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Left-Wing Political Strength, Inclusive Institutions, and the Evolution of Capitalist Systems

Teng, Xingan

University College London

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Left-Wing Political Strength, Inclusive Institutions, and the Evolution of Capitalist Systems

Xingan Teng

Abstract

This paper examines how left-wing political strength shapes the evolution of capitalist systems through the lens of income compositional inequality (IFC). Using LIS microdata for nearly 40 countries from 1978–2022, I construct an unbalanced panel of IFC and estimate two-way fixed-effects models with Driscoll–Kraay standard errors, complemented by dynamic panel GMM and a fuzzy RDD around close elections. Results show that stronger left representation reduces IFC and pushes economies toward liberal capitalism; a 10-percentage-point increase in left strength lowers IFC by about 0.0079—roughly 7.5% of the sample mean. Political checks and balances attenuate this distributive effect, while rule-of-law and property-rights institutions amplify it. Channel analysis based on the pseudo-Gini of capital indicates that the main pathway operates via reductions in capital inequality. The findings highlight that "inclusive institutions" are internally heterogeneous and interact with partisan power, offering a more granular account of distributive dynamics within democracies.

Keywords: Income composition inequality; Left-wing strength; Inclusive institutions; Capitalist systems

JEL: D31; D33; D72; P16; P51

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1) Introduction

As capitalism expanded globally and gradually became the dominant system (Milanovic, 2019), significant differences in development emerged among capitalist countries. These differences not only shaped their economic and social trajectories (Hall & Soskice, 2001), but also led to divergence in institutional stability, income inequality, and political governance. This makes the study of heterogeneity within capitalism both urgent and relevant. Against this background, distributional issues have once again moved to the center of scholarly debate.

In *Capital in the Twenty-First Century*, Piketty (2014) links the rising capital share to the increase in global income inequality, and introduces the well-known inequality $r > g$. Building on this insight, Milanovic (2017), and Ranaldi and Milanovic (2022) propose a distinction between liberal capitalism and classical capitalism. Further, Ranaldi (2022) develops a methodological innovation by constructing an index of functional income distribution inequality, which provides a quantitative basis for this concept. Later studies show that most countries are moving toward liberal capitalism, but with different speeds and intensities (Ranaldi, 2025; Ranaldi & Milanovic, 2022; Berman & Milanovic, 2024). These works highlight the diversity of capitalism from the perspective of factor income inequality and provide stylized facts for further analysis. However, this distributive view is mostly static and descriptive, lacking integration with

institutional analysis and dynamic causal mechanisms. In other words, few studies have explained the deeper causes behind differences and changes in capitalist systems.

By contrast, theories of parties and the welfare state have long tried to explain the dynamic link between left-wing politics and income distribution. The “power resources theory” argues that social equality and welfare states emerge because the working class, organized through left-wing parties, shapes social policy (Korpi, 1978, 1998; Esping-Andersen, 1988). Later theories criticize power resources theory for neglecting the broader environment in which the left operates, and seek to refine it with more detailed perspectives (Pierson, 1996; Hall & Soskice, 2001; Estevez-Abe et al., 2001). Many scholars argue that parties, as “players” within institutional rules, influence distribution only in relation to the broader institutional context (Iversen & Soskice, 2001; Ansell, 2008).

However, in the literature on institutions, political checks and balances and economic protection of property rights are often seen as the root causes of cross-country differences in economic growth (Acemoglu et al., 2001, 2005; Acemoglu & Robinson, 2006). Later debates expanded the explanatory power of institutions (Acemoglu et al., 2005). For example, human capital accumulation, technological progress, and even income and wealth distribution are treated as outcomes of inclusive institutions (Acemoglu et al., 2005; Acemoglu & Robinson, 2012). Yet many empirical studies show

that the relationship between distribution and inclusive institutions is not as clear as the theory suggests (Acemoglu et al., 2015; Sokoloff & Engerman, 2000). In addition, some scholars argue that institutional determinism relies too much on a binary divide such as “democracy vs. authoritarianism” or “inclusive vs. extractive.” This leaves limited attention to differences in distributional structures among democratic capitalist countries, and to the role of party ideology in shaping economic distribution (Bawn et al., 2012; Iversen & Soskice, 2006). Parties, as players under institutional rules, do not operate as simply as median voter theory assumes. The current rise of populism further leads many scholars to question the validity of the institution–median voter framework (Yang et al., 2020; Jones et al., 2022; Figueira, 2018).

Based on this, party power can be seen as influenced by institutions, but it differs from the “proximate causes” proposed by institutional determinism. Party power is not an institution itself, but rather the configuration of actors and the logic of governance within an institutional framework. Under similar democratic and rule-of-law arrangements, it systematically shapes functional income inequality and the evolution of capitalist systems. Thus, to explain distributional heterogeneity *within* democracies, one must examine political competition and governing coalitions inside the system. Party structure—especially the strength of the left—offers a strong and testable mechanism. This paper develops its analysis at this intersection. While continuing to use the IFC perspective to describe capitalist types, we introduce party structure into

the framework. We assess the role of left strength in the evolution of capitalist systems, and empirically control for traditional inclusive institutional factors to study both the moderating effects of institutional constraints and the distributive channels.

More specifically, this paper addresses three research questions. First, after controlling for traditional inclusive institutions, can left strength explain the evolution of capitalist system types? Second, how do political checks and balances and property rights protection shape the effect of left strength on capitalist evolution? Third, does left strength operate mainly through the concentration of capital income or through labor income inequality? Following Rodrik's (2004) distinction of geography, trade, and institutions in the growth literature, this paper treats institutional variables as baseline controls in the empirical analysis. This allows us to test whether left strength works within or beyond the explanatory boundaries of institutions. If left strength still has a significant effect on capitalist types after controlling for traditional institutions, it suggests that party structure provides explanatory power beyond institutional determinism for the transformation of capitalism through distribution. Using cross-nationally comparable data from the LIS database covering nearly 40 countries from 1978 to 2022, we construct an unbalanced panel dataset. We find that, after controlling for traditional institutional variables, left strength significantly promotes the shift of capitalism toward the liberal type. On average, a one percentage point increase in the net seat share of the left reduces the IFC index by about 0.0007 units.

Interestingly, when classifying by political and economic institutions, political institutions tend to weaken this effect, while economic institutions strengthen it. The findings remain robust across a series of robustness checks, alternative estimation strategies, and subsample regression discontinuity analyses.

By addressing these questions, this thesis makes the following marginal contributions. For theoretical contribution: Existing literature on functional distribution often focuses on stylized facts but lacks analysis of the driving forces. Institutional literature, on the other hand, cannot answer the question of heterogeneity within institutions under "democracy versus democracy." It is also often challenged in explaining determinants, with the risk of a single deterministic view. This thesis builds on this gap. It brings left‐wing strength into the debate on institutional determinism. On one side, it explores whether institutions can include party strength within their explanatory scope. On the other side, it helps the functional distribution literature to identify the determinants of capitalist institutional evolution. In addition, this thesis develops a framework of left‐wing strength × institutions. It further separates institutions into political and economic institutions. This provides a more detailed lens to capture heterogeneity inside inclusive institutions, which has been ignored by the determinism literature, and shows its redistributive economic consequences.

For empirical contribution: The literature on welfare states and income distribution

often focuses only on income inequality. It rarely analyzes structural inequality of factors under the classical functional distribution. This thesis uses the IFC index to systematically compare the concentration of capital in income distribution across countries. It argues that this measure better captures different types of capitalism.

For method and data contribution: Existing studies of IFC determinants use samples only from Europe, ignoring global diversity in capitalism. This thesis uses data covering many countries and multiple time periods. It shows the evolution paths of capitalism over time and across regions. This provides more general evidence for comparative political economy.

The structure of the paper is as follows. The next section reviews the literature on functional distribution and income composition inequality, party strength and the welfare state, and institutional determinism. Section 3 presents a simple political economy model to clarify the theoretical framework and testable hypotheses. Section 4 introduces the empirical research design, variable definitions, and data sources. Section 5 reports the regression results. The final section discusses the conclusions and directions for future research.

2) Literature review

2.1 Functional distribution, compositional inequality and capitalist systems

The functional distribution of income has long been a central topic in classical economics. The recent rise of the capital share (Piketty, 2014) has brought factor inequality between capital and labor back into public debate. As early as 1997, Atkinson highlighted both the importance of functional inequality for the study of overall inequality and the lack of research in this area (Atkinson, 1997). He recalled Ricardo's view in classical economics: the division of national income among labor, capital, and land is at the core of political economy's distributional problem. Glyn and Nolan (2011) also stressed the importance of studying functional distribution. They began from the division of national income into wages, profits, and rents, but added a class perspective. They argued that income inequality is closely related to functional distribution. If the bourgeoisie receive only capital income and workers only labor income, then changes in functional distribution directly affect income inequality. However, they did not further develop this class narrative. Milanovic (2017) gave this idea a clearer shape by defining two forms of capitalism. The first is *classical capitalism*, in which society consists mainly of wealthy capitalists and poor workers. Capitalists earn only capital income from asset ownership, while workers earn only labor income. The second is *liberal capitalism*, in which individuals can earn mixed incomes—for

example, labor income alongside significant capital income. Building on this, Ranaldi (2022) introduced the concept of *income compositional inequality* (ICI). Using micro survey data, he divided income into labor and capital components, constructed a Lorenz curve, and created the *Income Composition Factor* (IFC) index to measure ICI. Where IFC is high, capital income is concentrated at the top of the distribution, corresponding to classical capitalism. Where IFC is low, capital and labor incomes are more evenly spread, corresponding to liberal capitalism. This framework combines the classical concern with functional distribution and modern analysis of interpersonal inequality. In classical capitalism, because capital income is clustered at the top, when the rate of return on capital exceeds the growth rate of the economy ($r > g$, in Piketty's terms), inequality cannot converge. In the extreme case of pure classical capitalism (IFC at its maximum, with all capital income held by the top one person), the growth rate of inequality equals the difference between the rate of return on capital and the growth rate of the economy. Functional distribution thus has a strong impact on inequality. By contrast, in liberal capitalism, with a more balanced mix of labor and capital incomes, growth in capital returns does not necessarily increase inequality, so the impact of functional distribution is weaker.

