static model, Balmaceda and Fischer (2010) study the effect of changes in formal financial policies (bankruptcy and ex ante credit protection) on the interaction between inequality and access to credit and thus on the performance of an economy.

The paper proceeds as follows. Section 2 introduces our theoretical model. Section 3 describes the data and reports the summary statistics for the overall sample. Section 4 presents our empirical strategy and main results on the effect of inequality on private credit. Section 5 reports a set of additional robustness checks. Section 6 concludes.

2 The Model

We develop a Diamond overlapping generation model along the lines of Matsuyama (2004), as it provides a tractable framework in which to analyze the effects of income inequality in a dynamic model. Agents live for two periods. Each generation is assumed to receive earnings only when young. Thus, when young, agents work and at the end of the first period they allocate their income to finance old age consumption. They have two alternatives: i) they may deposit their earnings in the banking system and receive the fixed international gross rate per unit or ii) they may apply for a loan in order to invest in a risky project that produces physical capital that can be sold to firms. Capital cannot be consumed and fully depreciates in one period. Firms produce a single final good taking as input labour and physical capital supplied by the young and old generations respectively.

Each period, a continuum of young, risk neutral agents indexed by $z \in [0,1]$ is born, endowed with l^z units of observable labour.[2] Labor can be understood as human capital or labor capacity. Labour is distributed according to the cumulative probability function $\Gamma(l^z)$ with support $[0, l_{max}]$ and mean L^e

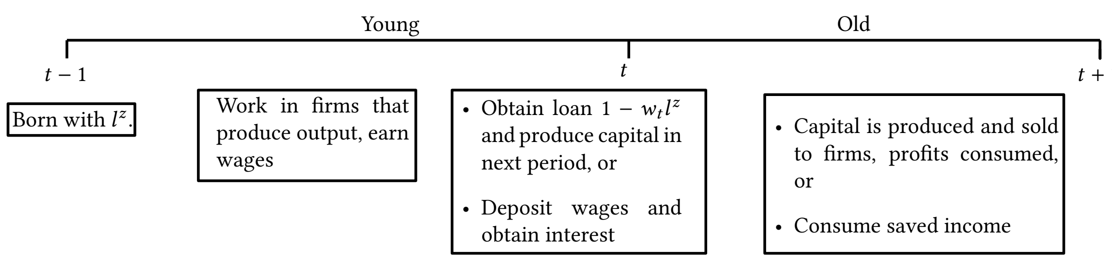
[1] A branch of the literature also focuses on the effects of financial reforms on income inequality (see,e.g., Bandiera et al., 2000; Christopoulos and McAdam, 2017).

[2] Risk neutrality is assumed for simplicity. In addition, we normalize the discount factor to one. All results still hold with a concave utility function and a discount factor lower than one.

(and density function $\gamma(l^z)$) which is also the aggregate labor supply each period. Earnings of agent z are given by $w_t l^z$, where w_t is the equilibrium wage in period t. Thus, $\Gamma(l^z)$ is the fraction of agents who earn less than $w_t l^z$. We normalize $L^e = 1$ to simplify notation. Therefore, the mean income of the distribution is w_t.

At the end of the first period t, young agents have two options. They can deposit their saved wages in the banking system and receive the international interest rate $(1+\rho^*)$. Then their second-period consumption is $(1+\rho)w_t l^z$. The second option is to apply for a loan to invest in a risky project that transforms output into physical capital that can be sold at a price p_{t+1} to firms at the beginning of the second period. See figure 1 for a timeline of the model. Lenders face moral hazard (*ex-ante* default) and bankruptcy (*ex-post* default) risk. The legal-institutional system provides imperfect protection against those risks. Thus, banks impose credit constraints that exclude poorer agents from the credit market and charge differentiated interest rates.

Figure 1: Lifecycle of agent z



As in Bernanke and Gertler (1989) there are different production technologies for output and capital. The capital production technology requires one unit of output. If the project succeeds, with independent probability θ, $\kappa > 1$ units of physical capital are produced. If the projects fails, the legal system recovers only a fraction $\upsilon \in [0,1]$ of the total investment. Thus, υ is a measure of the quality of the bankruptcy system, or alternatively, it can be interpreted as the liquidation value of the project. Lenders appropriate the entire liquidation value in the case of failure. Total physical capital produced in the economy determines total production of the output good. The technology of the consumption good is a standard linear homogeneous production function $Y_t = F(K_t, L_t)$, where L_t and K_t are aggregate labor and capital respectively. If there is a fraction η_t of young agents investing in a capital producing project at t, then total expected output at $t+1$ will be given by $F(\theta\kappa \cdot \eta_t, L_{t+1})$, where η_t measures credit penetration—i.e., the fraction of young agents who obtain a loan to invest in a capital-producing project at t.

We may then write the production function in per capita terms as $f(k_t)$, which satisfies $f'(k_t) > 0$, $f''(k_t) < 0$, $f(0) = 0$, $f''(0) = \infty$ and where $k_t \equiv \theta\kappa \cdot \eta_{t-1}$. The factor markets are competitive, thus the price of capital is $p_t = f'(k_t)$, and wages are $w_t = f(k_t) - k_t f'(k_t)$. As in Matsuyama (2004), we assume that $w(\theta\kappa) l_{max} < 1$—i.e. all agents need to borrow to start a project.

The future expected profit of a young agent z who becomes an entrepreneur at period t and obtains a loan $D_t^z = 1 - w_t l^z$ is:

$$\pi_{t+1}^z = \theta\kappa p_{t+1} - (1 + r_t^z)(1 - w_t l^z), \tag{1}$$

where r_t^z is the interest rate charged by lenders to entrepreneur z, which differs among agents.

Entrepreneurs who obtain a loan may decide to abscond instead of investing. In the second period they enjoy a private benefit from absconding given by $A(\phi, D_t^z)$ where ϕ measures the degree to which collateral laws protect creditors. We make the reasonable assumption that $A_1 < 0$, $A_2 > 0$ and $A_{22} < 0$, —i.e., better collateral protection (measured by ϕ) reduces the benefit from absconding, while the benefits of absconding increase in the size of the loan, at a decreasing rate. [3] Note that $A(\cdot)$ incorporates collateral laws in a very general sense.

The optimal contract must satisfy the incentive compatibility condition (IC) of agents, or the bank will not make a loan:

$$\pi_{t+1}^z \geq A(\phi, 1 - w_t l^z), \quad \forall l^z \geq \hat{l}_t(\phi, v), \tag{2}$$

where $w_t \cdot \hat{l}_t(\phi, v)$ is the minimum income needed to access the credit market at t. Therefore, $\Gamma(\hat{l}_t)$ is the measure of young agents who cannot access the credit market.

We assume that banks are competitive and have unlimited access to international funds at the gross rate $1+\rho^*$. A bank's expected future profits from lending to a young agent z who satisfies the IC constraint are:[4]

$$\Pi_{t+1}^z = [\theta(1 + r_{t+1}^z) - (1 + \rho^*)](1 - w_t l^z) + (1 - \theta) v. \tag{3}$$

2.1 Equilibrium

Since banks are competitive, $\Pi_{t+1}^z = 0$, the interest rate charged to an entrepreneur z in the second period is:

$$(1 + r_t^z) = \frac{1 + \rho^*}{\theta} - \frac{(1 - \theta) v}{\theta(1 - w_t l^z)}, \tag{4}$$

and substituting this into the profit function of entrepreneurs leads to $\pi_{t+1}^z = \theta\kappa p_{t+1} + (1 - \theta) v - (1 + \rho^*)(1 - w_t \hat{l}_t)$. Young agents become entrepreneurs only if the expected return is higher than for de-

[3] This means that the legal system is more efficient in pursuing the assets of agents who abscond with larger loans.

[4] Banks use the expected value, even though they can differentiate among agents, due to the fact that the shock θ is independent of individual income and by the Law of Large Numbers (since we have a continuum of agents).

positing their savings in the banking system—i.e., only if the following participation constraint (PC) is satisfied:

$$\theta\kappa p_{t+1} + (1-\theta)v \geq (1+\rho^*), \tag{5}$$

The fact that banks lend only to agents who do not abscond determines the minimum labour endowment required for a loan $\hat{l}_t$ at t:

$$\theta\kappa p_{t+1} + (1-\theta)v - (1+\rho^*)(1-w_t\hat{l}_t) - A(\phi, 1-w_t\hat{l}_t) = 0. \tag{6}$$

Finally, in order to ensure that agents can always repay their debt in the case of success, we impose the condition that the capital producing project be sufficiently productive.

Assumption 1. $\theta\kappa f'(\theta\kappa) \geq (1+\rho^*)[1-w(\theta\kappa)\hat{l}(\theta\kappa)] - (1-\theta)v$

Under Assumption 1, the agent $\hat{l}_t$ who has the largest loan can always repay his debt in the case of success.[5]

2.2 Dynamics and Steady State

We assume that the initial capital per capita is $k_0 > 0$. The dynamics of this model are completely described by the capital formation curve

$$k_{t+1}(k_t) = \begin{cases} \Upsilon(k_t) & \text{if } k_t < \hat{k}(\phi, v) \\ f'^{-1}\left(\frac{(1+\rho^*)-(1-\theta)v}{\theta\kappa}\right) & \text{if } k_t \geq \hat{k}(\phi, v), \end{cases} \tag{7}$$

where $k_{t+1} = \Upsilon(k_t)$ arises from the unique solution to $k_{t+1} = \theta\kappa[1-\Gamma(\hat{l}(k_t, k_{t+1})]$ and $\hat{l}(k_t, k_{t+1})$ is the solution to equation (6).

