Financial development, income inequality and governance institutions

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Abstract

The paper investigates empirically how governance institutions mediate the link between financial development and inequality. To this aim, we assemble a dataset of 48 middle- and high-income countries for the period 1996-2014. Results, obtained by means of instrumental variables dynamic panel data models, reveal that financial development is pro-inequality; however, the strength of the relationship is attenuated in contexts with stricter control of corruption, better regulatory quality, political stability and rule of law. Institutional domains less directly related to the market economy – political voice and accountability and government effectiveness – do not play any mediating role.

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1. Introduction

The global financial crisis exploded in 2008 and its devastating consequences on the real economy have revived the interest of the scientific community on the effects and the risks associated to the development of the financial sector. The widespread and dominant idea, maintained by an influential body of literature (see, for example, Ronald I. McKinnon 1973; Edward S. Shaw 1973; Raghuram G. Rajan and Luigi Zingales 1998; Robert G. King and Ross Levine 1993; Ross Levine and Sara Zervos 1998) that financial liberalisation should be promoted in order to trigger financial development and, consequently, economic growth, has been exposed to radical criticism. It has been shown that not always financial development has supported economic growth; rather, it has often increased macroeconomic fragility, instability and inequality (see, e.g., Enrique G. Mendoza and Marco E. Terrones 2008; Selim Elekdag and Yiqun Wu 2011; Carmen M. Reinhart and Kenneth S. Rogoff 2008; Moritz Schularick and Alan M. Taylor 2012; Joseph E. Stiglitz 2012).

Our paper adds to the empirical literature on the impact of financial development on income inequality by shedding light on whether and how specific institutional dimensions can affect this relationship. In particular, we investigate to what extent the direction and magnitude of the distributive effects of financial development are mediated by the quality of different governance dimensions. To this aim, we assemble a dataset of 48 middle- and high-income countries observed over the period 1996-2014 and employ empirical methods able to simultaneously account for endogeneity issues and the persistence over time of inequality. Such empirical approach is a first contribution of our paper to the existing knowledge, since these aspects have often been not fully accounted for by the empirical literature, despite their potential detrimental consequences on correct identification and inference. However, our main contribution lies on focusing on the role of specific institutional settings related to political governance (rule of law, control of corruption, political stability, regulatory quality, voice and accountability, government effectiveness), whereas most of previous studies make use of aggregate indicators or of single components.

Our empirical results support the side of the literature for which financial development increases income inequality; however, we find that better quality of some specific governance indicators (stricter control of corruption, better regulatory quality, political stability and rule of law) is able to mitigate the magnitude of this pro-inequality effect.

The remainder of the paper is organised as follows: in the next section we summarize the existing theoretical and empirical literature on the effects of financial development on inequality and how they can be affected by institutional settings. In section 3 we present our dataset and some descriptive statistics (3.1) and describe the empirical models and methods used in the econometric analysis (3.2). Section 4 presents and discusses the results obtained, whereas section 5 draws policy implications and concludes.

2. Financial development, income inequality and institutions

The first contributions on the specific link between financial development and income inequality date back to the early 1990s; the following decades marked the development of an extensive body of knowledge which has provided a variety of outcomes (see Jacob De Haan and Jan-Egbert Sturm 2017, Table A1 for a complete review). Most of the literature supports the idea that a more developed financial sector reduces income inequality by allowing the poor, previously excluded from borrowing, to gain access to credit and invest in human and physical capital assets that trigger their income (e.g., Abheijit V. Banerjee and Andrew F. Newman 1993; Oded Galor and Joseph Zeira 1993; Thorsten Beck, Asli Demirgüç-Kunt and Ross Levine 2007; George R.G. Clarke, Lixin Xu and Heng-fu Zou 2006; Rajen Mookerjee and Paul Kalipioni 2010; Luca Agnello, Sushanta K. Mallick and Ricardo M. Sousa 2012; Shigeyuki Hamori and Yoshihiro Hashiguchi 2012; Takuma Kunieda, Keisuke Okada and Akihisa Shibata 2014; Ruixin Zhang and Sami B. Naceur 2019;). Conversely, the literature claiming that financial deepening widens income inequality relies on the effects produced on the intensive margin. When the variety and the quality of financial services increase, rather than broadening access to credit to those previously excluded who still lack sufficient collaterals and credit history, financial development is beneficial to those (the rich) already enjoying the potential of financial services, hence widening the distribution of income (Florence Jaumotte, Subir Lall and Chris Papageorgiou 2013; Jie Li and Han Yu 2014; Era Dabla-Norris et al.2015; Oliver Denk and Boris Cournéde 2015).