Building on the work of Milanovic (2017) and Ranaldi (2022), recent research has begun to study the link between functional distribution and inequality in a quantitative way, and to connect this issue with the broader process of capitalist development in

political economy. In the writings of classical economists such as Adam Smith, Ricardo, and Marx, classical capitalism was often described as a society in which capitalists earn only capital income, workers earn only labor income, and capitalists dominate the top of the income distribution. Today, however, capitalism is seen as moving beyond this strict divide. More and more individuals earn wages from labor while also receiving returns from accumulated wealth. This creates a “mixed income” model. For example, Ranaldi (2025) uses microdata from the LIS database for 2000 and 2016, covering 80% of world income and 60% of the global population, to analyze income compositional inequality worldwide. The study finds that during this period the share of individuals with positive capital income increased, while global IFC declined. Ranaldi and Milanovic (2022) use the IFC index to classify countries into *classical capitalism* and *liberal capitalism*, and describe stylized facts of capitalism across regions. They conclude that in most countries, higher income compositional inequality is positively correlated with higher overall income inequality. At the national level, Iacono and Ranaldi (2023) use Italian microdata from 1989 to 2016 to measure ICI and argue that Italy is gradually moving toward liberal capitalism with mixed income. Berman and Milanovic (2024) use U.S. income survey data to study *homoploutia*—the share of individuals in the top decile who are simultaneously in the top decile of both capital income and labor income. They find that this share increased from 10% in 1950 to 30%, suggesting that U.S. top earners are increasingly mixed-income. The authors conclude that the strict capitalist–worker divide is losing relevance, and mixed income is

becoming the dominant mode of capitalism. Iacono and Palagi (2022) analyze microdata from 1975 to 2016 in four Nordic countries. They show that income compositional inequality has increased, capital's share has also grown, and together these trends have contributed to rising overall inequality. However, only Petrova and Ranaldi (2024) move beyond descriptive patterns to examine the determinants of IFC. Building on the power resources theory, they show that left-wing governments tend to reduce IFC, thereby extending the theory into the field of income compositional inequality and providing valuable evidence from European democracies. Their contribution opens the way for broader inquiry. While their focus is on advanced European countries, the question of whether similar dynamics hold across a wider set of democracies remains to be explored. In addition, the institutional context is treated in general terms, leaving room for more fine-grained analyses of how political and economic institutions condition the distributive role of party strength. Building on this foundation, the present study expands the scope to nearly forty countries over a longer time horizon, and explicitly distinguishes between institutional dimensions to better understand how partisan politics and institutions jointly shape the evolution of capitalism.

To summarize, the IFC index provides a useful measure to link the classical concept of functional distribution with today's urgent issue of inequality in capitalist development. Empirical studies show that most countries are moving toward liberal capitalism, with

the Nordic countries as a notable exception. This strand of literature highlights the internal dynamics and heterogeneity of capitalism from the angle of factor distribution and income inequality, and offers stylized facts for further research on determinants. While much of this work remains descriptive, Petrova and Ranaldi (2024) take an important step by introducing partisan politics into the analysis. Their findings that left governments tend to reduce IFC provide valuable evidence from European democracies and open the way for more systematic inquiry.

Building on this contribution, this paper approaches the relationship between party strength and IFC from a different perspective. By extending the data to 40 countries, we are able to analyze the determinants of IFC in a more global and generalizable way. The sample includes not only European democracies but also countries with greater institutional heterogeneity. On this basis, the paper incorporates institutional factors and treats parties as co-determinants alongside institutions, rather than as elements fully determined by them, in order to assess the explanatory boundaries of institutional determinism. Furthermore, the analysis introduces institutional interactions, exploring how party strength under different institutional settings affects IFC and, in turn, the evolution of capitalism.

2.2 Left-wing power and distribution

The discussion of parties and redistribution originates from the power resources theory. First proposed by Korpi in his study of class power and the welfare state, the theory defines power resources as the capacity of actors to reward or punish others. Based on this, Korpi explained how the class power of workers, organized through unions and related institutions, could influence welfare state policies (Korpi, 1978/2022; Korpi, 1985/1998). This perspective shifted scholarly attention toward how unions and left-wing parties convert these resources into policies and redistributive outcomes (Esping-Andersen, 1985/2017). Later, Huber and Stephens (2001) used cross-national data to examine redistribution under different welfare state regimes. They found that the long-term rule of different parties leads to major differences in welfare states, poverty, and inequality. Bradley et al. (2003) used the newer LIS database to reassess the validity of power resources theory. They concluded that left governments, through their impact on the size of the welfare state, are the strongest determinant of distribution identified so far. More recently, Petrova and Ranaldi (2024) extend the power resources perspective to income compositional inequality (ICI). Using data from 30 European countries between 2003 and 2017, they construct the IFC index and provide new empirical evidence. They show that under left-wing governments, the polarization between holders of capital and labor income declines. This supports the core claim of power resources theory: party strength is

shaped by the organization of the working class, rather than being an endogenous reflection of redistribution outcomes. At the same time, the literature also notes a weakening of the explanatory power of the left for welfare state variation. Cultural issues—such as gender and identity politics—are gaining importance in left mobilization. The left is shifting from representing primarily the economic interests of the working class to representing a broader set of marginalized groups. As a result, the traditional link between the left and redistribution is becoming less strong.

As welfare state theory evolved, new political theories and employer-centered approaches raised challenges to PRT. Pierson (1994, 1996) was among the first to critique power resources theory and welfare state research. Using case studies of governments and both quantitative and qualitative data from four countries (the UK, the US, Germany, and Sweden), he highlighted the limits of PRT and emphasized the importance of retrenchment and constraint policies. Building on this, Iversen and Soskice (2001) introduced an employer-centered explanation of income distribution. Using survey data from 11 advanced democracies, they argued that workers with specific skills and their employers share preferences for social policies that provide income insurance and protect skill investments. This approach was further developed with Hall and Soskice's (2003) influential *Varieties of Capitalism* (VoC) framework. Estevez-Abe et al. (2001) analyzed the link between specific-skill workers, general-skill workers, and social protection across capitalist types. They showed that in coordinated

economies, employers may also support certain social protections because these complement investments in specific skills and strengthen competitive advantage, thus fostering cross-class coalitions. Together, these studies suggest that, beyond the logic of power resources, heterogeneity within capitalism itself can shape the design of welfare state policies, which in turn influence income and wealth distribution.

Subsequent research, motivated by doubts about power resources theory, has examined how social and state contexts influence party preferences. Iversen and Soskice (2006) were the first to introduce institutions as a factor shaping parties. They built a theoretical model showing that electoral systems play a key role in redistributive heterogeneity within capitalism. Electoral rules shape the nature of parties, determine the composition of governing coalitions, and ultimately affect income distribution. Ansell (2008), using data from 22 OECD countries, studied party preferences through the lens of higher education heterogeneity. He found that left parties' support for or opposition to higher education depends on institutional context: in countries with different education policy arrangements, the attitudes of the left vary significantly. Persson and Tabellini (2003) examined the impact of constitutional arrangements on elections and economic policy. They showed that presidential versus parliamentary systems, and majoritarian versus proportional systems, influence the size of the public sector, government spending, and other political-economic outcomes. In addition, details of electoral rules—such as district magnitude and ballot structure—shape

corruption and the structure of economic growth.

To summarize, the power resources theory argues that welfare states and redistribution depend on the influence of the working class, unions, and left strength on state policy. This perspective successfully explains why left parties affect redistribution and highlights its historical path dependence. Evidence from European samples also shows that left strength, measured by net seat share, reduces IFC. However, as capitalism has developed, the new party theories and the varieties of capitalism literature suggest that welfare state expansion is not determined solely by class struggle. In different macro contexts, employers and capital holders may also support redistribution. Building on this, the latest party theories emphasize that power resources theory overlooks the role of institutional and contextual factors (Häusermann et al., 2013). Left parties influence distribution, but their impact is likely conditioned by the broader institutional environment. We therefore propose that the effect of left strength on IFC is heterogeneous across institutional settings, and should be tested within an “institutions × left strength” interaction framework.

2.3 Institutional determinism

The rise of the public choice school marked the systematic application of neoclassical methods in political economy, and is often seen as the starting point of modern

political economy. Beginning with Downs (1957) and Buchanan & Tullock (1965), scholars tried to incorporate political institutions into economic models using rational voting and median voter theories. In this framework, institutions were usually treated as the precondition for the median voter model, with democracy assumed as exogenous (Meltzer & Richard, 1981; Alesina & Rodrik, 1994). This view changed with North's (1990, 1991) work on institutions and growth. North introduced institutions into the causal chain and emphasized their origins and path dependence. He argued that institutions can be divided into informal norms (customs, taboos, traditions) and formal rules (constitutions, laws, property rights). Institutions shape society and influence economic performance; thus, they are an endogenous link in the causal chain rather than an exogenous "black box." Building on this, Acemoglu et al. (2001) used colonial mortality as an instrument and showed that inclusive institutions—such as political checks and balances and protection of property rights—generated long-term prosperity in colonial settings. Their conclusion, however, faced challenges from geography-based explanations (Sachs, 2003) and human-capital-based explanations (Glaeser et al., 2004). In response, Acemoglu et al. (2005) argued in the *Handbook of Economic Growth* that causal analysis should be separated into three layers. First, background variables—such as colonial origins or legal traditions (common law vs. civil law)—serve as "exogenous shocks" shaping institutional differences. Second, institutions themselves: economic institutions (property rights, market access, contract enforcement) determine long-run development, while political institutions

and power distribution shape economic institutions. Third, outcome variables: economic performance. Other factors, such as technology and human capital, are labeled *proximate causes*, while institutions are the *fundamental cause*. This distinction broadened the explanatory scope of institutional determinism. Later, Rodrik et al. (2004) treated institutions as controls when studying the impact of trade and geography on economic performance. They found that once institutions are controlled for, other factors lose significance. This result largely closed the debate on competing economic determinants.

It can be seen that within institutional determinism, the explanatory scope of institutions has grown stronger through these debates. In much of the literature, redistribution itself is also treated as determined by institutions. In *Why Nations Fail*, Acemoglu and Robinson (2012) link extractive institutions to inequality in their discussion of Soviet development, and argue that only inclusive institutions allow for long-term egalitarian redistribution. In Acemoglu et al. (2006), the authors go further by treating the future distribution of resources as an outcome of institutions, thereby expanding the explanatory boundary. Acemoglu and Robinson (2015) also engaged directly with Piketty's (2014) $r > g$ argument on wealth inequality. They argue that Piketty's claim overlooks institutions and endogenous technological change, making the conclusion incomplete. Drawing on historical cases such as South Africa and Sweden, they show how institutional differences shape the evolution of inequality.

Through these theoretical refinements and empirical strategies, institutional determinism has successfully explained “why some nations are rich while others remain poor.” It established the framework of institutions as the “fundamental cause” in cross-country comparison, extending its reach to wealth and income distribution as institutional outcomes. At the same time, however, this expansion creates the risk of an overly “one-size-fits-all” explanation for political economy variables.

Concerns about whether inclusive institutions, or democracy, reduce inequality are reflected in empirical studies. Acemoglu et al. (2015) use data from 184 countries between 1960 and 2010 to examine the relationship between democracy and redistribution. They find that democracy has a significant and strong effect on tax revenue as a share of GDP, but no significant effect on inequality. The authors argue that, in theory, democracy should promote redistribution and reduce inequality. However, when democracy is captured by wealthier groups through lobbying and other means, its redistributive effect may be undermined. Haggard and Kaufman (2020), through regional comparisons in East Asia, Latin America, and Eastern Europe, study the heterogeneity of democracy’s impact on redistribution. They conclude that democracy often favors the development of broader and more inclusive welfare systems, but the specific outcomes depend heavily on a country’s historical context, institutional legacy, and economic conditions.