The critical per capita level at which capital formation remains constant $\hat{k}(\phi, v)$ is defined implicitly by the solution to $\theta\kappa p(\Upsilon(\hat{k}(\phi, v))) + (1-\theta)v = (1+\rho^*)$. The participation constraint (5) starts to bind if $k_t \geq \hat{k}(\phi, v)$, while the incentive constraint binds below this value. Thus, to the left of $\hat{k}(\phi, v)$, investment is constrained by borrowing restrictions, while to the right of this value the effects of credit market imperfections vanish. Note that the capital formation curve starts above the diagonal—i.e.,

$$k_{t+1}(0) = f'^{-1}\left(\frac{(1+\rho^*) + A(\phi, 1) - (1-\theta)v}{\theta\kappa}\right) > 0,$$

[5] Here $\hat{l}(\theta\kappa)$ is the solution to (6) when all agents have access to credit and the return to entrepreneurs (the price of capital) is lowest. In the assumption, the economy has the highest possible level of capital per worker: $k_t = \theta\kappa$.

because $w = 0$ when $k = 0$. This fact has implications for the dynamics of the model.

The critical value $\hat{k}(\phi, v)$ decreases as legal protection measures ϕ, v increase.[6] Therefore, as the legal framework for credit improves, the incentive compatibility constraint becomes less restrictive and less capital is necessary to have access to credit. Since the parameter ϕ helps to determine which agents obtain loans to produce next period's capital, it determines the growth rate of the economy.

The capital formation curve is increasing in k_t for $k_t < \hat{k}(\phi, v)$. In fact, differentiating this condition with respect to k_t we obtain:

$$\frac{\partial k_{t+1}}{\partial k_t} = -\frac{k_t \hat{l}_t f''(k_t)}{\frac{w_t}{\theta\kappa\gamma(\hat{l}_t)} - \frac{\theta\kappa f''(k_{t+1})}{(1+\rho^* + A_2(\phi, 1-w_t\hat{l}_t))}} > 0. \tag{8}$$

This is explained by a credit multiplier effect: higher domestic investment increases the wages of young agents, which allows a larger fraction of them to invest in the capital producing project for the next period. This, in turn, leads to higher future wages and to a higher proportion of young agents having access to credit.[7] From equation (8), it is straightforward to see that the minimum labour required for a loan decreases as capital per capita increases $\frac{\partial \hat{l}_t}{\partial k_t} < 0$.

2.3 Financial Market Improvements and Income Redistribution Effects

In this section, we use the model to derive the results that we test empirically in the second part of the paper. We begin with the comparative statics of improvements in institutional variables such as credit protection, measured by ϕ and the efficiency of bankruptcy procedures, measured by v.

Lemma 1. *An improvement in the quality of laws that protect creditors or in the efficiency of bankruptcy procedures leads to a loosening of credit constraints,—i.e, $\hat{l}_t$ shifts left.*[8]

Proof. See the Appendix. □

Thus, we deduce that $\frac{\partial k_{t+1}}{\partial \phi} > 0, \frac{\partial k_{t+1}}{\partial v} > 0$. That is, a better legal framework for loans accelerates capital formation through credit channels.

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[6] Since $\frac{\partial \hat{k}(\phi,v)}{\partial x} = -\underbrace{\frac{\partial \Upsilon}{\partial x}}_{>0} / \underbrace{\frac{\partial \Upsilon}{\partial k}}_{>0} < 0, x = \phi, v$..

[7] A well-known issue of these types of models is that they can have multiple steady states, even without credit market imperfections (e.g. see Acemoglu, 2008, p. 332). The dynamics of the model are very similar to those of Matsuyama (2004, section 7.1). Since $\Upsilon(0) > 0$, there exist at least one crossing with the 45° line from above, that is, there is at least one stable steady state. See Matsuyama (2004) for the detailed analysis of the dynamics of this type of model.

[8] Note that at t, k_t is fixed because it was decided by young agents at $t-1$.

As a second topic, we examine the effects of income redistribution at t on credit penetration and on total debt. In the next proposition we describe the effect of an income redistribution of young agents on credit penetration and aggregate debt of the next period.

Recall that credit penetration is defined as the percentage of agents that can access the credit market at t, $\eta_t \equiv 1 - \Gamma(\hat{l}_t)$. We define aggregate debt as follows:

$$\mathscr{D}_t = \int_{\hat{l}_t}^{l_{max}} (1 - w_t l^z) \partial\Gamma(l^z). \tag{9}$$

In order to isolate income distribution from income effects, we consider pure income redistributions—i.e., Mean Preserving Spreads (MPS).[9] Since an MPS preserves the mean, total income among the young $\int w_t l^z \partial\Gamma(l^z) = w_t$ remains constant after an MPS of their incomes. In our setting, redistribution among the young at the beginning of period t does not change wages at t but will impact outcomes at $t+1$ through changes in credit access at t. Moreover, these effects are persistent.

Proposition 1. *Consider two countries,1 and 2, that are identical in all respects, except that the young of country 1 have an income distribution at t that is an MPS of that of country 2. If $w_t \hat{l}_t(\phi, \nu) >> w_t$, credit penetration and aggregate debt is higher in country 1 at t. The results are reversed if $w_t \hat{l}_t(\phi, \nu) << w_t$.*

Proof. See the Appendix. □

The proposition shows that the effects of inequality depend on the relationship between the minimum wage income required for a loan and the current average income of the young. In economies in which a young agent with average labor capacity does not have the wage income necessary to access credit (*credit-constrained countries*: $w_t \hat{l}_t >> w_t$), a regressive redistribution among the young leads to higher credit penetration and debt. A regressive redistribution has the opposite effect in a country in which a young agent with average labor capacity has access to credit, ($w_t \hat{l}_t << w_t$). Conversely, progressive redistribution reverses these effects in both credit constrained and credit-unconstrained countries. For countries with intermediate degrees of credit constraints, the direction of the effect of inequality on credit is ambiguous.

Note that a change in income distribution among the young has two effects: it changes the mass of agents that have access to the credit market, and it changes the future cost of capital in the opposite direction due to the adjustment in the future supply of capital. This second effect modifies period t credit requirements measured by $\hat{l}_t$. The directions of these effects depend on whether a country is

[9] A Mean Preserving Spread can be characterized by: Γ_1 is an MPS of Γ_0 if $\Gamma_1(l^z) > \Gamma_0(l^z)$ when $w_t l^z < w_t$ and $\Gamma_1(l^z) < \Gamma_0(l^z)$ when $w_t l^z > w_t$.

credit-constrained or credit-unconstrained. Equation (13) in the Appendix displays these two opposite effects.

Figure 2: Effects of an MPS in a wealthy country

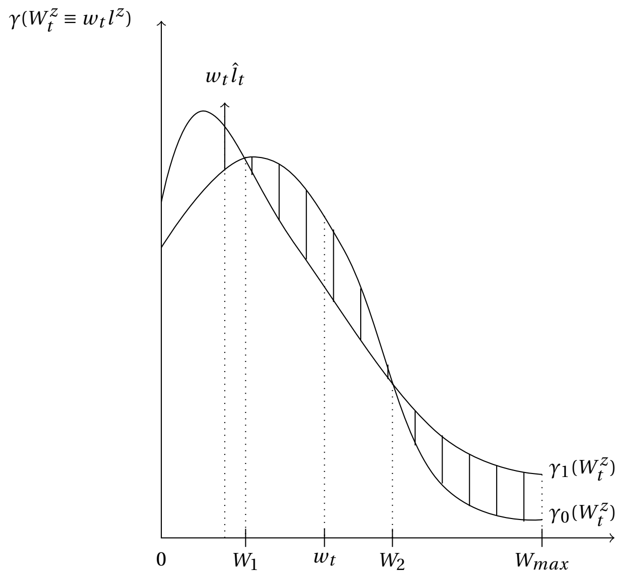

Figure 2 shows the effects of an MPS in a high income country (note that $w_t \hat{l}_t << w_t$). We can see from the figure that there are fewer agents with access to credit, as the mass of excluded agents, to the left of $w_t \hat{l}_t$, is larger. Moreover, since there are fewer agents in the range W_1, W_2 (who ask for larger than average loans), and there is more mass among the high income agents (which ask for smaller-than-average loans), total loan volume decreases.[10] A similar figure can be drawn to show that the effects of an MPS in a poor economy would increase not only access to credit but also the total volume of debt.[11]

It is important to emphasize that the direction of the effect of any MPS is clear as long as the country does not move from an initial state in which $w_t \hat{l}_t > w_t$ to a final one in which $w_t \hat{l}_t < w_t$ or viceversa. In an economy with an intermediate income level—i.e. where $w_t \hat{l}_t(\phi, \nu)$ is close enough to the average income—a change in the relationship between $w_t \hat{l}_t$ and w_t is more likely to occur. In fact, it is possible that an MPS can move $\hat{l}_t$ so that the economy jumps from being unconstrained to becoming constrained, or viceversa, leading to an ambiguous effect of an MPS on credit. For instance, consider an intermediate income country which initially satisfies that $w_t \hat{l}_t > w_t$ and where $\hat{l}_t$ is close to 1. An MPS that reduces inequality will initially decrease credit penetration, increasing the future price of capital, and therefore

[10] This figure is simplified, since there is also a small movement in $\hat{l}_t$ in response to the MPS.

[11] The effect of an MPS results from the combined effect of the change in the mass of agents with access to credit and the change in the mass of agents with a given income. Even when there is no change in credit penetration, a redistribution of income may have an effect on aggregate debt. The change in the extensive market is not a requirement for our results.

reducing $\hat{l}_t$. If after this change the country becomes unconstrained ($w_t\hat{l}_t < w_t$), the MPS will also increase credit to some extent. Overall, the net effect of the MPS on credit will not be clear. A similar analysis applies for an intermediate income economy such that $w_t\hat{l}_t < w_t$.