A third strand of the literature supplies evidence of a non-linear link between financial development and inequality, which materializes either as an inverted U-shaped relationship (i.e., inequality decreases only after a certain threshold of financial development: see Jeremy Greenwood and Boyan Jovanovic 1990; Dong-Hyeon Kim and Shu-Chin Lin 2011), or as a U-shaped relationship (e.g., Hui-Boon Tan and Siong-Hook Law 2012). In recent years many authors have attempted to provide direct empirical evidence on the idea that this non-linearity is commanded by institutional factors. Rajan and Zingales (2003) were among the first to put forward this idea: when institutions are weak, *de jure* political representation is dominated by *de facto* political influence and this enables established interests (of the affluent) to command access to credit for others, therefore benefitting from financial development exclusively or more than the poor. On the contrary, when institutions are strong enough to guarantee access to credit to those previously excluded, they are enabled to build capital assets and progress along the income ladder. A significant body of literature has also emphasised how specific governance institutional settings affect the way financial sector functions (e.g. Beck, Demirgüç-Kunt and Levine 2005, 2006; Ross Levine 2005, among others) and, in turn, shape income inequality. A good quality of governance is indeed crucial to prevent socially intolerable use of power either by ordinary individuals or by ruling elites (Amir N. Licht, Chanan Goldschmidt and Shalom H. Schwartz 2007). For example, effective rule of law in terms of security of property rights and contract enforceability, improves the functioning of financial markets (Levine 2005); weak enforcement of private contracting, on the contrary, undermines financial access (Stijn Claessens and Enrico Perotti 2007), because it is limited by poor investors protection (La Porta et al.1998). Similarly, since lobbying activities of established wealthier groups may limit access to funding to others (Daron Acemoglu and James A. Robinson 2013), financial access is higher in countries with better political accountability and stricter control of corruption. In more corrupted contexts, established wealthy groups, by needing less external finance, may lobby against the investor protection, hence distorting the allocation of bank credits and limiting access to finance (Enrico Perotti and Paolo Volpin 2007).

On the empirical ground, a number of papers offer interesting insights on the variety of approaches and results that can be obtained when exploring the idea that the quality of institutions conditions the link between financial development and inequality. Some contributions focus on aggregate measures of institutional quality; others on single components of governance indicators (e.g. law and order and transparency, control of corruption, rule of law). Manthos D. Delis, Iftehkar Hasan, and Pantelis Kazakis (2014), by applying 2SLS and GMM-system methods on a sample of eighty-four countries observed from 1997 to 2005, find that banking sector liberalisation decreases inequality only in developed countries with stronger institutions (better law and order and higher transparency). Using panel data on eighty-one countries for a broader period (1985-2010), Siong H. Law, Hui Boon Tan and W.N.W. Azman-Saini (2014) investigate the effect of financial development on inequality and, by means of static threshold regression methods, identify the possible mediating role of an aggregate institutional measure (the sum of five different governance indicators: corruption, law and order, bureaucratic quality, government repudiation of contracts, and risk of expropriation). They observe that the financial development-inequality nexus is insignificant until a certain threshold of institutional quality is reached; beyond this threshold, financial development reduces income inequality. Samuel Adams and Edem K. Klobodu (2016), on the other hand, focus on how control of corruption conditions the link between financial development and inequality for twenty-one sub-Saharan African countries over the period 1985-2011. By employing the pooled mean group estimator (PMG), they find that financial development boosts inequality; however, the inclusion of an interaction terms reveals that a stricter control of corruption mitigates the pro-inequality effect of financialisation. PMG estimator is employed also by Wang Chen and Takuji Kinkyo (2016) to investigate the role of financial development on inequality for a sample of eighty-eight countries from 1961 to 2012. After having summarized, by means of principal component analysis, six governance indicators into a single measure of governance quality, they find that financial development reduces inequality by providing inclusive growth in countries with good governance. The opposite occurs in countries with poor governance. De Haan and Sturm (2017) also sum up single indicators (bureaucratic quality, corruption and rule and order) to generate one aggregate measure of the quality of economic institutions, whereas democratic accountability is employed as a proxy for political institutions. Their empirical analysis is based on a panel fixed effects model augmented with interaction terms, applied on year averages of data on 121 countries from 1975 to 2005. They find that the quality of political institutions is able to condition the link between financial liberalization and inequality, but not the one between financial development and inequality.