To summarize this strand of literature, institutional determinism often relies on binary classifications (inclusive vs. extractive institutions; democracy vs. autocracy). Its focus is mainly on cross-country differences in global development, while paying less attention to the heterogeneous drivers of income and wealth distribution among countries that all share inclusive institutions or democracy. As a result, institutional determinism may have limited explanatory power for the diversity of capitalist forms within democracies. Institutional determinism treats inclusive institutions and democracy as the "fundamental cause." If this claim also holds within democracies, then controlling for institutions should eliminate the effect of the left. This paper proposes the opposite, testable hypotheses: H1 After controlling for political and economic institutions, left strength still significantly reduces IFC. H2: Political checks and balances weaken its effect through the channel of top capital returns. H3: Economic rule of law and property rights protection strengthen its effect through the channel of asset diffusion. H4: The overall outcome is a decline in capital inequality. These hypotheses are formalized in the model in Section 3.

3) Theoretical framework

This paper develops a simple political economy model to clarify the mechanisms and hypotheses. The core outcome variable is the form of capitalism. Here, *ICI* refers to income compositional inequality, and *IFC* is the measurement index. Following the

definition of Ranaldi (2022):

$$IFC \in [-1,1] \tag{1}$$

Where, a higher IFC index indicates greater income compositional inequality, meaning the society is closer to the classical capitalism type. The core explanatory variable of the model is left strength L_t, measured by the net parliamentary seat share of the left in a given country and year. To engage with the traditional institutional determinism literature, the model also includes measures of inclusive institutions: political checks and executive constraints Z_c, and property rights and rule of law strength in the economic sphere Z_p. More specifically, let Z_c denote government administrative capacity, meaning whether the government can take action. Let Z_p denote the economic institutional environment, meaning how effective government action is once taken. For simplicity, we assume that Z_c and Z_p are exogenously given in the first period, and that left strength (L_t) is approximately exogenous in the short run (due to bounded rationality leading the median voter to temporarily fail).

The model has two periods. In the first period, Z_c and Z_p are exogenously given as institutional variables, while left strength L_t is exogenously given as the distributive variable. In the second period, the left can choose two types of policy tools to redistribute capital income: One is a fiscal tax policy, denoted as $\tau(L_t)$. The other is a universal financial policy that lowers entry barriers to capital, denoted as $s(L_t)$. Both types of policies can reduce capital inequality in society. They have an increasing

effect: the stronger the left, the stronger the policies. But there are diminishing marginal returns. Based on this, the dynamics of capitalism can be written as the following equation:

$$\Delta IFC_{i,t+1} = -\theta L_t[a\kappa(Z_c) + b\phi(Z_p)] + \rho IFC_{it} + \mu_{i,t+1} \quad (2)$$

Here, $\kappa(Z_c)$ and $\phi(Z_p)$ denote, respectively, the strength with which political checks and balances suppress the capital tax, and the efficiency with which property-rights protection in the economy enables asset holding at the bottom of the income distribution. Like the policy tools, we assume both efficiency functions are concave. However, the political-rights function is decreasing $\kappa'(Z_c) < 0$, while the property-rights function is increasing $\phi'(Z_p) > 0$. The residual term $\rho IFC_{it} + \mu_{i,t+1}$ captures noise in the evolution of the capitalist type. In the empirical work, we address it by using lagged explanatory variables and covariates, or by taking averages of these variables over *t* lags for robustness. Given the evolution equation, the first-order effect of the left is:

$$\frac{\partial \Delta IFC_{i,t+1}}{\partial L_{it}} = -\theta[a\kappa(Z_c) + b\phi(Z_p)] < 0 \quad (3)$$

From this, we derive the first testable hypothesis of the model: after controlling for inclusive institutions, greater left strength reduces IFC, making income composition more equal and pushing the system toward “liberal capitalism.”

Since the institutional effect and the distributive effect of the left interact, we can test

the second-order effects of left strength with the two types of inclusive institutions as follows:

$$\frac{\partial^2 \Delta IFC_{i,t+1}}{\partial L_{it} \partial Z_c} = -\theta a \kappa'(Z_c) > 0 \qquad \text{Because } (\kappa' < 0) \tag{4}$$

$$\frac{\partial^2 \Delta IFC_{i,t+1}}{\partial L_{it} \partial Z_p} - \theta a \phi'(Z_p) > 0 \qquad \text{Because } (\phi' < 0) \tag{5}$$

From this, we derive the second and third testable hypotheses of the model: stronger political checks and balances restrict the left's ability to reduce top capital returns, thereby weakening its effectiveness in driving the shift toward liberal capitalism; by contrast, stronger rule of law and property rights increase the efficiency with which lower-income groups can turn transfers and public spending into secure assets, making the asset-diffusion channel more effective and strengthening the left's role in promoting the shift toward liberal capitalism.

In addition, following Ranaldi (2022) and Petrova & Ranaldi (2024), there are only four direct channels through which IFC can be reduced: (1) lowering capital income at the top, (2) lowering labor income at the bottom, (3) increasing labor income at the top, and (4) increasing capital income at the bottom. When we focus on capital inequality, measured by the pseudo-Gini of capital G_k, left tax policies $\tau(L_t)$ reduce capital income at the top. Meanwhile, universal financial policies—such as public pension funds, ESOPs, inclusive finance, baby bonds, or universal capital accounts—raise capital income at the bottom. Both policy tools work by reducing the capital pseudo-

Gini G_k , which in turn lowers IFC and shifts the system toward liberal capitalism. For the change in the capital pseudo-Gini, we have:

$$\Delta G_k^{i,t+1} = -\alpha\tau\left(L_{it}\right)\kappa(Z_c) - \beta s\left(L_{it}\right)\phi(Z_p) + \rho G_k^{it} + \varepsilon_{it} \tag{6}$$

The first-order condition of left policies for capital inequality is:

$$\frac{\partial \Delta G_k^{i,t+1}}{\partial L_{it}} = -\alpha\tau'\left(L_{it}\right)\kappa(Z_c) - \beta s'\left(L_{it}\right)\phi(Z_p) < 0 \tag{7}$$

From this, we derive the fourth testable hypothesis of the model: greater left strength reduces capital inequality, which in turn lowers IFC and drives the system toward liberal capitalism.

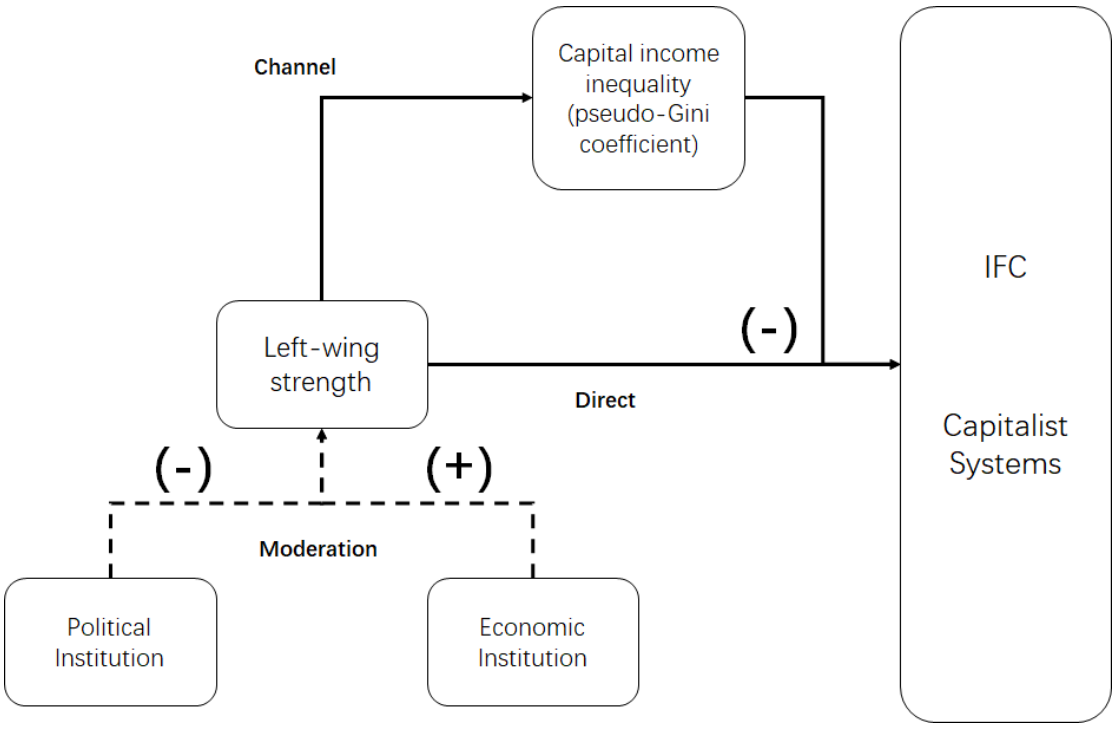

Figure 1 Conceptual framework: Left-wing strength, institutions, and IFC.

4) Empirical design and data

4.1 Variables

The dependent variable of this study is the IFC index. IFC measure income composition inequality (ICI) and can be regard as a proxy for the types of capitalism. Following the method proposed by Ranaldi (2022), the paper calculates the level of income composition inequality by dividing micro-level income data into capital income and labor income, and then constructing Lorenz curves based on the ranking of total income. Specifically, for the concentration of capital income, the IFC index is defined as follows:

$$IFC = \frac{A(\pi)}{B^{Max}(\pi)} \tag{8}$$

Where $A(\pi)$ is the area between the capital Lorenz curve and the capital Zero-Concentrate curve. $B^{Max}(\pi)$ is the area between the capital Zero-Concentrate curve and the capital Maximum-Concentrate curve. These three capital Lorenz curves are explained as follows. First, suppose that total income in society is split into capital income and labor income. In this setting, the capital share of society is the part of total income that comes from capital income, written as $\pi = \frac{\Pi}{Y}$. For each individual i, the capital share is the ratio of their capital income to their total income, written as $\pi_i = \frac{\Pi_i}{y_i}$. The capital concentration curve $f(\pi, p)$ shows how capital income is

distributed across income percentiles. To build this curve, all individuals are ranked by total income in ascending order. Then, for each percentile on the total income curve, the income is multiplied by the individual's capital share. This produces the capital concentration curve. This results in the capital concentration curve:

$$f(\pi, p) = \pi_i \bullet \sum_{i=1}^{n} y_i \tag{9}$$

For the Zero-Concentrate curve $f^e(\pi, p)$, it describes an ideal case where there is no inequality in income composition. In such a society, every person's capital share is the same as the overall capital share of the economy. To construct this curve, we first compute the total capital share of society, then multiply each point on the total income curve by this same capital share. The result is:

$$f^e(\pi, p) = \pi \bullet \sum_{i=1}^{n} y_i \tag{10}$$

The final curve is the Maximum-Concentrate curve, which represents a society with the highest possible level of income composition inequality. This curve is constructed in a specific way, as the IFC index may take negative values. Therefore, the maximum concentration curve is derived based on the relationship between the Zero-Concentrate curve (the ideal, equal-composition curve) and the actual capital concentration curve:

$$f^m(\pi, p) = \begin{cases} f(y, p) & \text{for p} > \text{p*} \\ \pi & \text{for p} \le \text{p*} \end{cases} \tag{11}$$

$$f^{M}(\pi, p) = \begin{cases} 0 & \text{for p} \leq \text{p*} \\ f(y, p) - \pi & \text{for p} > \text{p*} \end{cases} \quad (12)$$

When the Zero-Concentrate curve is above the capital concentration curve (the common case), the Maximum-Concentrate curve works as the horizontal axis for most of the distribution, that is $f^{M}(\pi, p)$. When the Zero-Concentrate curve is below the capital concentration curve, the Maximum-Concentrate curve is instead defined as $f^{m}(\pi, p)$.