Finally, note that if we consider two countries with the same income but different levels of legal rights, we could have that in the country with better legal rights, $w_t\hat{l}_t << w_t$. Thus, inequality reduces credit penetration and total credit. On the other hand, in the country with weaker institutions, it could be that $w_t\hat{l}_t >> w_t$, and, thus, increased inequality has the opposite effect. Note that there are ambiguous effects of inequality on private credit when countries have intermediate levels of both wealth and legal rights. This is consistent with the empirical results in section 4.

Apart from the short-run effects of a transitory increase in inequality, we can distinguish two kinds of long-run effects. First, the transitory changes in income distribution will persist through a credit multiplier type effect. The initial effect of increased inequality is to have higher credit penetration at period t, which raises capital formation and output at $t+1$. In turn, this raises the wages of the young at $t+1$. The young, face loosened credit constraints, thus increasing the number of agents who invest in the capital production technology, and these effects continue until the model converges to the steady state. Similar effects operate in the opposite direction when there is increased inequality in a high-income country.

Second, there can be jumps from the basin of attraction of one steady state to the basin of attraction of a higher stable steady state when there are multiple steady states. Conversely, more inequality in a high-income country can have a detrimental effect with, potentially, a shift to a lower steady state. Thus, transitory changes in the income distribution can have persistent effects along the entire transition path, and even lead to different equilibria in the long-run.

2.4 Mapping to Data

From Lemma 1, we know that the minimum income to obtain a loan is smaller in countries with better laws to protect creditors. Thus, a credit-constrained country can be understood as an economy in which aggregate income (w_t) is low relative to the minimum income required to receive credit ($w_t\hat{l}_t$), where the threshold $\hat{l}_t$ depends on the quality of institutions that protect creditors (ϕ, V).

In the empirical analysis aggregate income is measured by GDP per-capita, while the pair (ϕ, V) is measured by the strength of legal rights according to the World Bank's strength of legal rights index (SLR), which measures the degree to which collateral and bankruptcy laws protect lenders. In the analysis, we test the effect of inequality on credit penetration of the next period; that is, we test the effect presented in Proposition 1.

3 Data

To empirically test the predictions of Proposition 1, we employ a large panel of 155 emerging and advanced economies over the years 1982 to 2015, which we construct by using data from the World Bank's World Development Indicators (WDI), the Standardized World Income Inequality Database (SWIID) by Solt (2009) and other sources. Our dependent variable is the level of private credit by deposit money banks and other financial institutions as a fraction of GDP. This measure has been used mainly as a continuous proxy for the size of domestic private credit (see, e.g., Gennaioli et al., 2014).

In line with our model, our key explanatory variables of interest are income inequality, average income, and the quality of collateral laws and of bankruptcy legislation (the pair (ϕ, υ) in the model). We measure income inequality by using the Gini index from both the WDI and the SWIID, and the income share held by the top 10% from the WDI. According to Leigh (2007), the top 10% income share represents a suitable alternative proxy to measure broader income inequality.[12] We proxy for a country's average income using GDP per capita and for the quality of creditor protection using the strength of legal rights (SLR) index. The SLR index measures the degree to which collateral and bankruptcy laws protect the rights of borrowers and lenders. The SLR index ranges from 0 to 12. Higher scores indicate that these laws are better designed to expand access to credit.[13]

For robustness purposes, we also control in some specifications for a large set of variables that may directly affect private credit: abundance of natural resources, inflation, the share of GDP generated by the manufacturing sector, economic growth, bank concentration and financial openness. These same variables have been used to study the links between financial development and growth (Levine, 2005) and the determinants of financial development and inclusion (Allen et al., 2014; Fischer and Valenzuela, 2013). Finally, we employ a number of variables as instruments for income inequality and per capita GDP, average income. We instrument inequality and GDP per capita with the residual variation in inequality that is not due to private credit to GDP. Following Braun et al. (2018), we also instrument inequality with three measures of ethnic, linguistic and religious fractionalization.[14]

Table 1 reports summary statistics for the overall sample. The Appendix presents the description and sources of all the variables (Table A.1) and provides the list of countries included in our sample (Table A.2).

[12]We also check the robustness of our results by using the income share held by the top 20%, the lowest 10%, and the lowest 20%.

[13]The data on SLR index start only in 2004. Our baseline estimations reported in Table 2 utilizes the earliest value available. This procedure may be satisfactory because the SLR index exhibits substantial persistence over time.

[14]Braun et al. (2018) instrument inequality by using fractionalization from Alesina et al. (2003) to explore the effect of income inequality on economic growth and innovation.

4 Empirical Strategy and Main Results

In this section, we perform panel data regression analyses to check whether the data support Proposition 1 of our model. That is, we test whether the effect of inequality on private credit depends on the levels of GDP per capita and the strength of legal rights. Thus, our baseline econometric model takes the following form:

$$\begin{aligned} Private\ Credit_{i,t} &= \alpha_i + \nu_t + \beta_1 Inequality_{i,t-1} + \beta_2 GDPpc_{i,t-1} + \beta_3 Legal\ Rights_{i,t-1} \\ &\quad \beta_4 Inequality_{i,t-1} \times GDPpc_{i,t-1} + \beta_5 Inequality_{i,t-1} \times LegalRights_{i,t-1} + \epsilon_{i,t}, \end{aligned} \tag{10}$$

where $Private\ Credit_{i,t}$ is the level of private credit by deposit money banks and other financial institutions as a fraction of GDP in country i at time t; $Inequality_{i,t-1}$ is the lagged value of either the Gini index or the income share held by the top 10%; $GDPpc_{i,t-1}$ is the lagged value of GDP per capita; and $Legal\ Rights_{i,t-1}$ is the lagged value of the SLR index. We include the two interaction terms to examine whether, consistent with Proposition 1, the direction of the effect of inequality on private credit depends on the capital constraints of individual countries. In this context, a credit constrained economy (see section 2.4) is a country with low GDP per capita and weak Legal Rights, while a country that exhibits higher GDP per capita and stronger SLR index is less constrained or unconstrained. We lag all of our independent variables to attenuate potential biases associated with reverse causality.

To attenuate potential endogeneity biases associated with omitted variables, all of our regressions include country and time fixed effects. The term α_i represents country fixed effects that control for all time-invariant country-specific factors affecting both private credit and inequality. The term ν_t captures time fixed effects that control for common shocks affecting all countries such as global financial crises or changes in the world business cycle.

According to the model presented in equation (10), the effect of income inequality on private credit at different levels of GDP per capita and legal rights can be calculated by examining the following partial derivatives:

$$\frac{\partial PrivateCredit_{i,t}}{\partial Inequality_{i,t-1}} = \beta_1 + \beta_4 GDPpc_{i,t-1} + \beta_5 LegalRights_{i,t-1}. \tag{11}$$

Based on Proposition 1, we expect that $\beta_1 > 0$, $\beta_4 < 0$ and $\beta_5 < 0$. In other words, greater within-country income inequality leads to higher private credit in economies with low GDP per capita and weak legal rights, but this effect vanishes and even becomes negative in economies with higher GDP per capita and stronger legal rights. Note that the direction of the effect of income inequality on private credit depends on the levels of β_1, β_4 and β_5 as well as on the levels of GDP per capita and the SLR index.

Table 2 reports our main results from estimating various specifications of equation (10) using ordinary least squares (OLS) regressions with standard errors clustered by country. In columns 1 to 3, we measure income inequality by using the Gini index from WDI. In columns 4 to 6, we measure income inequality with the income share held by the top 10% from WDI. In columns 7 to 9, we measure income inequality by using the Gini index from SWIID. In some columns—1, 2, 4, 5, 7 and 8—we consider only one of the two interaction terms at a time (setting $\beta_4 = 0$ or $\beta_5 = 0$). In columns 3, 6 and 9, we simultaneously include the two interaction terms of interest.

As expected, our three measures of income inequality (the two Gini indexes and the income share held by the top 10%), GDP per capita, and the SLR index enter with positive and statistically highly significant coefficients in almost all of our regressions. Moreover, in line with Proposition 1, the interaction term between inequality and GDP per capita and the interaction between inequality and the SLR index enter in our regressions with negative coefficients that are also in general highly statistically significant.[15] Overall, the significant positive coefficient on inequality and the negative coefficients on the interaction terms confirm that greater within-country income inequality leads to higher private credit in capital-constrained economies, but that this effect vanishes and even becomes negative in capital-unconstrained economies. As reported in columns 3, 6 and 9, our results are robust to simultaneously including the two interaction terms. Although our baseline regressions are very parsimonious, it is important to highlight that, due to the inclusion of country and time fixed effects, they are able to explain a very large proportion of the variance in private credit to GDP, as shown by the adjusted R-squared.

Figures 3 and 6 show that the effect of income inequality on private credit is heterogeneous across credit constrained economies and credit unconstrained economies. Specifically, Figures 3 and 4 show graphically the marginal effect of the two measures of inequality on private credit to GDP conditional on values of the natural logarithm of GDP per capita.[16] The figures report 95% confidence bands and the number of countries by GDP per capita. Consistent with the predictions of our theoretical model, Figure 3 clearly shows that the marginal effect of the Gini index on private credit to GDP is positive and statistically significant in low income economies, not significant in middle and high income economies, and negative and statistically significant in a few very high income economies. Countries like India, Chile and Norway are in the low, middle and high, and very high income groups, respectively. Thus, our results emphasize that increased income inequality may have distinct effects on aggregate credit in economies

[15] In unreported regressions we also checked the robustness of our results to alternative measures of inequality (i.e., income share held by the top 20%, the lowest 10%, and the lowest 20%). The results remains qualitatively unchanged.