More recently, Yi-Bin Chiu and Chien-Chiang Lee (2019), by using unbalanced data for the years 1985-2015 for fifty-nine countries and panel smooth transition methods, analyse the non-linear effects of both financial development and country risks on income inequality. Their results suggest that financial development is pro-inequality when economic stability is low and financial and political stability is high. However, the effects are heterogeneous across income levels, since for high income countries financial development reduces inequality when economic and financial country-risk is low. Zhang and Naceur (2019), by using instrumental variable (IV) regression methods on an extensive sample of 143 countries over the period 1961-2011, provide evidence that rule of law is able to enhance the inequality-reducing effect of finance. Lastly, Dong-Hyeon Kim, Joyce Hsieh and Shu-Chin Lin (2019) investigate the role played by democratization on the link between finance and inequality for seventy-seven countries from 1989 to 2011. A distinctive feature of their study is the distinction of financial development into the stock market and the banking components. By means of GMM methods, they find that stock market development mitigates income inequality in less democratic countries. Conversely, banking development tends to exacerbate income inequality, but higher democratization alleviates this pro-inequality effect.

Compared to this body of literature, the main distinctive feature of our study is that we address the role of six distinct types of governance indicators in shaping the financial development-inequality nexus. Chen and Kinkyo (2016) is the only study considering the same set of governance indicators (but aggregated into a single measure). In order to derive more refined policy implications, we consider crucial to dig into the specificities of single and narrower institutional dimensions. Our empirical methods (section 3.2) also allow, compared to previous studies, addressing simultaneously various challenges posed by the data, particularly related to identification issues due to omitted variable bias and endogeneity.

3. Data and methods

3.1 Data

The empirical analysis relies on a balanced panel dataset of 48 countries observed annually over the period 1996-2014. The list of variables used, their definition and source are provided in Table A1 in the Appendix; Table A2 reports the countries included in the sample, grouped according to the World Bank classification (lower-middle, upper-middle and high-income countries).

Our indicator of inequality, the Gini coefficient on disposable incomes (0-100), is obtained from the Standardized World Income Inequality Database (SWIID), assembled by Frederick Solt (2009). SWIID is a combination of data by Luxembourg Income Study (LIS), University of Texas Inequality Project (UTIP) and World Top Incomes Database (WTID). Compared to other available inequality measures, it has the broadest possible coverage of countries and years; also, it guarantees high comparability of inequality statistics in terms of: (i) population covered, geography, age and employment status; (ii) the variables used as a proxy for welfare; (iii) the equivalence scale; (iii) the treatment of various other items, such as nonmonetary income and imputed rents (Solt, 2016).

As for our second main variables of interest, financial development, we follow the majority of the empirical literature (see, for example, Beck et al. 2007; Sebastian Jauch and Sebastian Watzka 2016; De Haan and Sturm, 2017, for a recent review) and use the amount of credit to the private sector relative to GDP from the World Development Indicators – WDI (World Bank 2018). The measure excludes credit issued by central or development banks and also credits to public sector; it includes credit to individuals and enterprises from banks and other financial corporations (finance and leasing companies, money lenders, insurance corporations, pension funds and foreign exchange companies). Compared to alternative measures, such as bank loans to the private sector, it has the advantage of including credit provided by non deposit taking institutions, which started to play a relevant role in the last decades (Elekdag and Wu 2011).