Regarding the classification of labor income and capital income, income survey datasets often categorize income into labor income, capital income, operating income, and transfer income. There are methodological differences in how operating income is divided between labor and capital components. For example, a common approach is to assign 30% of operating income to capital income and 70% to labor income (Alvaredo et al., 2016). In studies covering a wide range of countries, such as those including Latin America where agriculture plays a large role, operating income is sometimes entirely classified as labor income (Ranaldi and Milanovic, 2022). This is often due to agricultural income being included in operating income in these datasets. In the baseline definition (IFC_def1), this paper assigns 30% of operating income to capital income and 70% to labor income. As a robustness check, we also test an alternative definition (IFC_def2), which assigns all operating income to labor income. Details on the LIS variables used in the construction of the IFC index can be found in

the appendix Table 8 & Table 9 and in the code sections for *def1* and *def2*.

The explanatory variable of this paper is left government strength. We use cross-nationally comparable data from the Database of Political Institutions (DPI). The DPI provides, for each country-year, the ideology (left or right) and parliamentary seats of the three largest governing parties and the largest opposition party, as well as the total number of seats. Together, these four parties account for most of the legislature (see Appendix Table 10, where their combined share is about 85%). Based on this, we measure left strength as follows: we take the total seats of left parties among these four parties, subtract the total seats of right parties, and then divide this net left seat number by the total seats in parliament. This gives the left strength in a given country-year. Specifically: Let $S_1^{gov}, S_2^{gov}, S_3^{gov}$ denote the number of seats held by the three largest governing parties, S_1^{opp} denotes the number of seats held by the largest opposition party, and S_{tot} denotes the total number of parliamentary seats. Define $I\frac{L/R}{gov,j}$ as an indicator variable that takes the value +1 if party j is classified as left-wing, −1 if right-wing, and 0 otherwise. The numerator therefore represents the number of left-wing seats minus right-wing seats among these four major parties, while the denominator is the total number of parliamentary seats. Thus, left-wing strength is defined as：

$$left = \frac{I\dfrac{L/R}{gov,1}\cdot S_1^{gov} + I\dfrac{L/R}{gov,2}\cdot S_2^{gov} + I\dfrac{L/R}{gov,3}\cdot S_3^{gov} + I\dfrac{L/R}{opp,1}\cdot S_1^{opp}}{S_{tot}} \quad (13)$$

The control variables for traditional institutions are taken from the institutional determinism literature. First, for political checks and balances, we use *exconst* from the Polity IV database. In Acemoglu et al. (2001), this variable is treated as a measure of "limited government," capturing the extent of constraints on the executive and thus reflecting the security of property rights. Second, for property rights, Acemoglu et al. (2001) used the *expropriation rate* from the PRS Group's *International Country Risk Guide* (ICRG) to measure the risk of property confiscation. In the latest version of the ICRG, however, this variable has been merged into the broader *Investment Profile* (IP). In response, some recent studies following institutional determinism have adapted by using different proxies: for example, Azzimonti (2018) directly uses IP, while Lei & Luo (2022) use *Law and Order* (LO) from ICRG, together with three other indicators, normalized to create an institutional index. Both IP and LO are close to the economic institutions emphasized in the literature, but they also contain unrelated components, which may contaminate estimates. Therefore, this paper chooses the *Rule of Law* (RL) indicator from the World Bank's *Worldwide Governance Indicators* (WGI) as the main measure of property rights protection. The World Bank defines RL as a measure of confidence in and compliance with society's rules, including contract enforcement and property rights protection (Kaufmann et al., 2009). Overall, RL better replicates the

expropriation risk used in Acemoglu et al. (2001). In the empirical analysis, we use *Rule of Law* as the main measure of property rights protection, and employ *Law and Order* (LO) and *Investment Profile* (IP) as alternative controls in robustness checks.

This paper follows the principle of controlling as much as possible for country characteristics when selecting other control variables. Specifically, we include: the log of GDP per capita (*lnpgdp*) to control for development level; the log of population (*lnpop*) to control for country size; GDP growth (*gdp_growth*) to capture cyclical fluctuations; and population growth (*pop_growth*) to reflect demographic pressure. In addition, we control for the share of final consumption expenditure in GDP (*con*) to capture the demand-side structure of the economy; the ratio of FDI inflows to GDP (*fdi*) to capture cross-border capital flows and openness; and the ratio of trade to GDP (*trade*) to capture overall openness. Finally, we include the number of years of primary education (*edu*) to measure the institutional basis of human capital accumulation.

4.2 Model

Based on the theoretical analysis and testable hypotheses above, this paper employs a two-way fixed effects panel model for the baseline estimation. The baseline regression model is specified as follows, to test Hypothesis 1:

$$IFC_{it} = \alpha_0 + \alpha_1 left_{it} + \alpha_2 institution_{it} + \alpha_3 control_{it} + \delta_i + \gamma_t + \varepsilon_{it} \quad (14)$$

Where IFC_{it} is the compositional inequality index and represents the type of capitalism in country i at year t. α_0 is the intercept. $left_{it}$ is the Left-wing strength of country i in year t. The term $institution_{it}$ denoted the control variable of institutions, which include $xconst_{it}, LO_{it}, IP_{it}$ and RL_{it}. The term $control_{it}$ represents the other control variables in country level. δ_i and γ_i are country fixed-effect and time fixed-effect, respectively. Finally, ε_{it} is the error term.

To test Hypotheses 2 and 3, this paper employs an interaction model between institutions and left strength, while controlling for country and year fixed effects, in order to identify the moderating effects of political and economic institutions on left strength:

$$IFC_{it} = \alpha_0 + \alpha_1 left_{it} + \alpha_2 institution_{it} + \alpha_3 left_{it} \times institution_{it} + \alpha_3 control_{it} + \delta_i + \gamma_t + \varepsilon_{it} \quad (15)$$

Where $left_{it} \times institution_{it}$ denoted the interaction of left-wing strength and institutional variables.

To test Hypothesis 4, this paper calculates the pseudo-Gini coefficient of capital income (*mu_p*) and the pseudo-Gini coefficient of labor income (*mu_w*) from their respective Lorenz curves, in order to examine the channels through which left strength affects IFC:

$$\mu_{it}^{p/w} = \alpha_0 + \alpha_1 left_{it} + \alpha_2 institution_{it} + \alpha_3 control_{it} + \delta_i + \gamma_t + \varepsilon_{it} \quad (16)$$

Where μ^p denotes the pseudo-gini coefficient of capital income, and μ^w represents the pseudo-gini coefficient of labor income.

4.3 Data

The empirical data of this paper come from several cross-nationally comparable databases. For the measurement of the IFC index, we follow Ranaldi and Milanovic (2022) and use capital income and labor income data from the LIS database. Left strength is drawn from the Database of Political Institutions (DPI). Among inclusive institutions, political checks are measured with Polity IV, while economic institutions are taken from the PRS Group's *International Country Risk Guide* (ICRG) and the World Bank's *Worldwide Governance Indicators* (WGI). Other control variables are obtained from the World Bank's *World Development Indicators* (WDI). Due to data availability in LIS and WDI, the final regression sample is an unbalanced panel. The descriptive statistics are presented below:

Table 1 Descriptive statistic

Variable	Obs	Mean	Std. Dev.	Min	Max	Data Source
IFC_def1	858	.093	.196	-.512	.753	LIS
IFC_def2	809	.216	.209	-.44	.79	LIS
left	748	-.035	.301	-1	1	DPI
xconst	698	6.795	.667	1	7	POLITY IV
RL	618	.962	.877	-1.123	2.015	WGI
IP	770	9.527	1.918	3.167	12	ICRG
LO	770	4.687	1.344	1	6	ICRG
con	828	75.705	8.509	35.27	107.924	WDI
fdi	802	4.676	15.484	-28.968	252.987	WDI
trade	828	81.719	56.617	14.391	382.661	WDI
lnpgdp	829	10.029	.9	6.794	11.63	WDI
lnpop	830	16.775	1.649	12.576	21.062	WDI
gdp_growth	829	2.712	3.339	-14.839	24.616	WDI
pop_growth	830	.663	.759	-2.517	3.455	WDI

edu	830	5.595	.947	3	8	WDI

5) Empirical results

5.1 Baseline regression

Given the presence of contemporaneous common shocks and serial correlation in cross-country panel data, we use Driscoll–Kraay robust standard errors in the baseline estimation. This approach is robust to cross-sectional dependence and allows for serial correlation in panels with long *T*. All baseline tables report DK standard errors, while Appendix Table 12 provides results with conventional cluster-robust errors for comparison. The baseline regression results are presented in Table 2:

Table 2 Baseline regression

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	IFC	IFC	IFC	IFC	IFC	IFC
left	-0.079***	-0.076***	-0.070***	-0.042**	-0.073***	-0.071***
	(-3.33)	(-3.10)	(-3.12)	(-2.09)	(-2.96)	(-2.96)
xconst			-0.000			
			(-0.01)			
RL				-0.101***		
				(-3.08)		
IP					0.010***	
					(2.88)	
LO						0.019
						(1.47)
con		-0.001	-0.001	-0.001	-0.001	-0.000
		(-0.41)	(-0.32)	(-0.62)	(-0.32)	(-0.12)
fdi		-0.000	0.000	-0.000	-0.000	-0.000
		(-0.67)	(0.37)	(-0.28)	(-0.74)	(-0.70)
trade		0.001	0.001	-0.000	0.001	0.001
		(1.31)	(1.00)	(-1.02)	(1.29)	(1.17)

lnpgdp		0.019	0.059	0.051	0.008	0.021
		(0.27)	(0.76)	(0.76)	(0.10)	(0.31)
lnpop		0.026	0.034	-0.081	-0.050	-0.029
		(0.28)	(0.28)	(-0.53)	(-0.47)	(-0.28)
gdp_growth		0.004***	0.003**	0.005**	0.004**	0.005***
		(2.72)	(2.44)	(2.67)	(2.06)	(2.96)
pop_growth		0.054***	0.050***	0.049***	0.046***	0.050***
		(4.72)	(3.65)	(6.65)	(4.17)	(5.12)
edu		0.011	0.015	-0.023	0.000	-0.003
		(0.46)	(0.58)	(-0.91)	(0.01)	(-0.11)
Constant	-0.126***	-0.852	-1.253	0.000	0.615	0.101
	(-6.02)	(-0.46)	(-0.59)	(.)	(0.32)	(0.05)
Observations	748	696	612	539	653	653
Country FE	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

Column (1) reports the estimation with only two-way fixed effects, without controlling for institutional variables. The results show that when the left's seat share relative to the right increases by 10 percentage points, IFC decreases by 0.0079 units, significant at the 1% level. Column (2) adds macro-level controls, and left strength still significantly reduces IFC. To avoid serious multicollinearity between institutional variables, following the institutional determinism literature, we do not include political and economic institutions in the same specification. Columns (3)–(6) report results controlling for political institutions and, separately, each of the three economic institutional indicators. In all cases, left strength continues to significantly reduce IFC. For model choice, Hausman test results comparing random and fixed effects are presented in Appendix Table 11, supporting the use of fixed effects. Results with

conventional cluster-robust standard errors are reported in Appendix Table 12 and remain statistically significant. Overall, the baseline regressions support Hypothesis 1: after controlling for traditional institutional factors, left strength significantly reduces income compositional inequality, facilitating the transition from classical to liberal capitalism.