[16] We conduct this exercise using the results reported in columns 1 and 4 of Table 2.

with relatively low levels of inequality but with different levels of per capita income (e.g., India versus Norway). Figure 4 shows that the marginal effect of the 10% top income share on private credit to GDP is positive and statistically significant in low income economies, while it is not significant in middle and high income economies. The results shown in Figure 4 are in line with previous studies that do not find a significant relationship between private credit and income inequality in high income economies. In a sample of 14 advanced economies from 1920 to 2008, Bordo and Meissner (2012) reject any relationship between private credit and the income share held by the top 1%, 5% and 10%.[17]

Figures 5 and 6 show the marginal effect of the previously mentioned measures on inequality on private credit to GDP conditional on values of the SLR index. These figures also report 95% confidence bands and the number of countries by SLR. Figures 5 and 6 show that the marginal effect of inequality on private credit to GDP is positive and statistically significant in economies with weak legal rights, not significant in economies with strong legal rights, and negative in a few economies with the strongest levels of legal rights (i.e., economies with a SRL index of 10). Countries like Venezuela, South Africa and United Kingdom are in the weak, strong, and strongest legal rights groups, respectively.

5 Robustness

In this section, we test whether our main empirical findings are robust to a set of additional robustness checks, including using instrumental variables estimations, different sub-samples of countries and a comprehensive set of country-level time-variant control variables. In what follows, to keep the size of the paper manageable, we only report the estimations by using the inequality measures from the World Bank.[18]

5.1 Instrumental Variables Estimation

Although country and time fixed effects mitigate potential endogeneity concerns associated with omitted variables, they do not correct for endogeneity biases associated with reverse causality. This is an important concern, given a potential effect running from private credit to inequality. Demirgüç-Kunt and Levine (2009) emphasize that improvements in domestic financial markets are likely to expand eco-

[17]The countries analyzed by Bordo and Meissner (2012) are: Australia, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Spain, Sweden, Switzerland, United Kingdom, and the United States.

[18]Although the sample size increases considerably when we use the index from SWIID, we report all our results by using the Gini index from WDI as there are some issues associated with the imputation model employed by the SWIID (Jenkins, 2015). In any case, the main tables of the paper were also replicated using the Gini index from SWIID. The findings remain largely unchanged.

nomic opportunities and, thus, reduce inequality.

To attenuate potential endogeneity associated with reverse causality, we employ instrumental variables estimations. Given the lack of suitable country-level time-varying instruments to address potential endogeneity in our context,[19] we use these estimations as a robustness test—one that needs to be interpreted with caution. Specifically, to clean the potential effect of private credit to GDP on inequality, we follow two different strategies. First, we construct inequality and GDP per capita variables that contain all of the dimensions of inequality and income that are unrelated to expansions of private credit to GDP. For this purpose, we follow an instrumental variables strategy similar to that of Brückner (2013) and Brueckner and Lederman (2015). We do so by first regressing inequality and GDP per capita (in natural logarithm) on private credit to GDP (together with country and time fixed effects) and by then computing the residual variation in inequality and GDP per capita that is not due to private credit to GDP. [20] Brückner (2013) utilizes this instrumental variables strategy to explore the effect of foreign aid on economic growth, while Brueckner and Lederman (2015) use it to examine the effect of inequality on output. Second, following Braun et al. (2018), we instrument inequality using measures of ethnic, linguistic and religious fractionalization provided by Alesina et al. (2003). This is consistent with the idea that fractionalization may accentuate between-groups differences and conflicts, leading to persistent national inequality.

Table 3 reports our results from re-estimating our baseline models using instrumental variables regressions. We instrument the Gini index and GDP per capita using the residual variation of each variable that is not due to private credit to GDP. The main empirical findings remain qualitatively unchanged. Additionally, in Table A.3 from the Appendix, we instrument the Gini index and 10% top income shares with the previously mentioned measures of fractionalization. Given that these instruments are time-invariant, we do not consider country fixed effects. Once again, our previous results remain qualitatively unchanged, although the interpretation is somewhat different. That is, income inequality is related to higher ratios of private credit to GDP in economies with low incomes and weak legal rights, while this effect disappears and even becomes negative in economies with higher incomes and stronger legal rights. [21]

[19] Perugini et al. (2015) employ labour market regulation and economic integration as instruments for inequality. However, these instruments are unlikely to satisfy the exclusion restriction as they directly affect private credit. Stronger labor regulation reduces credit capacity of firms (Balmaceda and Fischer, 2010). It is also well documented that financial integration affects corporate credit risk and domestic private credit (Fischer and Valenzuela, 2013; Andreasen and Valenzuela, 2016).

[20] Blanchard and Perotti (2002) and Fatás and Mihov (2003) employ an equivalent strategy to construct instruments in the literature on fiscal policy.

[21] In unreported estimations, first-stage regressions report that these instruments are strong ($F >> 10$). The instruments coefficients are, in general, consistent with proposed fractionalization theory and also statistically significant.

5.2 Sub-Samples

It is likely that our main results on the relationship between income inequality and private credit are driven by periods of financial distress. To address this concern, Table 4 explores the robustness of our findings once these periods of financial distress are excluded from our sample. Specifically, we merge our database with the one from Laeven and Valencia (2013) and drop all country-year observations in which a country experienced a systemic banking crises. Overall, our main results remain largely qualitatively unchanged.

Table 5 replicates our baseline regressions (columns 1 to 3 of Table 2) for different sub-samples of countries.[22] Columns 1 to 6 separate our sample into two groups of countries based on income level: 1) low- and lower-middle income economies; and 2) upper-middle and high-income economies. These estimations offer an additional robustness test to our baseline regressions because they focus on more homogeneous samples of countries in terms of economic development. Once again, our previous results are robust to the mentioned sub-sample of countries, as the signs on the coefficients on all our variables of interest are in the expected direction, and most of them are statistically highly significant. The loss of significance on a few coefficients is expected, given the smaller sizes of the samples. Regarding the magnitude of the effects, the results reported in columns 1 and 4 of Table 5 indicate, for example, that the marginal effect of the Gini index on private credit for a country with a GDP per capita (in ln) of 5 is equal to 0.271 (i.e., $1.876 - 0.321 \times 5$), while that for a country with a GDP per capita of 9 is equal to -0.038 (i.e., $3.355 - 0.377 \times 9$).

The regressions in columns 7 to 9 exclude African countries, while those in columns 10 to 12 exclude Latin American Countries. It is well-documented that some African and LAC countries are characterized by capital markets highly constrained and/or by very high levels of inequality. Therefore, these estimations offer a robustness test to the private credit regressions estimated using the whole sample because they ensure that potential outliers are not driving our main results. Most coefficients of interest remain statistically highly significant in the expected directions. Additionally, finding similar results for subsets of countries with different levels of income is in line with the possibility of multiple steady states in the theoretical model.

[22] For space considerations, we only present the results obtained when we measure inequality with the Gini index. The main findings remain qualitatively unchanged if we use the 10% top income share to measure inequality.

5.3 Additional Control Variables

As a final test, Table 6 reports the results from estimating our baseline regression, controlling for a comprehensive set of variables that are commonly accepted as determinants of the private credit (see e.g Allen et al., 2014; Fischer and Valenzuela, 2013). These variables include abundance of natural resources; inflation; the share of GDP generated by the manufacturing sector; economic growth; banking concentration and the Chinn-Ito financial openness index (Chinn and Ito, 2006). All of our previous main findings remain qualitatively unchanged. Despite the fact that the sample size drops dramatically from 1,341 to 739 observations (i.e., a 45% drop).

6 Conclusions

Although rich bodies of research and economic history highlight a close relationship between inequality and credit, whether inequality improves or weakens access to credit remains an open question. In this paper, we explore this relationship from both a theoretical and an empirical perspective. We develop an overlapping generation model in which the effect of income inequality on private credit depends on the countries' per capita income and on the quality of laws protecting creditor rights, and we provide empirical evidence that is consistent with the main prediction of the model.

In our model, credit market imperfections arise from the fact that banks are not perfectly protected against default. In the case of default, only a fraction of the loan is recovered by the legal system and repaid to banks. A credit-constrained country is understood as an economy in which initial aggregate income is low relative to the quality of institutions that protect creditors. In this framework, both the severity of credit constraints and income inequality determine the mass of agents with access to the credit market and, thus, capital formation in the next period. The general lesson from our model is that a more-unequal income distribution in an economy with binding credit constraints (because of low aggregate income and/or weak creditor protection) leads to higher credit penetration and aggregate debt. By contrast, a more-unequal income distribution in an economy in which credit constraints are not binding leads to smaller credit penetration and aggregate debt.

Using panel data models for a large number of developed and emerging economies over the last three decades, we show empirical evidence consistent with our theoretical arguments. Our empirical analysis suggests that within-country increases in income inequality lead to a higher ratio of private credit to GDP in economies with low incomes and weak legal rights, while this effect vanishes and even becomes negative in economies with higher incomes and stronger legal rights. The main empirical finding of our paper is robust to the inclusion of country and time fixed effects, to alternative measures of income inequality, to alternative sub-samples of countries, and to controlling for potential endogeneity problems.

Our analysis point towards a mechanism through which economic development and legal institutions may shape the impact of inequality on financial markets. In light of the ample evidence for a linkage between economic and financial development, we believe that our main finding also contributes to the growth literature by suggesting a novel channel (i.e., the credit channel) through which inequality may affect economic growth.