The third key quantitative information for our work consists of the institutional governance indicators provided by the World Bank Worldwide Governance Indicators - WGI (Daniel Kaufman, Aart Kraay and Massimo Mastruzzi 2011). The informative extent of the dataset, despite having received some critiques (e.g., Laura Langbein and Stephen Knack 2010), is largely acknowledged in economics and political sciences and its employment in academic research is widespread (see, for example, Alberto Chong and Mark Gradstein 2007; Licht et al. 2007; Edinaldo Tebaldi and Ramesh Mohan 2010; Ralph De Haas and Neeltje Van Horen 2012; Gai Wolfsfeld, Elad Segev and Tamir Sheafer 2013; Marcel Fratzscher, Philipp J. König and Claudia Lambert 2016). The indicators range from -2.5 to 2.5 with higher values corresponding to better quality and report on six main domains: voice and accountability (VA), control of corruption (CC), regulatory quality (RQ), political stability and absence of violence (PV), government effectiveness (GE) and rule of law (RL). Following for example of Chen and Kinkyo (2016) or Law et al. (2014), we also compute an aggregate governance indicator (GOV) by averaging the six indicators.

Inequality is a multifaceted phenomenon and the distribution of income is the observable outcome of the effect of many interrelated economic, social and political forces. As a consequence, to correctly identify the impact of the variables of interest, the empirical model needs to minimize the probability of omitting any important explanatory variable and therefore running into biased estimates of the core parameters. To this aim, besides using fixed effects and a dynamic specification (see section 3.2), we include an array of control variables that the literature has identified as key determinants of inequality. Per capita income (in log), GDP growth and the share of industrial sector on total value-added are used to capture the effects of the development level, growth and structural change (e.g., Celine Gimet and Thomas Lagoarde-Segot 2011; Kunieda et al. 2014). A dummy variable (*crisis*) controls for the effects of the global crisis and is coded as one for the years 2008 to 2014 and zero for the remaining ones. Inflation (consumer price index – CPI) and government spending over GDP are included as indicators for macroeconomic stability and for the degree of state intervention via redistributive expenditures (as in Clarke et al. 2006; Kim and Lin 2011), respectively. Macroeconomic instability, particularly price instability, can be pro-inequality since the poor are less able than the rich to protect themselves from high inflation as they hold more cash rather than a variety of financial assets. If the tax and transfers system is aimed at redistributing resources towards low-income groups, a negative relationship between government spending and inequality is to be expected. On the other hand, if the rich have political strength to capture policy makers, reverse effects are likely to occur (see Clarke et al. 2006). Higher government spending may also correspond to an aggregate demand stimulus, often beneficial to low-wage sectors and this way creating downward pressure on inequality.

Table 1. Summary Statistics (48 countries, 1996-2014)

Variable	Mean	Std. Dev.	Min	Max
Gini	37.189	9.099	22.246	54.492
Private credit on GDP (cred)	73.744	50.601	1.385	253.262
(ln) Income per capita (lninc)	9.488	1.144	7.197	11.425
GDP growth (growth)	2.813	3.575	-14.8	16.226
Capital account openness (kaopen)	0.715	0.335	0.000	1.000
Trade openness (trade)	79.089	54.097	15.636	441.604
Government expenditures (govt)	16.514	4.564	4.787	27.935
Human capital (hc)	2.832	0.493	1.452	3.734
Inflation (inf)	7.659	36.983	-4.479	1058.374
Share of Industry on total value added (indust)	30.223	6.648	10.693	57.796
Total factor productivity (tfp)	0.724	0.245	0.186	1.617
Institutions				
Voice and Accountability (VA)	0.576	0.793	-1.749	1.801
Control of corruption (CC)	0.549	1.109	-1.400	2.470
Regulatory quality (RQ)	0.644	0.847	-1.815	2.233
Political stability/absence of violence (PV)	0.215	0.869	-2.374	1.760
Government effectiveness (GE)	0.643	0.969	-1.227	2.437
Rule of law (RL)	0.513	1.022	-1.916	2.100
Governance (GOV – Average of the six)	0.523	0.881	-1.382	1.970
Legal system (legalsys)	6.262	1.776	1.880	9.620
Credit market deregulation (credmark)	8.456	1.310	3.003	10.000
Labour market deregulation (labormark)	5.491	1.366	2.29	9.280