5.2 Moderation effect

Next, we extend the baseline regressions by adding interaction terms between traditional institutional factors and left strength. This allows us to test how political and economic institutions moderate the effect of left strength on reducing IFC. The results are presented in Table 3:

Table 3 Moderation analysis

VARIABLES	(1) IFC left#xconst	(2) IFC left#RL	(3) IFC left#IP	(4) IFC left#LO
left	-0.518***	0.015	-0.077	0.014
	(-3.08)	(0.55)	(-0.64)	(0.20)
xconst	0.027			
	(1.38)			
c.left#c.xconst	0.066**			
	(2.55)			
RL		-0.109***		
		(-3.18)		
c.left#c.RL		-0.058***		
		(-2.97)		
IP			0.010***	
			(3.00)	
c.left#c.IP			0.000	
			(0.04)	

LO				0.021
				(1.60)
c.left#c.LO				-0.018
				(-1.17)
con	-0.001	-0.001	-0.001	-0.000
	(-0.34)	(-0.77)	(-0.32)	(-0.18)
fdi	0.000	-0.000	-0.000	-0.000
	(0.33)	(-0.28)	(-0.74)	(-0.70)
trade	0.001	-0.000	0.001	0.001
	(0.89)	(-0.91)	(1.29)	(1.25)
lnpgdp	0.056	0.068	0.008	0.020
	(0.71)	(1.11)	(0.10)	(0.30)
lnpop	0.025	-0.087	-0.050	-0.028
	(0.21)	(-0.59)	(-0.47)	(-0.28)
gdp_growth	0.003**	0.005**	0.004**	0.005***
	(2.66)	(2.50)	(2.07)	(2.89)
pop_growth	0.047***	0.050***	0.046***	0.051***
	(3.35)	(6.65)	(4.41)	(5.11)
edu	0.023	-0.026	0.000	-0.006
	(0.87)	(-1.09)	(0.02)	(-0.26)
Constant	-1.297	0.000	0.617	0.128
	(-0.61)	(.)	(0.32)	(0.07)
Observations	612	539	653	653
Country FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

Column (1) reports the moderating effect of political institutions. At the 1% level of significance, political institutions significantly weaken the effect of left strength on reducing IFC, thus constraining the role of the left in driving capitalist evolution. This result supports Hypothesis 2. Column (2) reports the moderating effect of economic institutions. At the 1% level, economic institutions significantly strengthen the effect of left strength in reducing IFC, thereby amplifying the left's role in promoting capitalist transformation. To illustrate the heterogeneity brought by institutions more directly,

this paper also plots marginal effects. Figure 2 shows the marginal effect of political institutions (*xconst*): when political constraints are low (*xconst = 1*), left strength significantly reduces IFC; but as political constraints increase, the effect of the left declines sharply (at *xconst = 7*, the effect of the left on IFC is only about –0.1 and almost loses significance).

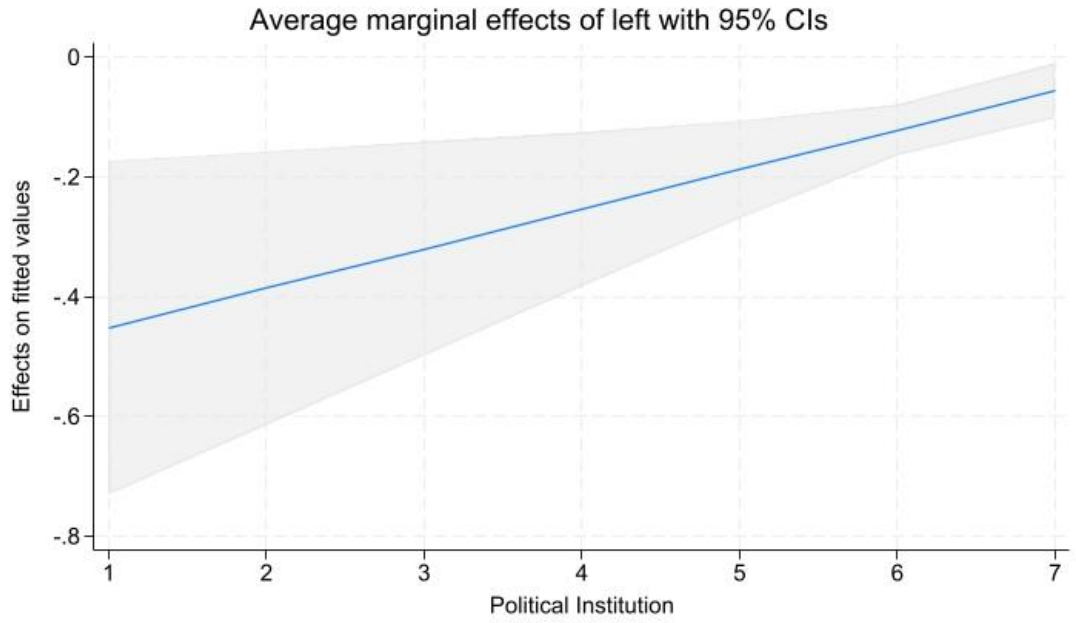

Figure 2 Marginal effect of xconst

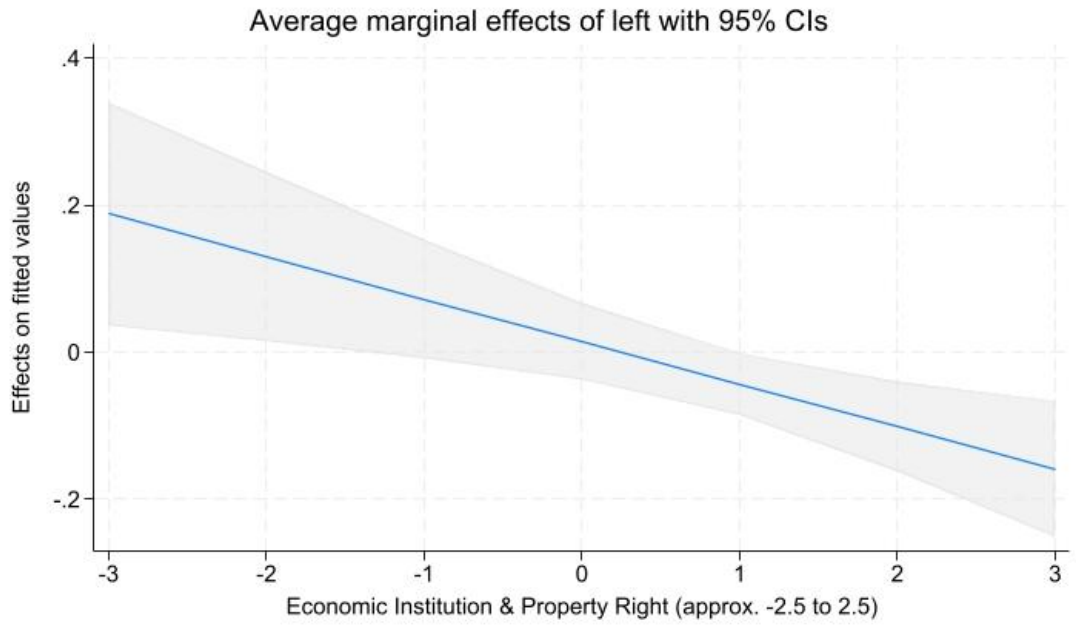

Figure 3 Marginal effect of RL

Figure 3 shows the marginal effect of economic institutions (*RL*): when the rule of law is low (*RL = –2*), left strength significantly increases IFC (the confidence interval does not cross zero); as the level of economic institutions rises, the effect of the left in reducing IFC becomes stronger, and once *RL* exceeds 1, left strength significantly lowers IFC.

Also, we use cluster-robust standard errors to further test if the moderation effect is still significant. Results are reported in Appendix Table 13. Moreover, to address potential noise, we introduce lagged explanatory variables and covariates, and test the moderating effects using different time windows. The results remain robust (see Appendix Table 15). Taken together, these empirical findings support Hypotheses 3 and 4.

Columns (3) and (4) use two alternative proxies for traditional economic institutions—*Investment Profile* (IP) and *Law and Order* (LO). As anticipated, the interaction between IP and left strength shows a sign reversal and is not significant. For LO, the interaction still points in the expected direction, indicating that it helps the left reduce IFC, but the effect is small (t = 1.17) and lacks statistical significance. We interpret this as arising from the fact that both IP and LO contain many components unrelated to economic institutions or property rights (especially IP), which introduces bias into the estimates. This outcome is consistent with our expectations.

5.3 Channel analysis

Next, we introduce the pseudo-Gini coefficient of capital (*mu_p*) to analyze Hypotheses 4 and 5. Following Petrova & Ranaldi (2024), we calculate *mu_p* as one minus the area under the concentration curve of capital income, using it as a proxy for capital inequality. A larger *mu_p* indicates greater capital inequality in society. The results of the channel analysis are presented in Table 4:

Table 4 Channel analysis

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	mu_p	mu_p	mu_p	mu_w	mu_w	mu_w
	Capital inequality			Wage inequality		
left	-0.026***	-0.024***	-0.015*	0.006	0.008*	0.002
	(-3.03)	(-2.91)	(-2.07)	(1.63)	(1.79)	(0.89)
xconst		-0.011			-0.010***	
		(-1.66)			(-3.22)	

RL			-0.039***			0.005
			(-2.91)			(1.62)
con	-0.001	-0.001	-0.001	-0.001***	-0.001***	-0.001**
	(-0.97)	(-1.01)	(-1.02)	(-3.61)	(-4.10)	(-2.57)
fdi	-0.000	0.000	-0.000	0.000	0.000	-0.000
	(-0.30)	(0.84)	(-0.54)	(1.15)	(1.48)	(-0.77)
trade	0.000*	0.000	-0.000	0.000	0.000	0.000***
	(1.68)	(1.60)	(-0.54)	(0.61)	(1.29)	(2.92)
lnpgdp	-0.043	-0.034	-0.050*	-0.066***	-0.077***	-0.078***
	(-1.26)	(-0.94)	(-1.92)	(-7.05)	(-8.40)	(-6.65)
lnpop	0.015	0.032	0.039	-0.024**	-0.007	0.027***
	(0.42)	(0.70)	(0.60)	(-2.09)	(-0.58)	(3.27)
gdp_growth	0.001	0.000	0.001*	-0.000	-0.001	-0.000
	(1.23)	(0.57)	(1.92)	(-1.42)	(-1.50)	(-0.93)
pop_growth	0.019***	0.016**	0.013***	0.001	-0.000	-0.003**
	(3.56)	(2.54)	(4.12)	(0.53)	(-0.14)	(-2.63)
edu	0.005	0.006	-0.012	0.000	-0.001	0.003
	(0.47)	(0.61)	(-1.11)	(0.12)	(-0.35)	(1.54)
Constant	0.799	0.549	0.000	1.796***	1.694***	0.000
	(1.11)	(0.65)	(.)	(7.71)	(6.97)	(.)
Observations	696	612	539	696	612	539
Country FE	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

Column (1) reports the effect of left strength on capital inequality without controlling for institutional factors. The results show that left strength significantly reduces capital inequality at the 1% level. Column (2) shows that this effect remains robust after controlling for political institutions. Column (3) adds economic institutions, where the significance of left strength declines slightly but remains statistically meaningful, confirming the robustness of the channel analysis. Columns (4)–(6) present the results

for the pseudo-Gini of labor income. In all cases, left strength does not reduce labor inequality. We therefore conclude that left strength does not lower IFC through labor inequality. This finding is consistent with Hypothesis 4: left strength reduces IFC primarily by lowering capital inequality.