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Figure 3: Marginal effect of the Gini index on private credit to GDP conditional on the values of GDP per capita (in logs). The dotted lines are 95% confidence bands. This exercise was conducted using the results reported in column (1) of Table 2.

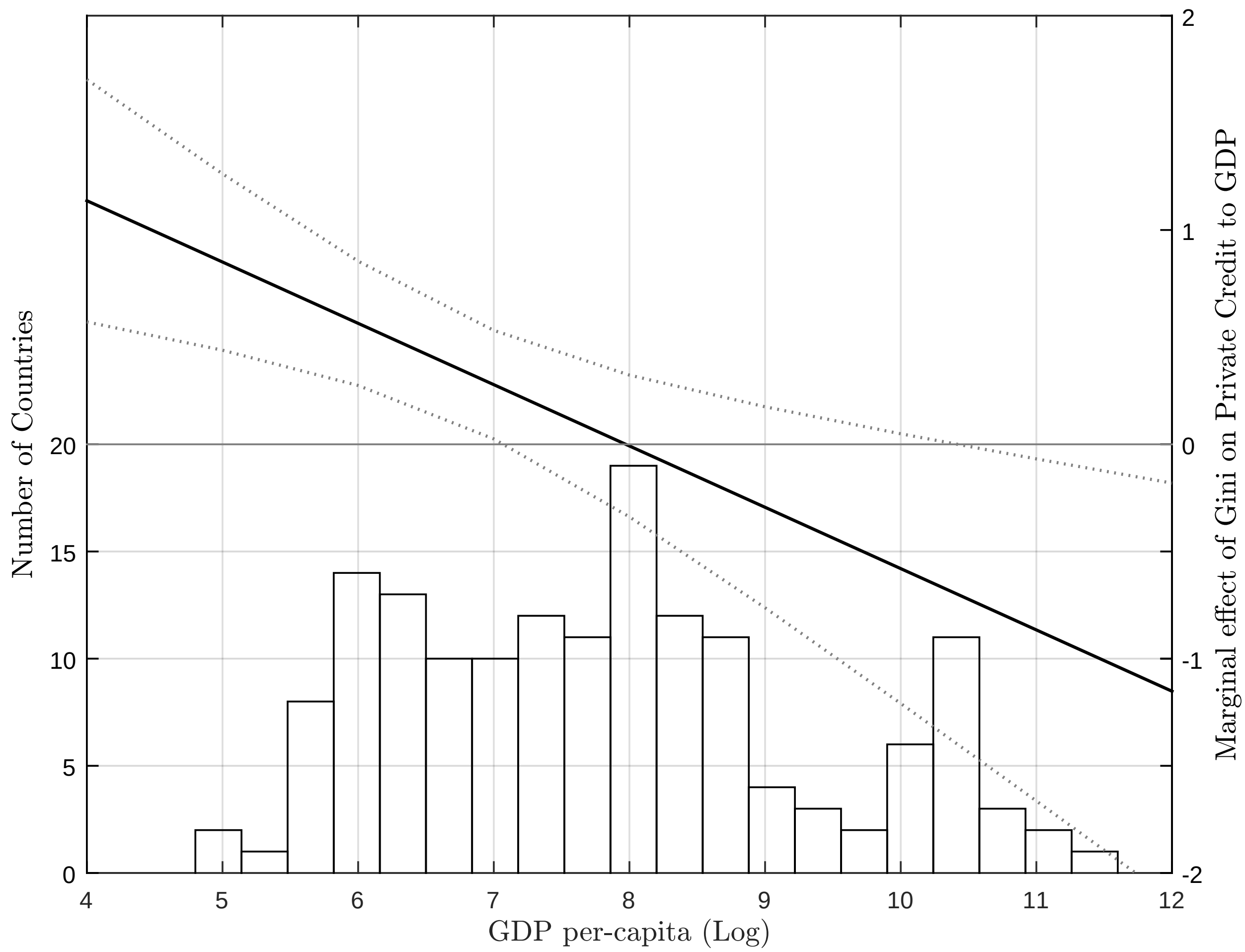

Figure 4: Marginal effect of 10% top income share on private credit to GDP conditional on the values of GDP per capita (in logs). The dotted lines are 95% confidence bands. This exercise was conducted using the results reported in column (4) of Table 2.

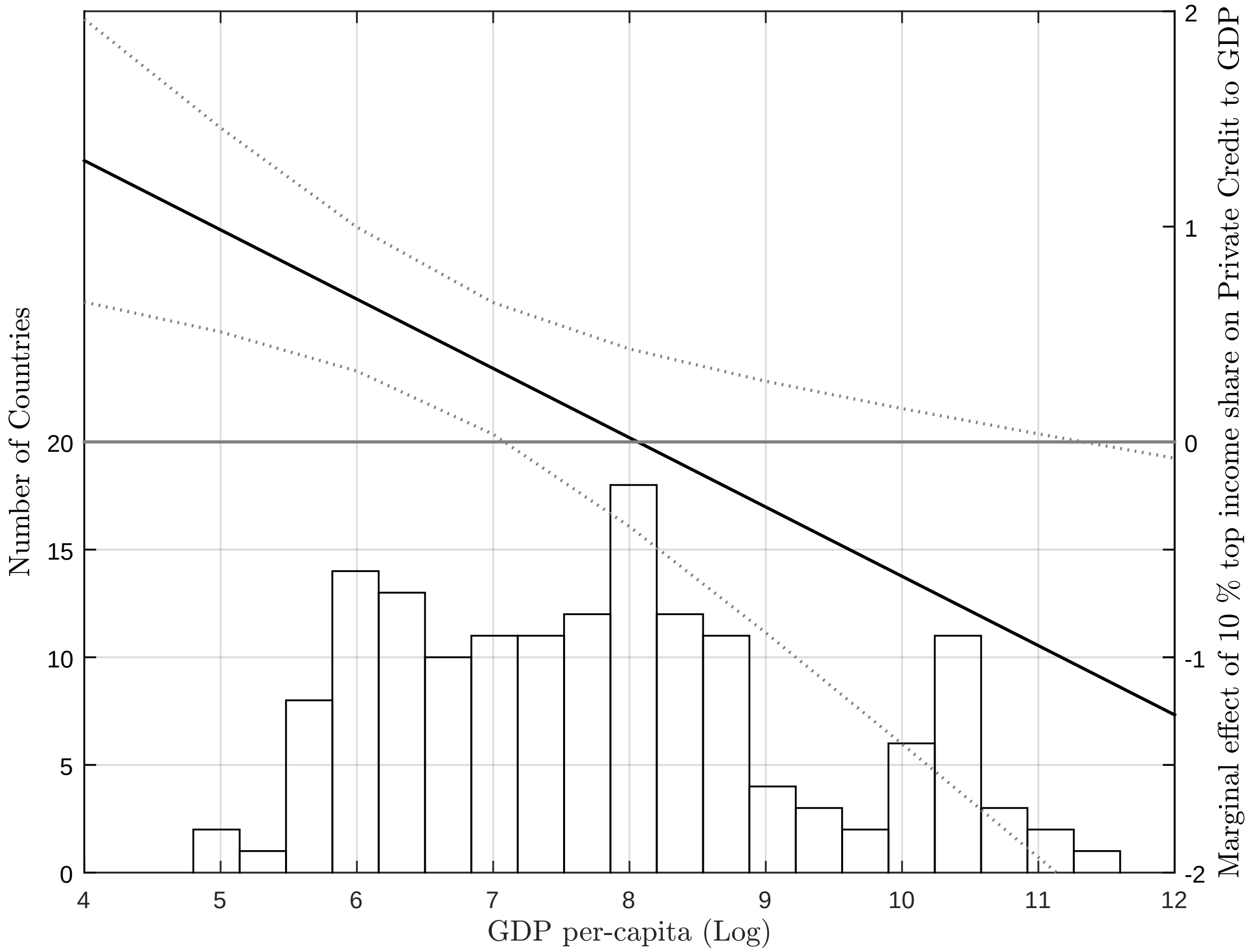

Figure 5: Marginal effect of the Gini index on private credit to GDP conditional on the values of the Legal Rights Index. The dotted lines are 95% confidence bands. This exercise was conducted using the results reported in column (2) of Table 2.