Source: own elaborations (sources of data in Table A1)

James K. Galbraith (2007) argues that within country inequality has been triggered on a big scale by macro and global forces rather than country-specific, micro effects; an extensive body of literature, based on classical and more recent trade theories has studied the effects of globalisation on income distribution, mainly due to effects on relative demand of skills (Hongyi Li, Lyn Squire and Heng-fu Zou 1998; Robert J. Barro 2000; Adrian Wood 1995; David Dollar and Aart Kraay 2004; Nathalie Chusseau, Michel Dumont and Joël Hellier 2008). As standard in the literature (see, e.g., De Haan and Sturm 2017 and the many references cited therein), we use as a proxy for globalisation a metric of trade openness - the sum of exports and imports to GDP. Following Silke Bumann and Robert Lensink (2016) and Davide Furceri and Prakash Loungani (2018), we also include among the regressors the capital account openness index constructed by Menzie D. Chinn and Hiro Ito (2008), defined as a measure of *de jure* capital openness and globalization of finance.

The effects of the race between technological change and skilled labour supply on the labour market (the so-called skilled-bias technological change hypothesis) are other factors commonly identified of possible determinants of inequality (Lawrence F. Katz and David H. Autor 1999; Maarten Goos and Alan Manning 2007; avid H. Autor, Frank Levy and Richard J. Murnane 2003; Daron Acemoglu 1998; Robert J. Barro and Jong W. Lee, 2013). To account for these drivers, we use a measure of total factor productivity (TFP) and a metric for human capital, both from the Penn World Tables. The index of human capital combines information on average years of schooling and return to education based on the indicators constructed by Barro and Lee (2013) and Daniel Cohen and Laura Leker (2014).

Apart from the WGI governance indicators, we consider in our empirical analysis three more strictly economic institutional dimensions: labour market (de)regulation, legal system & property rights and credit market (de)regulations, supplied by the Fraser Institute (Gwartney et al. 2009). The first one is included as a control variable, as labour market deregulation (in the form of lower employment protection, less unionisation, more decentralised and less coordinated wage bargaining) has been shown to directly lead to greater inequality (David Card and John E. DiNardo 2002; David Card, Thomas Lemieux and W. Craig Riddell 2004; Tito Boeri and Pietro Garibaldi 2007; Cristiano Perugini and Fabrizio Pompei 2017; Daniele Checchi and Cecilia Garcia-Penalosa 2010). The remaining two metrics are used as instrumental variables in order to account for the potential endogeneity of financial development with respect to inequality (see following section).

3.2 Empirical models and methods

Our empirical strategy starts from a baseline estimation of the drivers of income inequality, which includes our main variable of interest (financial development). For a correct identification we need first of all to reduce to a minimum the exposure of the model to omitted variable bias. To this aim, and at the cost of restricting the time and geographical coverage of the sample, we include in the baseline estimation all proxies for the factors affecting income inequality we were able to collect as well as the lagged level of the dependent variable. This is functional to account for fact that within country income inequality is characterized by high inertia and can be viewed as a time-persistent phenomenon (see, among others, Dilip Mookherjee and Debraj Ray 2003). We therefore consider the following baseline dynamic model:

$$Gini_{i,t} = \alpha_i + \tau_t + \rho Gini_{i,t-1} + \beta_1 Cred_{i,t} + \beta_2 Inst_{i,t} + X_{i,t}{}'\gamma_n + Trend_t + \varepsilon_{i,t} \qquad [1]$$