5.4 Endogeneity check

Although the baseline regressions with two-way fixed effects account for unobserved country and year effects, they can only partly address the possibility that left strength lowers IFC and pushes society toward liberal capitalism. Potential bidirectional causality and omitted variables may still bias the estimates. Unlike the Gini coefficient, the IFC index does not directly reflect the kind of inequality that classical political economy models expect to drive demand for the left (Meltzer & Richard, 1981). Income composition inequality between capital and labor is less salient to voters than interpersonal income inequality, which creates sustained political pressure. Still, it is possible that countries with higher IFC also tend to have higher Gini inequality, which in turn makes left parties more likely to come to power. To address this, we apply the generalized method of moments (GMM), constructing a dynamic panel by including lagged IFC. However, given the long time dimension and gaps in data coverage, the panel is relatively unbalanced. In the initial attempts, even after collapsing instruments and restricting lag lengths of both the dependent variable and explanatory variables

to two periods, the number of instruments still exceeded the number of countries once year dummies were added (about 53 instruments vs. 30 countries). This caused the Hansen test to fail. To resolve this issue, we group the years 1978–2022 into 11 four-year periods and use these as period dummies in the instrument set. The system GMM estimation results are reported as Table 5:

Table 5 GMM

	(1)	(2)	(3)
VARIABLES	IFC	IFC	IFC
	SYS-GMM	DIF-GMM	SYS-GMM
L.IFC_def1	0.000	0.000	0.981***
	(.)	(.)	(5.72)
left	-0.117*	-0.133*	-0.077*
	(-1.98)	(-1.97)	(-1.95)
xconst	0.067***	0.051***	0.012
	(6.38)	(5.29)	(0.14)
con	-0.002***	-0.002***	0.006
	(-3.50)	(-4.13)	(1.01)
fdi	0.003***	-0.001**	-0.001
	(3.83)	(-2.60)	(-1.05)
trade	-0.001***	-0.001***	0.001
	(-9.99)	(-6.68)	(0.94)
lnpgdp	0.087***	0.091***	0.027
	(11.99)	(13.16)	(0.66)
lnpop	-0.042***	-0.041***	0.027
	(-10.88)	(-9.11)	(0.96)
gdp_growth	0.010***	0.000	0.006
	(5.06)	(0.09)	(1.29)
pop_growth	0.043***	0.042***	0.023
	(7.28)	(6.24)	(0.74)
edu	-0.060***	-0.043***	0.000
	(-9.99)	(-8.32)	(0.02)
Constant	0.000	0.000	-1.047
	(.)	(.)	(-0.93)
AR(1)	P=0.000***	P=0.027**	P=0.005***
AR(2)	P=0.653	P=0.131	P=0.272

Hansen test	P=1.000	P=1.000	P=0.626
Observations	446	446	446
IV Number	53	53	31
Country Number	28	28	28
Period FE	NO	NO	YES
Year FE	YES	YES	NO

Robust t-statistics in parentheses

*** p<0.01, ** p<0.05, * p<0.1

The results show that across all model specifications, left strength significantly reduces IFC, supporting the argument that it promotes the transition toward liberal capitalism. All three models pass the AR tests: first-order autocorrelation is significant, as expected given the inclusion of the lagged dependent variable, while second-order autocorrelation is not significant, consistent with the assumption that higher-order lags are valid instruments. This confirms the necessity of including lags of two periods or more. However, instrument validity remains a concern. Although the Hansen test rejects the null of weak instruments, this test is unreliable when the number of instruments is large. In models (1) and (2), despite collapsing the instruments and restricting endogenous lags to two periods, the inclusion of year fixed effects still produces too many instruments for a relatively small sample. This raises the risk of weak instruments and reduces explanatory power. In column (3), we address this issue by replacing yearly dummies with 4-year period dummies as instruments. In this case, the Hansen test yields a p-value of 0.626, indicating that the instrument set is valid and avoiding the problem of instrument proliferation. This provides stronger support for the model specification. Nonetheless, when experimenting with alternative lag structures, the direction of the results remains consistent but significance becomes

unstable. This suggests that while the GMM results are broadly supportive, their explanatory strength is relatively weak.

To address the problems of instrument proliferation and unstable results in dynamic panel estimation, this paper introduces a fuzzy regression discontinuity design (Fuzzy RDD) for the endogeneity analysis. Specifically, using DPI data, we calculate the "net left votes" as the vote share of the three largest governing parties plus the largest opposition party that are left-wing, minus those that are right-wing (see Appendix Table 10, where these parties account for about 85% of total votes). When the net left votes are just above zero, the probability of a left party coming to power increases sharply. This provides the basis for identifying left incumbency as the treatment variable. Because DPI does not include complete vote data for all parties, we apply a fuzzy RDD and interpret it as an instrumental variables strategy. In the first stage, the sign of net left votes serves as the instrument for left incumbency. In the second stage, left incumbency is used to estimate its causal effect on IFC. It is important to note that the RDD analysis differs from the baseline regressions in both the explanatory variable and the sample scope. The baseline regressions use *left strength* (seat share) as a continuous measure reflecting the overall political balance, while RDD relies on a binary treatment—*left incumbency*—to identify a quasi-experimental causal effect at the cutoff. Moreover, RDD identification depends on bandwidth choice and is based only on subsamples close to the cutoff. This reduces the number of observations

compared to the baseline regressions, but precisely because of this local focus, RDD offers a cleaner causal test of the theoretical hypothesis. In sum, we treat the RDD as a complementary robustness check: a sub-sample analysis, using an alternative measurement of the explanatory variable, that provides a more credible causal identification of the same theoretical claim.

In terms of model specification, the classic literature argues that RDD relies mainly on the local discontinuity at the cutoff and therefore does not require many control variables (Hahn & Todd, 2001). In practice, however, approaches vary. For example, Lee et al. (2004), studying the relationship between left incumbency and government spending in advanced democracies, included only time effects. Their reasoning was that the running variable is a country-level feature, and adding country fixed effects could create multicollinearity. By contrast, Ferreira & Gyourko (2009), in the U.S. context, used both country and time fixed effects, along with additional controls. Following these approaches, this paper conducts robustness checks of the RDD results under different model specifications. All fuzzy RDD models are estimated using kernel-weighted local polynomial regression with a triangular kernel, which assigns greater weight to observations close to the cutoff and progressively lower weight to those farther away. The regression results are reported in Table 6:

Table 6 Fuzzy RDD

VARIABLES	(1)	(2)	(3)	(4)

	OLS	Year FE	Year FE	Two-way FE
	First Stage Regression			
	inpower	inpower	inpower	inpower
vote_win	0.641***	0.612***	0.583***	0.593***
	(6.905)	(3.72)	(3.24)	(3.40)
	Second Stage Regression			
	IFC	IFC	IFC	IFC
inpower	-0.261***	-0.383*	-0.350*	-0.011
	(-3.48)	(-1.97)	(-1.70)	(-0.17)
vote_win		0.033*	0.021	0.014
		(1.82)	(0.70)	(1.69)
vote_win#win			0.191	0.016
			(0.65)	(-1.74)
Observations	632	174	174	168

z&t-statistics in parentheses
Raw (2) (3) and (4) used robust standard error.
*** p<0.01, ** p<0.05, * p<0.1

Column (1) reports the pooled RDD results. The first stage is significant, showing that missing party vote information in the DPI does not bias our estimation. The second stage indicates that left incumbency significantly reduces IFC, pushing society toward liberal capitalism. Column (2) adds year fixed effects and robust standard errors. The effect of left incumbency remains significant at the 10% level. Column (3) further includes interactions between the running variable and the instrument to ensure robustness of the discontinuity at the cutoff. The results continue to show that left incumbency significantly lowers IFC. In Column (4), country fixed effects are added. While the coefficient sign remains consistent, significance disappears. This is mainly because RDD already relies on local samples around the cutoff, and country fixed effects absorb additional cross-national variation. Identification then depends only on

a few election switches within each country, reducing effective variation and inflating standard errors. Similar to Lee et al. (2004), this supports the choice to include time effects but not country effects, since the running variable itself is defined at the country level and adding country fixed effects risks multicollinearity. Appendix Table 18 reports results with alternative bandwidths: a wider bandwidth (main × 1.2) and a narrower one (main × 0.8). In both cases, the effect of left incumbency remains significant. Appendix Table 19 and Table 20 provide diagnostic checks. The McCrary density test shows that the running variable (left net votes, *vote_win*) is continuous at the cutoff, ruling out manipulation. Placebo tests shifting the cutoff by ±3 and ±6 lead to coefficient sign reversals or large declines in significance. Taken together, these findings confirm that the RDD results are robust.

5.5 Other robustness checks

This section reports the remaining robustness checks.

First, we change the definition of capital and labor shares in IFC. In the baseline regressions, following Alvaredo et al. (2016), we classify 30% of self-employment income as capital income and 70% as labor income. However, Ranaldi and Milanovic (2022) argue that since the LIS database includes agricultural income under self-employment, this traditional split may underestimate IFC in agriculture-intensive

regions (e.g., Latin America). Therefore, we construct an alternative definition, *IFC_def2*, which assigns all self-employment income to labor. Column (1) of Table 7 shows that under this definition, and after controlling for two-way fixed effects and institutional variables, left strength still significantly reduces IFC.

Column (2) of Table 7 reports results after 1% winsorization of all variables to eliminate the influence of extreme values. The effect of left strength on reducing IFC remains significant once two-way FE and institutional controls are included.