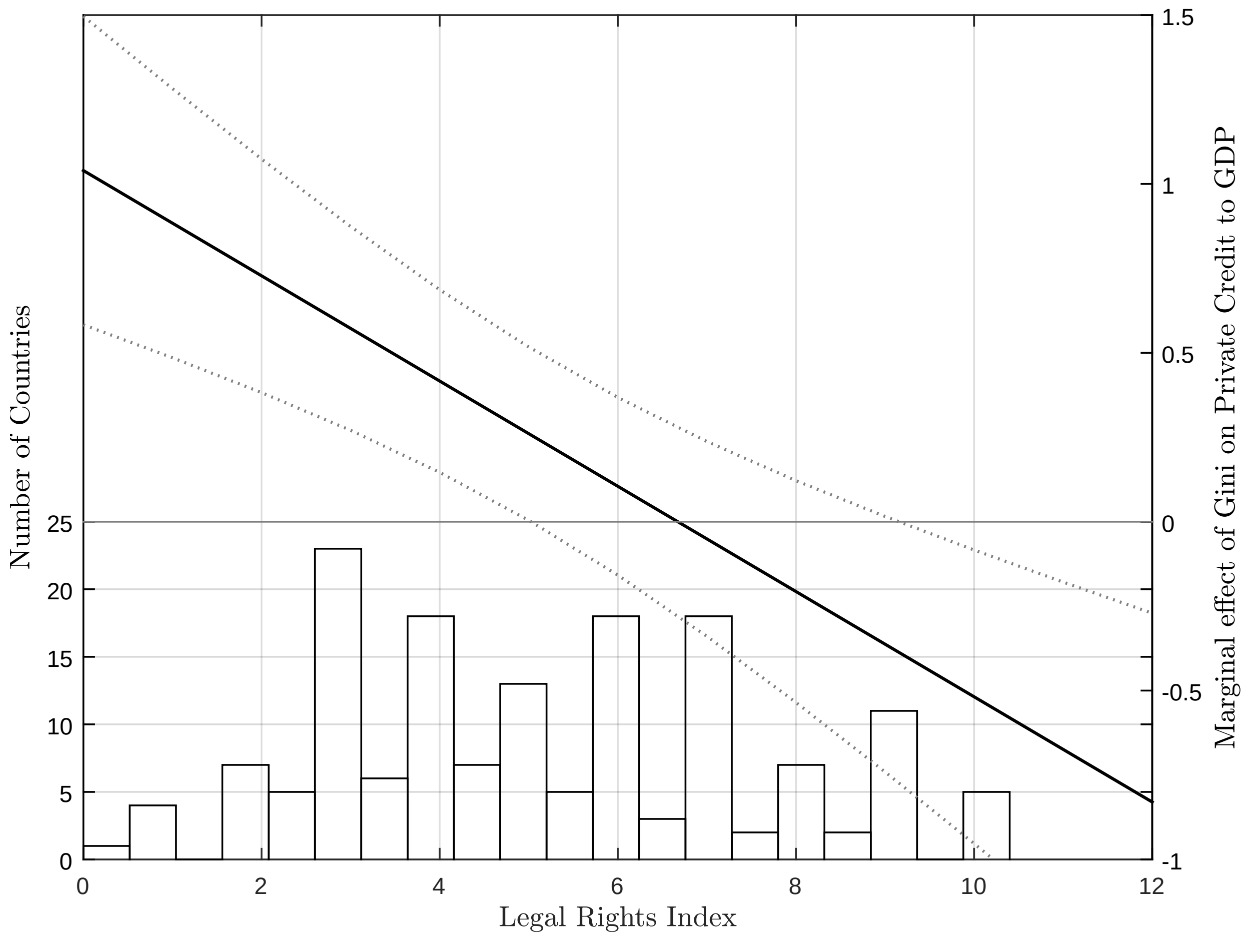

Figure 6: Marginal effect of the 10% top income share on private credit to GDP conditional on the values of the Legal Rights index. The dotted lines are 95% confidence bands. This exercise was conducted using the results reported in column (5) of Table 2.

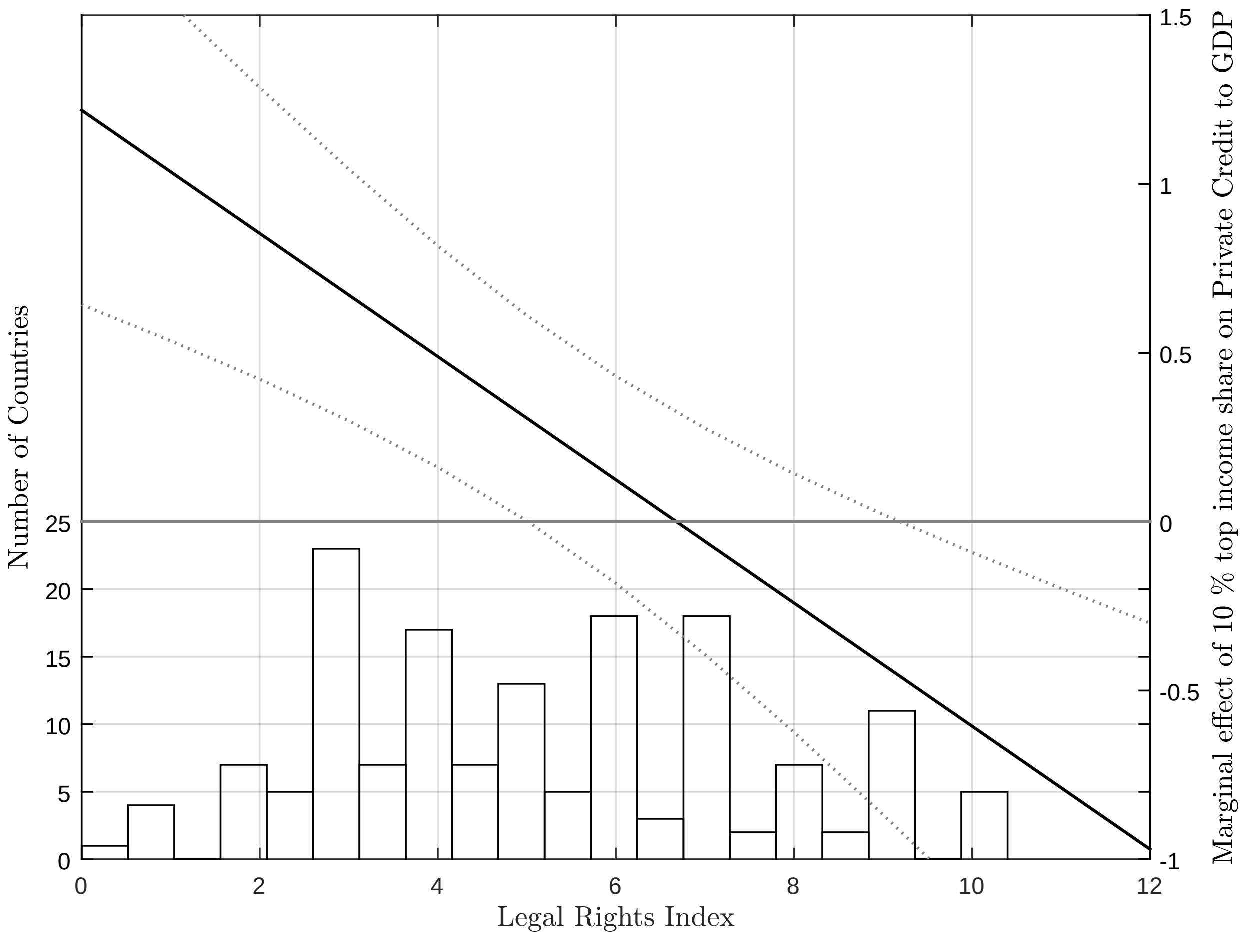

Table 1: Descriptive Statistics

VARIABLES	**Obs.**	**Mean**	**S.D.**	**Min.**	**Max.**
Private Credit / GDP	1341	50.45	44.87	0	312.2
Gini (WDI)	1341	39.72	9.978	16.23	74.33
10% top income share	1339	31.19	7.759	17.14	65
Gini (SWIID)	3853	38.85	8.576	19	63.30
Log(GDP per capita)	1341	8.245	1.514	4.782	11.64
Legal Rigths Index	1341	5.397	2.374	0	12
Natural resources	1340	6.503	10.20	0	71.90
Inflation	1317	23.39	179.3	-10.07	4735
Manufacturing/GDP	1238	16.36	6.348	0.382	43.54
Growth GDP per capita	1335	2.710	4.511	-29.14	30.34
Chinn-Ito index	1271	0.565	1.562	-1.904	2.374
Bank Concentration	798	67.31	18.88	21.40	100
Ethnic fractionalization	1296	0.404	0.234	0	0.930
Linguistic fractionalization	1258	0.340	0.268	0.00800	0.923
Religious fractionalization	1290	0.398	0.214	0.00230	0.860

Table 2: Inequality, Capital Constraints and Private Credit

Private credit to GDP	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**
Gini (WDI)	2.281***	1.040***	2.607***						
	(0.640)	(0.233)	(0.641)						
10% top income share				2.594***	1.219***	3.036***			
				(0.758)	(0.294)	(0.756)			
Gini (SWIID)							5.821***	1.754***	5.935***
							(0.521)	(0.286)	(0.553)
Log(GDP per capita)	25.87***	15.16***	24.12***	24.84***	15.30***	23.44***	36.36***	13.64***	36.12***
	(4.158)	(2.251)	(4.193)	(3.853)	(2.250)	(3.868)	(3.115)	(1.111)	(3.107)
Legal Rights Index	1.434***	7.569***	6.528***	1.411***	7.125***	6.350***	0.372	3.557*	1.592
	(0.464)	(1.688)	(1.696)	(0.465)	(1.668)	(1.672)	(0.348)	(2.133)	(2.137)
Gini (WDI) x Log(GDP per capita)	-0.286***		-0.239**						
	(0.0922)		(0.0933)						
Gini (WDI) x Legal Rights Index		-0.156***	-0.128***						
		(0.0372)	(0.0376)						
10% top income share x Log(GDP per capita)				-0.322***		-0.275**			
				(0.111)		(0.111)			
10% top income share x Legal Rights Index					-0.182***	-0.156***			
					(0.0463)	(0.0467)			
Gini (SWIID) x Log(GDP per capita)							-0.624***		-0.619***
							(0.0753)		(0.0749)
Gini (SWIID) x Legal Rights Index								-0.0833*	-0.0306
								(0.0500)	(0.0501)
Observations	1,341	1,341	1,341	1,342	1,342	1,342	3,853	3,853	3,853
Adjusted R-squared	0.871	0.871	0.872	0.871	0.871	0.872	0.858	0.854	0.858
Country fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES
Time fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES

Table 3: Instrumental Variables Estimation

Private credit to GDP	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10)**	**(11)**	**(12)**
Gini (WDI)	4.594***	0.999***	4.979***				1.839**	1.180***	2.180***			
	(0.630)	(0.220)	(0.636)				(0.757)	(0.262)	(0.760)			
10% top income share				5.579***	1.215***	6.161***				1.988**	1.372***	2.461***
				(0.791)	(0.286)	(0.794)				(0.870)	(0.322)	(0.867)
Log(GDP per capita)	41.20***	16.00***	38.91***	40.47***	16.08***	38.41***	-5.409	-12.83***	-6.930	-6.724	-12.90***	-7.866
	(4.057)	(2.094)	(4.096)	(3.938)	(2.091)	(3.927)	(5.697)	(3.275)	(5.740)	(5.260)	(3.294)	(5.281)
SLR index	1.515***	10.59***	7.767***	1.484***	10.50***	8.211***	1.932***	7.890***	7.213***	1.895***	7.541***	7.070***
	(0.430)	(1.631)	(1.608)	(0.434)	(1.680)	(1.621)	(0.517)	(1.735)	(1.748)	(0.515)	(1.688)	(1.714)
Gini (WDI) x Log(GDP per capita)	-0.669***		-0.608***				-0.201*		-0.152			
	(0.0901)		(0.0910)				(0.108)		(0.110)			
Gini (WDI) x SLR index		-0.232***	-0.156***					-0.150***	-0.132***			
		(0.0362)	(0.0360)					(0.0390)	(0.0394)			
10% top income share x Log(GDP per capita)				-0.818***		-0.750***				-0.213*		-0.165
				(0.116)		(0.115)				(0.126)		(0.127)
10% top income share x SLR index					-0.290***	-0.212***					-0.179***	-0.163***
					(0.0475)	(0.0461)					(0.0475)	(0.0487)
Observations	1,324	1,324	1,324	1,325	1,325	1,325	1,324	1,324	1,324	1,325	1,325	1,325
Adjusted R-squared	0.890	0.892	0.891	0.890	0.892	0.890	0.872	0.872	0.873	0.872	0.872	0.873
Instrument for Gini index: Statistic	YES	YES	YES	YES	YES	YES	NO	NO	NO	NO	NO	NO
Instrument for GDP per capita: Statistic	NO	NO	NO	NO	NO	NO	YES	YES	YES	YES	YES	YES
Country fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Time fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES

Table 4: Periods of Financial Stability

Private credit to GDP	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Gini (WDI)	1.732***	1.041***	2.010***			
	(0.625)	(0.229)	(0.624)			
10% top income share				1.961***	1.277***	2.358***
				(0.723)	(0.273)	(0.724)
Log(GDP per capita)	21.59***	14.04***	19.59***	20.79***	14.18***	19.06***
	(4.036)	(2.286)	(4.096)	(3.683)	(2.295)	(3.701)
Legal Rights Index	0.973**	6.843***	6.099***	0.958**	6.645***	6.075***
	(0.440)	(1.657)	(1.691)	(0.440)	(1.621)	(1.630)
Gini (WDI) x Log(GDP per capita)	-0.202**		-0.149*			
	(0.0887)		(0.0904)			
Gini (WDI) x Legal Rights Index		-0.147***	-0.127***			
		(0.0362)	(0.0371)			
10% top income share x Log(GDP per capita)				-0.222**		-0.166
				(0.104)		(0.104)
10% top income share x Legal Rights Index					-0.179***	-0.160***
					(0.0444)	(0.0446)
Observations	1,233	1,233	1,233	1,233	1,233	1,233
Adjusted R-squared	0.873	0.874	0.874	0.873	0.874	0.874
Country fixed effects	YES	YES	YES	YES	YES	YES
Time fixed effects	YES	YES	YES	YES	YES	YES

Table 5: Sub-Samples by Income Level and Region

	Low and Middle Lower Income			Upper Middle and High Income			Excluding Africa			Excluding LAC		
Private credit to GDP	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10)**	**(11)**	**(12)**
Gini (WDI)	1.876**	0.500**	1.737**	3.355***	1.200***	3.670***	2.817***	1.092***	3.091***	0.741	0.911***	1.179
	(0.728)	(0.241)	(0.706)	(1.076)	(0.401)	(1.077)	(0.834)	(0.280)	(0.823)	(0.782)	(0.328)	(0.819)
Log(GDP per capita)	30.02***	17.13***	25.51***	28.00***	14.68***	27.06***	27.45***	14.81***	25.45***	16.65***	13.77***	15.20***
	(5.144)	(2.577)	(5.158)	(5.902)	(3.010)	(5.938)	(4.932)	(2.638)	(5.036)	(5.136)	(2.928)	(5.156)
SLR index	2.720***	7.827***	6.534***	0.692	7.236**	5.629*	1.697***	7.915***	6.676***	1.722***	6.782***	6.621***
	(0.583)	(1.934)	(1.772)	(0.794)	(3.251)	(3.341)	(0.521)	(1.934)	(1.969)	(0.628)	(2.202)	(2.189)
Gini (WDI) x Log(GDP per capita)	-0.321***		-0.217*	-0.377***		-0.342**	-0.343***		-0.290**	-0.0864		-0.0413
	(0.115)		(0.111)	(0.139)		(0.141)	(0.114)		(0.116)	(0.116)		(0.117)
Gini (WDI) x SLR index		-0.133***	-0.0992**		-0.158**	-0.116		-0.162***	-0.127***		-0.141***	-0.136**
		(0.0418)	(0.0397)		(0.0707)	(0.0730)		(0.0438)	(0.0451)		(0.0545)	(0.0537)
Observations	435	435	435	906	906	906	1,144	1,144	1,144	985	985	985
Adjusted R-squared	0.822	0.823	0.825	0.878	0.877	0.878	0.880	0.880	0.881	0.904	0.905	0.905
Country fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Time fixed effects	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES

Table 6: Additional Control Variables

Private credit to GDP	**(1)**	**(2)**	**(3)**
Gini (WDI)	2.112**	0.596**	1.267
	(0.823)	(0.259)	(0.865)
Log(GDP per capita)	22.01***	9.746***	13.34**
	(4.878)	(2.825)	(5.372)
SLR index	1.081**	4.105**	3.831**
	(0.487)	(1.771)	(1.780)
Gini (WDI) x Log(GDP per capita)	-0.246**		-0.0949
	(0.116)		(0.123)
Gini (WDI) x SLR index		-0.0820**	-0.0737*
		(0.0409)	(0.0409)
Natural resources	-0.190*	-0.109	-0.118
	(0.102)	(0.0938)	(0.0946)
Inflation	0.105***	0.104***	0.104***
	(0.0244)	(0.0239)	(0.0234)
Manufacturing/GDP	-0.358	-0.455	-0.446
	(0.298)	(0.282)	(0.279)
Growth GDP per capita	-0.687***	-0.688***	-0.667***
	(0.162)	(0.160)	(0.159)
Bank Concentration	-0.103**	-0.0944**	-0.0928**
	(0.0479)	(0.0453)	(0.0452)
Chinn-Ito index	0.0155	0.177	0.174
	(0.826)	(0.821)	(0.821)
Observations	739	739	739
Adjusted R-squared	0.962	0.964	0.964
Country fixed effects	YES	YES	YES
Time fixed effects	YES	YES	YES

Appendix

A Additional Tables

Table A.1: Description of Variables

Variable	Description	Source
Private Credit / GDP	Domestic credit to private sector (% of GDP)	WDI (2015)
Gini (WDI)	GINI index (World Bank estimate)	WDI (2015)
Gini (SWIID)	GINI index (Standardized World Income Inequality Database)	Solt (2009)
10% top income share	Income share held by highest 10%	WDI (2015)
Log(GDP per capita)	Log of GDP per capita (constant 2010 US$)	WDI (2015)
SLR index	Index of collateral and bankruptcy laws protection of borrowers and lenders (0-12).	WDI (2015)
Natural resources	Total natural resources rents (% of GDP)	WDI (2015)
Inflation	Inflation, consumer prices (annual %)	WDI (2015)
Manufacturing/GDP	Manufacturing, value added (% of GDP)	WDI (2015)
Growth GDP per capita	GDP growth (annual %)	WDI (2015)
Bank Concentration	Assets of the three largest banks as a share of assets of all banks.	Financial Development and Structure (Beck et al., 2010)
Chinn-Ito index	Financial openness index.	Chinn and Ito (2006)
Ethnic fractionalization	Degree of ethnic fractionalization (0-1)	Alesina et al. (2003)
Linguistic fractionalization	Degree of linguistic fractionalization (0-1)	Alesina et al. (2003)
Religious fractionalization	Degree of religious fractionalization (0-1)	Alesina et al. (2003)

Table A.2: List of Countries in Sample

Africa

Algeria
Angola
Benin
Botswana
Burkina Faso
Burundi
Cabo Verde
Cameroon
Central African Republic
Chad
Comoros
Congo, Dem. Rep.
Congo, Rep.
Cote d'Ivoire
Djibouti
Egypt, Arab Rep.
Ethiopia
Gabon
Gambia, The
Ghana
Guinea
Guinea-Bissau
Kenya
Lesotho
Liberia
Madagascar
Malawi
Mali
Mauritania
Mauritius
Morocco
Mozambique
Namibia
Niger
Nigeria
Rwanda
Sao Tome and Principe
Senegal
Seychelles
Sierra Leone
South Africa
Sudan
Swaziland
Tanzania
Togo
Tunisia
Uganda
Zambia

Asia

Afghanistan
Bangladesh
Bhutan
Cambodia
China
India
Indonesia
Iran, Islamic Rep.
Iraq
Israel
Japan
Jordan
Kazakhstan
Kyrgyz Republic
Lao PDR
Malaysia
Maldives
Mongolia
Nepal
Pakistan
Philippines
Russian Federation
Sri Lanka
Syrian Arab Republic
Tajikistan
Thailand
Timor-Leste
Turkey
Vietnam
West Bank and Gaza
Yemen, Rep.