where subscripts *i* and *t* refer to countries and years, respectively (*i = 1, ..., 48; t = 1996, ..., 2014)*; α_i and τ_t are country and time specific effects and $\varepsilon_{i,t}$ is the error term. *Cred* (private credit on GDP) is the proxy for financial development, our main variable of interest, and *Inst* is the metric describing the quality of institutions. This variable will be either the summary measure of quality of governance (GOV) or one of its six components described in section 3.1. β_1 and β_2 measure the direct effect of financial development and governance quality on income inequality, respectively. $X_{i,t}$ is the control variables matrix and γ_n is the vector of associated coefficients. A time trend variable is also included in order to prevent a possible spurious relation between variables driven by a common time pattern[1]. A major

[1] As a preliminary step we have run panel unit root tests for stationarity in our core variables. In view of the macroeconomic nature of our dataset, we opted for the Fischer-type test, which performs an Augmented Dickey-Fuller (ADF) unit-root test on each panel's series separately, then combine the p-values to obtain an overall test of whether the panel series contains a unit root. The null hypothesis being tested (using the STATA command *xtunitroot*) is that all panels contain a unit root. For a finite number of panels, the alternative is that at least one panel is stationary. As for the variable *Gini* in our dataset, all four tests rejected the null hypothesis that all the panels contain unit roots. As In Choi's (2001) simulation results suggest that the inverse normal Z statistic offers the best trade-off between size and power, we report its values here (the remaining ones are available upon request). The statistics for *Gini* amounts to -8.0636 (p-value 0.000) when the number of lags is set at 2 and to -7.8237 (p-value 0.000) with three lags. Similarly, for the financial development variable (*Cred*) the corresponding inverse normal Z statistics are -9.4947 (p-value: 0.000) and -10.6037 (p-value 0.000), respectively. The use of alternative tests, such as the Levin–Lin–Chu (LLC) which assumes that all panels share a common autoregressive parameter, equally indicate a rejection of the null hypothesis of a unit-root, for both *Gini* and *Cred* (adjusted T statistic -4.1140 [0.000] and -3.3638 [0.000], respectively).

advantage of the panel approach is the inclusion of country and year specific effects, which account for unobservable or imperfectly measured drivers of income inequality across time and space.

In order to capture the mediating role of institutions on the effects that financial development produces on inequality, the model is augmented with an interaction term between *Cred* and *Inst* as follows:

$$Gini_{i,t} = \alpha_i + \tau_t + \rho Gini_{i,t-1} + \beta_1 Cred_{i,t} + \beta_2 Inst_{i,t} + \beta_3 Cred_{i,t} \cdot Inst_{i,t} + X_{i,t}'\gamma_n + Trend_t + \varepsilon_{i,t} \qquad [2]$$

While β_1 and β_2 measure again the direct (main) effects of financial development and the quality of institutions on inequality, respectively, β_3 measures how institutional settings impact on the effect of financial development on inequality. The inclusion of the interaction term (as done by Adams and Klobodu 2016 and De Haan and Sturm 2017, in similar contexts) is a relatively simple and straightforward approach, as the significance and the sign of β_3 directly indicate whether and how the effect of financial development on inequality (measured by β_1) changes due to differences in institutional quality. In addition, it can be employed easily in econometric models able to account for the many challenges posed by the estimation of our model (in particular, as explained below, the presence of the lagged dependent variable and endogeneity issues). Other approaches, such as panel threshold (Bruce E. Hansen 1999) or panel smooth transition regression (González et al. 2017), offer important informative advantages but empirical models and estimation tools needed for our analysis (i.e., dynamic panel threshold model with endogeneity, as in Myung H. Seo and Yongcheol Shin 2016) are still in an early stage of development. Hence, we prefer here to use different approaches, now standard in the empirical literature, able to address satisfactorily the problems mentioned and to compare the outcomes, particularly the relationships of interest, step-by-step.