Next, since institutional effects may vary dynamically over time, and to address possible noise in the dependent variable, reverse causality, and the need for smoother explanatory variables, we construct three-year lags (average of lagged values from *t–3* to *t–1*) for the explanatory and control variables. The results, shown in Column (3) of Table 7, indicate that left strength continues to significantly reduce IFC, providing evidence that causality runs from left strength to IFC. Appendix Table 14 reports results using a five-year lag window (*t–5* to *t–1*) as an additional robustness check.

Finally, to further rule out reverse causality, we conduct a placebo test by advancing left strength by one period. If IFC were driving left strength, then lagged IFC should predict future left strength, and advanced left strength should explain contemporaneous IFC. Otherwise, the result should be insignificant. Column (4) of

Table 7 shows that one-period-ahead left strength does not explain current IFC. Combined with the lagged results in Column (3), this suggests that the baseline findings are not driven by reverse causality.

Table 7 Other robustness check

VARIABLES	(1) IFC Change_def	(2) IFC Winsor 1%	(3) IFC m1-3	(4) IFC left(F1)
left	-0.076***	-0.070***	-0.098***	-0.002
	(-3.57)	(-3.10)	(-4.37)	(-0.08)
xconst	-0.025	0.006	0.010***	0.006
	(-1.19)	(0.28)	(3.87)	(0.55)
con	-0.004	-0.001	0.000	-0.002
	(-1.48)	(-0.42)	(0.12)	(-0.93)
fdi	-0.000	0.000	0.001	-0.000
	(-0.45)	(0.43)	(1.12)	(-0.94)
trade	-0.000	0.001	0.001	-0.000
	(-0.17)	(1.25)	(1.49)	(-0.75)
lnpgdp	0.075	0.061	-0.008	0.023
	(0.94)	(0.95)	(-0.12)	(0.31)
lnpop	0.047	0.023	-0.039	-0.044
	(0.29)	(0.18)	(-0.63)	(-0.34)
gdp_growth	0.003	0.004**	0.012***	0.002**
	(1.59)	(2.17)	(4.52)	(2.14)
pop_growth	0.064***	0.051***	0.036*	0.045***
	(3.99)	(3.90)	(1.82)	(5.80)
edu	-0.029	0.015	0.007	-0.035
	(-1.57)	(0.58)	(0.32)	(-1.37)
Constant	-0.662	-1.130	0.503	0.867
	(-0.24)	(-0.54)	(0.39)	(0.37)
Observations	606	612	603	446
Country FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

In the main text, we present only the results controlling for political institutions. The results controlling for the other three economic institutional variables remain significant and are reported in Appendix Table 16 and Table 17.

6) Conclusion

This paper examines the effect of left strength, measured by net parliamentary seat share, on IFC across nearly 40 countries worldwide. Following Milanovic (2017) and Ranaldi & Milanovic (2022), we treat the IFC index as a proxy for liberal versus classical capitalism and investigate the determinants of capitalist dynamics. We also engage with the institutional determinism literature, which has steadily expanded its explanatory scope, and ask whether institutions alone can account for distributive outcomes. The empirical results show that, after controlling for traditional institutions, an increase in left net seats significantly reduces IFC. A 10-percentage-point increase in left strength lowers IFC by about 0.0079 units—roughly 7.5% of the sample mean and 3.6% of the standard deviation—indicating substantive economic significance. This finding empirically challenges the view of institutions as the sole determinant, and brings party strength and distribution back into the debate. Building on Acemoglu et al. (2001), we further distinguish political and economic institutions to study heterogeneity in the effect of left strength. The results show that political institutions (executive constraints, *xconst*) weaken the effect of left strength in reducing IFC, while

economic institutions (rule of law/property rights, RL) amplify it. Channel analysis using the pseudo-Gini of capital shows that left strength reduces IFC mainly by lowering capital inequality, thereby fostering capitalist transformation. However, this mechanism is itself constrained by political institutions. On identification, the RDD provides consistent local evidence around the cutoff, while placebo tests with advanced lags are insignificant. The GMM results point in the same direction but are weakened by instrument proliferation and sensitivity to lag length, indicating limited explanatory strength.

This study makes two main contributions. First, it moves beyond the existing literature that focuses only on stylized facts of functional income distribution and capitalist types, by identifying the determinants and exploring the dynamics driving their evolution. Second, it empirically challenges the single-determinism view that institutions alone determine party power and even income distribution. After clarifying the explanatory boundaries of institutions, the paper uses the IFC index as a lens to examine how the effect of left strength in pushing societies toward liberal capitalism is moderated by different institutional settings. Compared with previous work, this study provides a more detailed perspective that complements institutional determinism, functional distribution theory, and power resources theory.

This paper nonetheless has several limitations:

First, the strength of causal identification remains limited. Two-way fixed effects in panel data help address omitted variables, but are insufficient for resolving reverse causality. Beyond fixed effects, we rely on three strategies for cleaner causal inference: GMM, RDD, and lag/lead placebo tests. However, in long panels, both difference and system GMM suffer from instrument proliferation, which weakens identification and makes the estimates less reliable. Although aggregating years into periods partly mitigates this issue, it also ignores within-period shocks (such as the 2008–09 crisis), and thus cannot fully substitute for year effects in GMM. RDD provides clean identification but only for subsamples and a binary treatment definition, so it cannot rule out reverse causality in the baseline regressions. Lagged explanatory variables and lead placebo tests help, but they do not provide identification as strong as instrumental variables or structural models.

Second, the theoretical model remains at the level of mechanism illustration. It does not yet incorporate micro-level heterogeneity or rational choice foundations, which would allow for a more general analysis.

Third, the DPI database covers about 85% of total votes/seats through the three largest governing parties and the largest opposition party, which is sufficient as a proxy in most cases. However, around 15% of votes remain unexplained, which is especially

problematic in multiparty systems such as those in Europe.

To address these limitations, future work could explore potential instrumental variables such as historical union density or the number and intensity of past revolutions. Provided these instruments affect IFC only through left strength, an FE-2SLS approach could deliver cleaner identification. The theoretical model could also be extended by introducing agent heterogeneity in an overlapping generations (OLG) framework, with two types of individuals making rational choices. This would allow us to study how left strength, under different institutional constraints, shapes the distribution of capital income at the micro level. Finally, left strength could be measured using alternative political databases to check robustness against the limitations of the DPI's coverage of party seats and votes.

7) Appendix

1.Here is the definition of IFC with the corresponding LIS variables:

Table 8 IFC_def1 and IFC_def2

Definition 1 (baseline, IFC_def1):	**Definition 2 (robustness, IFC_def2):**
Labor income = $pi11+(2/3)\bullet pi12$	Labor income = $pi11+pi12$
Capital income =	Capital income = $hicapital/nhhmen$

$\left[hicapital + (1/3) \bullet hi12 \right] / nhhmen$	

Where：

Table 9 LIS database variables name

pi11 = wage and salary income,
pi12 = self-employment income,
hicapital = household capital income,
hi12 = self-employment income (household level),
nhhmen = number of household members.

2. For the seat share, we use data from the DPI database. Specifically, we sum the seats of the three largest governing parties and the largest opposition party (*gov1seat + gov2seat + gov3seat + opp1seat*) and divide by the total number of seats (*totalseats*):

$$seat\_share = \frac{gov1seat + gov2seat + gov3seat + opp1seat}{totalseats} \tag{17}$$

Table 10 Proportion of known vote/seat

	(1)	(2)	(3)	(4)	(5)
VARIABLES	N	mean	sd	min	max
prop_known_vote	3,867	0.869	0.149	0	1
prop_known_seat	6,860	0.843	0.201	0	1.228

3. Before adding time fixed effects, we conducted a Hausman test to choose between random effects and fixed effects. The results reject the null in favor of random effects,

indicating that fixed effects are significantly preferred.

Table 11 Hausman test

Test of H0: Difference in coefficients not systematic
chi2(10) = (b-B)'[(V_b-V_B)^(-1)](b-B)=90.62
Prob > chi2 = 0.0000 (V_b-V_B is not positive definite)

4.Baseline regression with cluster-robust sd:

Table 12 Baseline regression with cluster-robust standard errors

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	IFC	IFC	IFC	IFC	IFC	IFC
left	-0.079**	-0.076**	-0.070*	-0.042*	-0.073**	-0.071**
	(-2.40)	(-2.27)	(-2.00)	(-1.98)	(-2.54)	(-2.46)
xconst			-0.000			
			(-0.01)			
RL				-0.101*		
				(-1.74)		
IP					0.010	
					(1.36)	
LO						0.019
						(1.38)
con		-0.001	-0.001	-0.001	-0.001	-0.000
		(-0.40)	(-0.31)	(-0.63)	(-0.34)	(-0.13)
fdi		-0.000	0.000	-0.000	-0.000	-0.000
		(-1.00)	(0.29)	(-0.25)	(-1.06)	(-1.02)
trade		0.001	0.001	-0.000	0.001	0.001
		(1.07)	(0.79)	(-0.94)	(1.08)	(1.03)
lnpgdp		0.019	0.059	0.051	0.008	0.021
		(0.22)	(0.57)	(0.51)	(0.08)	(0.24)
lnpop		0.026	0.034	-0.081	-0.050	-0.029
		(0.14)	(0.15)	(-0.43)	(-0.24)	(-0.14)
gdp_growth		0.004*	0.003	0.005**	0.004*	0.005**

		(2.00)	(1.35)	(2.38)	(1.78)	(2.38)
pop_growth		0.054**	0.050*	0.049***	0.046**	0.050**
		(2.12)	(1.78)	(2.80)	(2.15)	(2.24)
edu		0.011	0.015	-0.023	0.000	-0.003
		(0.27)	(0.36)	(-0.62)	(0.01)	(-0.07)
Constant		-0.852	-1.253	1.162	0.536	0.003
		(-0.26)	(-0.32)	(0.39)	(0.16)	(0.00)
Observations	748	696	612	539	653	653
R-squared	0.413	0.463	0.455	0.202	0.444	0.443
Country FE	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES

Robust t-statistics in parentheses

*** p<0.01, ** p<0.05, * p<0.1

5.Moderation effect with cluster-robust sd：

Table 13 Moderation analysis with cluster-robust standard errors

	(1)	(2)	(3)	(4)
VARIABLES	IFC	IFC	IFC	IFC
left	-0.518***	0.015	-0.077	0.014
	(-3.19)	(0.61)	(-0.66)	(0.14)
xconst	0.027			
	(1.20)			
c.left#c.xconst	0.066**			
	(2.67)			
RL		-0.109*		
		(-1.88)		
c.left#c.RL		-0.058**		
		(-2.08)		
IP			0.010	
			(1.32)	
c.left#c.IP			0.000	
			(0.05)	
LO				0.021
				(1.47)
c.left#c.LO				-0.018
				(-0.88)
con	-0.001	-0.001	-0.001	-0.000
	(-0.33)	(-0.70)	(-0.34)	(-0.19)

fdi	0.000	-0.000	-0.000	-0.000
	(0.26)	(-0.27)	(-1.05)	(-1.03)
trade	0.001	-0.000	0.001	0.001
	(0.69)	(-0.80)	(1.08)	(1.08)
lnpgdp	0.056	0.068	0.008	0.020
	(0.55)	(0.67)	(0.08)	(0.22)
lnpop	0.025	-0.087	-0.050	-0.028
	(0.11)	(-0.47)	(-0.24)	(-0.14)
gdp_growth	0.003	0.005**	0.004*	0.005**
	(1.46)	(2.24)	(1.78)	(2.36)
pop_growth	0.047*	0.050***	0.046**	0.051**
	(1.69)	(2.92)	(2.20)	(2.25)
edu	0.023	-0.026	0.000	-0.006
	(0.52)	(-0.68)	(0.01)	(-0.17)
Constant	-1.297	1.121	0.538	0.032
	(-0.34)	(0.40)	(0.16)	(0.01)
Observations	612	539	653	653
R-squared	0.460	0.210	0.444	0.445
Country FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Robust t-statistics in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