Europe

Albania
Armenia
Austria
Azerbaijan
Belarus
Belgium
Bosnia and Herzegovina
Bulgaria
Croatia
Cyprus
Czech Republic
Denmark
Estonia
Finland
France
Georgia
Germany
Greece
Hungary
Iceland
Ireland
Italy
Latvia
Lithuania
Luxembourg
Macedonia, FYR
Moldova
Montenegro
Netherlands
Norway
Poland
Portugal
Romania
Serbia
Slovak Republic
Slovenia
Spain
Sweden
Switzerland
Ukraine
United Kingdom

Latin America and the Caribbean

Argentina
Belize
Bolivia
Brazil
Chile
Colombia
Costa Rica
Dominican Republic
Ecuador
El Salvador
Guatemala
Guyana
Haiti
Honduras
Jamaica
Mexico
Nicaragua
Panama
Paraguay
Peru
St. Lucia
Suriname
Trinidad and Tobago
Uruguay
Venezuela, RB

Northern America

Canada
United States

Oceania

Australia
Fiji
Micronesia, Fed. Sts.
Papua New Guinea
Samoa
Solomon Islands
Tonga
Vanuatu

Table A.3: Instrumental Variables Estimation: Fractionalization

Private credit to GDP	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Gini (WDI)	2.335	0.842*	4.480***			
	(1.966)	(0.443)	(1.727)			
10% top income share				4.623*	1.086*	7.092***
				(2.604)	(0.596)	(2.318)
Log(GDP per capita)	31.02***	16.00***	32.05***	36.47***	15.50***	36.03***
	(8.414)	(0.775)	(6.753)	(8.865)	(0.823)	(7.278)
SLR index	2.794***	16.92***	15.41***	2.836***	17.88***	16.62***
	(0.458)	(4.252)	(3.818)	(0.455)	(4.496)	(4.060)
Gini (WDI) x Log(GDP per capita)	-0.396*		-0.442**			
	(0.232)		(0.187)			
Gini (WDI) x SLR index		-0.356***	-0.315***			
		(0.107)	(0.0955)			
10% top income share x Log(GDP per capita)				-0.698**		-0.709***
				(0.307)		(0.252)
10% top income share x SLR index					-0.489***	-0.444***
					(0.145)	(0.130)
Observations	1,254	1,254	1,254	1,254	1,254	1,254
Adjusted R-squared	0.468	0.433	0.456	0.471	0.419	0.453
Instrument for Inequality: Fractionalization	YES	YES	YES	YES	YES	YES
Country fixed effects	YES	YES	YES	YES	YES	YES
Time fixed effects	YES	YES	YES	YES	YES	YES

B Proofs

Lemma 1. *An improvement in the quality of laws that protect creditors or in the efficiency of bankruptcy procedures leads to a loosening of credit constraints,—i.e, $\hat{l}_t$ shifts left.*

Proof. From condition (6) and the fact that k_t is given at t we have:

$$\frac{\partial \hat{l}_t}{\partial \phi} = \frac{A_1(\phi, 1 - w_t \hat{l}_t)}{1 + \rho^* + A_2(\phi, 1 - w_t \hat{l}_t) - (\theta\kappa)^2 \gamma(\hat{l}_t) f''(k_{t+1})} < 0$$
$$\frac{\partial \hat{l}_t}{\partial v} = -\frac{(1-\theta)}{1 + \rho^* + A_2(\phi, 1 - w_t \hat{l}_t) - (\theta\kappa)^2 \gamma(\hat{l}_t) f''(k_{t+1})} < 0$$

□

Proposition 1. *Consider two countries,1 and 2, that are identical in all respects, except that the young of country 1 have an income distribution at t that is an MPS of that of country 2. If $w_t \hat{l}_t(\phi, v) >> w_t$, credit penetration and aggregate debt is higher in country 1 at t. The results are reversed if $w_t \hat{l}_t(\phi, v) << w_t$.*

Proof. We concentrate the proof on MPSs whose densities cross the original distribution only twice. It is easy to deal with the case of more crossings, but the proofs are more cumbersome. Since the average income remains unchanged after an MPS, we can obtain the effect of a labour redistribution on the interest variables and then scale the results in terms of income.

Let $\Gamma_1(l^z)$ be an MPS of $\Gamma_0(l^z)$ and define the convex combination as $\Gamma_\lambda = \lambda\Gamma_1 + (1-\lambda)\Gamma_0$. Differentiation of condition (6) with respect to λ leads to (income distribution of young agents at t does not impact w_t):

$$\frac{\partial \hat{l}_t}{\partial \lambda} = -(\theta\kappa)^2 \frac{f''(k_{t+1}) \frac{\partial \eta_t}{\partial \lambda}}{(1 + \rho + A_2) w_t} \tag{12}$$

Differentiation of credit penetration leads to:

$$\frac{\partial \eta_t}{\partial \lambda} = \Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) - \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} \tag{13}$$

Thus, replacing condition (12), we obtain:

$$\frac{\partial \eta_t}{\partial \lambda} = \frac{\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t)}{1 - (\theta\kappa)^2 \frac{f''(k_{t+1})\gamma(\hat{l}_t)}{(1+\rho+A_2) w_t}} \tag{14}$$

By definition of an MPS, if $\hat{l}_t > 1$, then $\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) > 0$. Since $f'' < 0$, the denominator of the RHS is positive, and, thus, $\frac{\partial \eta_t}{\partial \lambda} > 0$, and by substituting into (12), $\frac{\partial \hat{l}_t}{\partial \lambda} > 0$. With higher credit penetration, the future cost of capital falls $\frac{\partial p_{t+1}}{\partial \lambda} < 0$. Note that in the case of a reversed MPS (less inequality), we have $\frac{\partial \hat{l}_t}{\partial \lambda} > 0$, and it may be that $\hat{l}_t < 1$ for the economy with less inequality. In order to guarantee that this does not happen, we require that $\hat{l}_t >> 1$ (i.e., $w_t \hat{l}_t >> w_t$).

On the other hand, if $\hat{l}_t < 1$, then $\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) < 0$ and $\frac{\partial \eta_t}{\partial \lambda} < 0$, $\frac{\partial p_{t+1}}{\partial \lambda} > 0$ and $\frac{\partial \hat{l}_t}{\partial \lambda} < 0$. By the same arguments used in the paragraph above, we need to impose the condition $\hat{l}_t << 1$ to guarantee that the result is not overturned[23].

In the case of the result for debt, note that we can rewrite condition (9) as follows:

$$\mathscr{D}_t = (1 - \Gamma(\hat{l}_t)) - \int_{\hat{l}_t}^{l_{max}} w_t l^z \partial \Gamma(l^z) = (1 - \Gamma(\hat{l}_t)) + \int_0^{\hat{l}_t} w_t l^z \partial \Gamma(l^z) - w_t$$

Differentiating debt with respect to λ leads to:

$$\frac{\partial \mathscr{D}_t}{\partial \lambda} = \int_0^{\hat{l}_t} w_t l^z \partial(\Gamma_1 - \Gamma_0) + w_t \hat{l}_t \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} + \left(\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) - \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} \right) \tag{15}$$

Recall that we assumed that the MPS distributions cross at only two points l_1, l_2 and that $l_1 < 1 < l_2$. Assume first that $\hat{l}_t < l_1$. We have that $\frac{\partial \eta_t}{\partial \lambda} = \Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) - \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} < 0$, so the last term in (15) is negative. In addition, we showed that $\frac{\partial \hat{l}_t}{\partial \lambda} < 0$, so the middle term is also negative. However, the first term is positive. When $\hat{l}_t < l_1$, an upper bound for (15) is:

$$\begin{aligned}
\frac{\partial \mathscr{D}_t}{\partial \lambda} &< -w_t \hat{l}_t \left[\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) - \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} \right] + \left(\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) - \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} \right) \\
&= [1 - w_t \hat{l}_t] \left(\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) - \gamma(\hat{l}_t) \frac{\partial \hat{l}_t}{\partial \lambda} \right) \\
&= \underbrace{[1 - w_t \hat{l}_t]}_{>0} \underbrace{\frac{\partial \eta_t}{\partial \lambda}}_{<0} \\
&< 0
\end{aligned}$$

Thus, when $\hat{l}_t < l_1$, more inequality leads to lower aggregate debt. When $\hat{l}_t \in (1, l_2)$, we obtain a positive

[23]Note that it could be the case that $\hat{l}_t = 0$ or $\hat{l}_t \to l_{max}$. In these particular cases an MPS may not have any affect on η_t, because $\Gamma_0(\hat{l}_t) - \Gamma_1(\hat{l}_t) = 0$. Nevertheless, based on the proof we present below, there still would exist an effect on aggregate debt.

lower bound using the logic of the first case:

$$\frac{\partial \mathscr{D}_t}{\partial \lambda} > \underbrace{[1 - w_t \hat{l}_t]}_{>0} \underbrace{\frac{\partial \eta_t}{\partial \lambda}}_{>0} > 0$$

When $\hat{l}_t \in (l_1, 1)$ or when $\hat{l}_t > l_2$ it is straightforward to see that is $\frac{\partial \mathscr{D}_t}{\partial \lambda}$ negative or positive, respectively.

We conclude that $\frac{\partial \mathscr{D}_t}{\partial \lambda} > 0$ if $\hat{l}_t > 1$ and that $\frac{\partial \mathscr{D}_t}{\partial \lambda} < 0$ if $\hat{l}_t < 1$. Note, finally, that since w_t is unchanged by the value of λ, these results can be expressed more conveniently in terms of the minimum income required to have access to credit, $w_t \hat{l}_t$. Additionally, this proof applies to a numerable crossing points.

□