As a first pass, we estimate a static standard fixed effects model (with no lagged dependent variable as a regressor), followed by its dynamic specification; this way, we show how much not accounting for the persistence over time of inequality affects the specification of the model. However, the presence of the lagged dependent variable, due to its potential correlation with the composite error $\alpha_i + \varepsilon_{i,t}$, may lead to inconsistent parameter estimates also when country heterogeneity is accounted for by means of conventional fixed- or random-effects estimators (Badi H. Baltagi 2001). This is due to the so-called dynamic panel bias—although if T (the time dimension) is sufficiently large, as in our case, this becomes insignificant. Under such circumstances, a standard straightforward fixed-effects estimator can be employed (David Roodman 2009). Yet, this approach fails to address the problems of endogeneity due to potential reverse causality, which is a crucial issue here. Our interest lies indeed in correctly identifying the effect of financial development on inequality and a large body of literature (reviewed, for example in Cristiano Perugini, Jens Hölscher and Simon Collie 2016) has emphasised that income inequality might be one of the drivers of the expansion of credit, through various channels. Similarly, inequality can contribute shaping various types of institutional settings (see Chong and Gradstein 2007; Philip Keefer and Stephen Knack 2002; William Easterly 2001). Despite other regressors are at risk of endogeneity too, given the purpose of the paper we focus on the treatment of the potential endogeneity of the variable (*Cred*) and of the institutional indicators, using two different approaches. The first relies on a fixed effect instrumental variable estimator based on the Lars P. Hansen (1982) original Generalized Methods of Moments (GMM), which allows the instrumentation of variables at risk of endogeneity, as well as providing standard errors robust to heteroskedasticity and autocorrelation. As instruments, we use a mix of internal (lagged levels of *Cred*) and external variables. Standard (Hansen) tests of overidentifying restrictions, reported in the relevant tables, indicate the validity of instruments. Based on the available literature, as external instruments we use information related to the general legal setting and to the credit market institutional environment. An extensive literature points out that legal systems with better enforcement of property rights and protection of legal rights of investors promote financial development (see La Porta et al. 1998, 2000; Beck and Levine, 2008; Kim and Lin 2011). Similarly, financial liberalisation is normally, although not necessarily (Elkhuizen et al. 2018), correlated to financial development (Chinn and Ito 2006; Michael D. Bordo and Christopher Meissner 2012). We use here two indicators available in the Economic Freedom of the World – EFW – database (see Table A1), which range from 0 to 10 in ascending order of progress/deregulation. The first one (*legalsys*) is defined as Legal System and Property Rights and summarizes information on the soundness of the legal system with reference to: judicial independence, impartial courts, protection of property rights, military interference in rule of law and politics, integrity of the legal system, legal enforcement of contracts, regulatory restrictions on the sale of real property, reliability of policy and business costs of crime. This variable, compared to other widely used ones (such as legal origins, see La Porta et al.1997), has the advantage of varying over time, therefore accounting for developments in the field across the period considered (see Armour et al., 2009), besides being of a less problematic use in a panel data setting. The second indicator (*credmark*) is the credit market (de)regulation index, widely used in the existing empirical literature (e.g., Domenico Giannone, Michele Lenza and Lucrezia Reichlin 2011; John W. Dawson 2006; Petar Stankov 2012) and summarizes four dimensions of liberalisation related to: (i) ownership of banks; (ii) foreign bank competition; (iii) private sector credit; (iv) interest rate controls/negative interest rates.