6. The baseline regressions using the average of explanatory variables lagged from one to five periods (*t–1 ~ t–5*) yield the following results:

Table 14 Baseline regression with the mean of the explanatory variable lagged one to five periods

	(1)	(2)	(3)	(4)
VARIABLES	IFC	IFC	IFC	IFC
m5_left	-0.134***	-0.089***	-0.117***	-0.118***
	(-5.14)	(-4.77)	(-3.93)	(-4.63)
m5_xconst	0.002			
	(1.14)			
m5_RL		-0.133***		
		(-3.82)		
m5_IP			0.019***	
			(2.80)	
m5_LO				0.028*

				(1.99)
m5_con	-0.000	-0.001	-0.000	0.001
	(-0.00)	(-0.47)	(-0.22)	(0.25)
m5_fdi	0.001	-0.001	0.000	0.000
	(1.20)	(-0.65)	(0.07)	(0.29)
m5_trade	0.001**	-0.000	0.002***	0.001**
	(2.48)	(-0.14)	(3.20)	(2.31)
lnm5_pgdp	-0.022	-0.021	-0.120	-0.105
	(-0.35)	(-0.32)	(-1.28)	(-1.29)
lnm5_pop	-0.037	-0.121	-0.122	-0.086
	(-0.66)	(-1.02)	(-1.40)	(-1.13)
m5_gdp_growth	0.014***	0.012***	0.008**	0.012***
	(4.49)	(4.73)	(2.18)	(3.35)
m5_pop_growth	0.018	0.035*	0.012	0.026
	(0.90)	(1.86)	(0.65)	(1.30)
m5_edu	-0.002	-0.013	-0.017	-0.020
	(-0.11)	(-0.38)	(-0.68)	(-0.75)
Constant	0.705	2.555	2.863*	2.071*
	(0.57)	(1.32)	(1.90)	(1.76)
Observations	597	508	639	639
Country FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

7. Moderation effect using lagged averages of explanatory variables (t–1 to t–3 and t–1 to t–5):

Table 15 Moderation effect with the mean of the explanatory variable lagged one to three and one to five periods

VARIABLES	(1) IFC m1-3	(2) IFC m1-3	(3) IFC m1-5	(4) IFC m1-5
left	-0.330***	-0.019	-0.444***	-0.026
	(-4.00)	(-0.57)	(-3.87)	(-1.08)

xconst	0.002		0.010***	
	(0.43)		(3.46)	
c.left#c.xconst	0.036***		0.048**	
	(2.90)		(2.68)	
RL		-0.113***		-0.134***
		(-3.52)		(-3.94)
c.left#c.RL		-0.053***		-0.070***
		(-3.97)		(-8.22)
con	-0.001	-0.001	-0.001	-0.002
	(-0.32)	(-0.68)	(-0.56)	(-1.08)
fdi	0.001	-0.000	0.001	-0.001
	(0.95)	(-0.92)	(1.00)	(-0.71)
trade	0.001	-0.000	0.001**	0.000
	(1.33)	(-0.54)	(2.09)	(0.01)
lnpgdp	-0.037	-0.008	-0.047	-0.013
	(-0.60)	(-0.17)	(-0.77)	(-0.23)
lnpop	0.006	-0.096	-0.022	-0.114
	(0.09)	(-0.96)	(-0.39)	(-0.95)
gdp_growth	0.012***	0.011***	0.013***	0.011***
	(4.46)	(4.33)	(4.18)	(4.15)
pop_growth	0.029	0.041**	0.012	0.034*
	(1.45)	(2.52)	(0.57)	(1.84)
edu	0.012	-0.018	0.006	-0.009
	(0.52)	(-0.53)	(0.29)	(-0.26)
Constant	0.156	2.027	0.689	2.406
	(0.12)	(1.28)	(0.61)	(1.20)
Observations	603	511	597	508
Country FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

8.Robustness checks—including an alternative IFC definition, 1% winsorization, baseline regressions with explanatory variables averaged over t–1 to t–3, and placebo tests using IFC at t+1 while controlling for other institutional factors—confirm the

stability of the results.

Table 16 Other robustness check with other institutional control variables 1

VARIABLES	(1) IFC	(2) IFC	(3) IFC	(4) IFC	(5) IFC	(6) IFC
	Change_def			Winsor 1%		
left	-0.100**	-0.104***	-0.099***	-0.039*	-0.072***	-0.070***
	(-2.43)	(-3.42)	(-3.34)	(-2.02)	(-2.94)	(-2.93)
RL	-0.047			-0.092***		
	(-0.59)			(-3.11)		
IP		0.018***			0.010***	
		(3.14)			(2.73)	
LO			0.040***			0.021
			(2.89)			(1.64)
con	-0.002	-0.005**	-0.004*	-0.002	-0.000	0.000
	(-0.45)	(-2.29)	(-1.87)	(-0.89)	(-0.15)	(0.11)
fdi	-0.000	-0.000	-0.000	-0.000	-0.000	-0.000
	(-1.13)	(-0.43)	(-0.44)	(-0.93)	(-0.15)	(-0.08)
trade	0.000	0.000	0.000	-0.001	0.001	0.001
	(0.42)	(0.81)	(0.71)	(-1.26)	(1.61)	(1.51)
lnpgdp	0.066	0.008	0.035	0.071	0.015	0.030
	(0.53)	(0.07)	(0.35)	(1.37)	(0.24)	(0.51)
lnpop	0.331	0.002	0.048	-0.115	-0.042	-0.024
	(1.29)	(0.01)	(0.28)	(-0.80)	(-0.40)	(-0.23)
gdp_growth	0.006**	0.003	0.006**	0.006***	0.005**	0.006***
	(2.36)	(1.23)	(2.15)	(3.46)	(2.37)	(3.59)
pop_growth	0.051***	0.060***	0.066***	0.043***	0.047***	0.051***
	(5.41)	(4.50)	(6.45)	(7.41)	(4.80)	(6.38)
edu	-0.079***	-0.042	-0.047	-0.016	0.003	0.000
	(-2.97)	(-1.47)	(-1.57)	(-0.66)	(0.13)	(0.01)
Constant	0.000	0.372	-0.761	0.000	0.000	0.000
	(.)	(0.13)	(-0.28)	(.)	(.)	(.)
Observations	529	631	631	539	653	653
Country FE	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

Table 17 Other robustness check with other institutional control variables 2

VARIABLES	(1) IFC	(2) IFC	(3) IFC	(4) IFC	(5) IFC	(6) IFC
		m1-3			left(F1)	
left	-0.068**	-0.087***	-0.082***	-0.010	-0.007	-0.015
	(-2.88)	(-3.71)	(-3.76)	(-0.33)	(-0.28)	(-0.55)
RL	-0.110***			-0.081**		
	(-3.39)			(-2.45)		
IP		0.013**			0.012**	
		(2.03)			(2.65)	
LO			0.032**			0.007
			(2.44)			(0.52)
con	-0.001	-0.000	0.000	-0.002	-0.003	-0.003
	(-0.37)	(-0.11)	(0.22)	(-0.82)	(-1.48)	(-1.33)
fdi	-0.000	0.000	0.000	-0.000	-0.000	-0.000
	(-0.91)	(0.17)	(0.24)	(-0.58)	(-1.07)	(-0.95)
trade	-0.000	0.001*	0.001	-0.001*	-0.000	-0.000
	(-0.67)	(2.02)	(1.59)	(-1.99)	(-0.59)	(-0.75)
lnpgdp	-0.018	-0.107	-0.102	0.025	-0.031	-0.010
	(-0.36)	(-1.16)	(-1.26)	(0.39)	(-0.38)	(-0.13)
lnpop	-0.095	-0.100	-0.090	-0.079	-0.069	-0.031
	(-0.93)	(-1.47)	(-1.37)	(-0.51)	(-0.55)	(-0.25)
gdp_growth	0.012***	0.009***	0.012***	0.004**	0.002	0.003
	(4.84)	(2.89)	(4.87)	(2.39)	(1.14)	(1.66)
pop_growth	0.042**	0.034*	0.043**	0.054***	0.048***	0.053***
	(2.54)	(1.74)	(2.26)	(7.38)	(7.15)	(8.32)
edu	-0.020	-0.007	-0.006	-0.086***	-0.038	-0.039
	(-0.59)	(-0.30)	(-0.22)	(-3.26)	(-1.36)	(-1.45)
Constant	2.092	0.000	0.000	0.000	0.000	0.000
	(1.32)	(.)	(.)	(.)	(.)	(.)
Observations	511	650	650	434	482	482
Country FE	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES

Robust t-statistics (Driscoll–Kraay) in parentheses.

*** p<0.01, ** p<0.05, * p<0.1

9.Wide and narrow bandwidth of RDD estimate.

Table 18 RDD bandwidth robustness check

VARIABLES	(1)	(2)	(3)	(4)
	Narrow bandwidth		Wide bandwidth	
	OLS	Year FE	OLS	Year FE
	First Stage Regression			
	inpower	inpower	inpower	inpower
vote_win	0.425***	0.539***	0.591**	0.633***
	(2.669)	(3.10)	(2.169)	(3.75)
	Second Stage Regression			
	IFC	IFC	IFC	IFC
inpower	-0.551**	-0.485*	-0.317***	-0.350*
	(-2.10)	(-1.98)	(-4.105)	(-1.83)
vote_win		0.055*	0.021	0.027
		(1.90)	(0.70)	(1.69)
Observations	632	138	632	194

10. Placebo tests with shifted RDD cutoffs.

Table 19 RDD cut-off placebo check

VARIABLES	(1)	(2)	(3)	(4)
	Cutoff+3	Cutoff+6	Cutoff-3	Cutoff-6
	First Stage Regression			
	inpower	inpower	inpower	inpower
vote_win	-1.510***	-0.194***	0.395***	-0.072
	(-3.17)	(-3.06)	(5.78)	(-0.62)
	Second Stage Regression			
	IFC	IFC	IFC	IFC
inpower	0.168	0.875	0.115*	-0.747
	(0.68)	(1.31)	(1.79)	(-0.64)
Observations	632	632	632	632

z-statistics in parentheses

*** p<0.01, ** p<0.05, * p<0.1

11. RDD continuity test at the cutoff (McCrary density test)

Table 20 McCrary density check

Variable (running variable)	Vote_win	Cut off (0)
Method	T	P>T
Robust	-0.6544	0.5129

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