As a check for the robustness of the results obtained we also approach the endogeneity issue using the System GMM estimation techniques (Manuel Arellano and Olympia Bover 1995; Richard Blundell and Stephen Bond 1998). The GMM-sys estimator employs as instruments the lagged values of the endogenous explanatory variables. Variables in levels are instrumented with lagged first differences; then, in order to consider these additional moments as valid instruments for levels, it is require the identifying assumption that past changes of the explanatory variables are uncorrelated with current errors in levels, which include fixed effects (Roodman 2009). The validity of the moment conditions can be verified by means of the test of overidentifying restrictions proposed by Denis J. Sargan (1958) and Hansen (1982) and by testing the null hypothesis of no second order serial correlation in the error term. The GMM-Sys estimator has the advantage of allowing instrumentation of endogenous variables with internal lags but it is designed for large N small T panels to deal efficiently with dynamic panel bias. Its employment to dataset like ours (in which the time dimension is relatively large) may cause a natural proliferation of the number of instruments (Roodman 2009; Clive G. Bowsher 2002). However, system GMM estimation allows some flexibility by means of several specification choices. In particular, given the structure of our panel, we use the one-step estimator and correct the standard errors to account for small-sample bias and heteroskedasticity, by applying the Huber and White robust variance estimator. Furthermore, to address the problem of instrument proliferation, we use a combined strategy obtained by collapsing instruments (i.e., creating one instrument for each variable and lag distance only, with zero substituted for any missing values) and restricting the number of lags used as instruments.

4. Results

Table 2 reports the results of the estimations for our baseline empirical model (equation 1) that includes, among the drivers of income inequality, the aggregate measure of governance quality (GOV). The first column illustrates the outcomes of standard static fixed effect estimation; in the second one, the lagged dependent variable is added to the set of regressors. Its significance and positive coefficient indicate that inequality is a persisting feature of economic systems. Along with the sharp increase in the explanatory power of the model, this evidence indicates that a static approach would expose the model to severe omitted variable bias issues, which would seriously undermine statistical inference. For these reasons, and different from the existing literature that neglects such aspects (e.g., Law et al.

2014), we opt for a dynamic specification throughout our empirical analysis. Columns 3 in Table 2 displays the results obtained with the Hansen GMM estimator described in section 3.2, which is able to address issues related to the dynamic panel bias and to potential endogeneity of the regressors of main interest here (*cred* and *GOV*). The tests at the bottom of the Table indicate that the variables hypothesised as endogenous must indeed be treated as such and that the instruments used in the first stage (their first lags and variables related to the legal system – *legalsys* - and the regulation of the credit market - *credmark*) are valid. Outcomes of the robustness check implemented by means of the GMM-sys estimators are reported in Table A3 in the Appendix.

With reference to the variables used as controls, Table 2 shows a remarkable stability of the outcomes obtained with different estimators; the signs of the coefficients, where significant, are generally consistent with the existing literature and our expectations. The level of development (approximated by the log of per capita GDP) is linearly associated to lower inequality (its squared term always turned out being not significant and originating multicollinearity issues, it is therefore not included in the model). Similarly, growth is pro-poor. Variables measuring the degree of globalisation (trade and capital account openness) are not significant, whereas a stronger role of the state is associated to lower inequality. A higher endowment of human capital seems able to reduce income disparities, consistent with the idea that larger shares of highly educated workers reduce the dimension of inequality related to wage differences across education groups. Similarly, a larger share of the industrial sector on total value added is associated to lower inequality; this is consistent with the evidence that wage inequality in the secondary sector (manufacturing in particular) is normally lower than in services or agriculture due, among other things, to a stronger role of unions and more pervasive wage setting institutions (higher wage coordination and centralisation). The variable describing labour market institutional settings explicitly accounts for such aspects and suggests that more deregulated labour markets are conducive to higher inequality, allegedly due to wider wage gaps between different labour market segments. The dummy variable controlling for global crisis has a negative, significant coefficient. Although seemingly counterintuitive, this outcome is not entirely new, as some empirical studies have highlighted how in some contexts a crisis does not necessarily translate into higher inequality. For example, Paul De Beer (2012) found that in roughly half the EU member states inequality declined and poverty rates dropped with the 2008-2010 crisis. Similarly, the World Bank (2016) showed, again with reference to the most recent global crisis, that between 2008 and 2013, the number of countries experiencing declining inequality was twice the number exhibiting widening inequality. Other influential contributions found little or no effects of the crisis on household income inequality, mainly due to stabilization measures and protection guaranteed by tax and benefit systems. On the contrary, the austerity measures that followed produced much greater changes in the income structure (Jenkins et al.2013). The time trend variable is not statistically significant, indicating that a possible linear trend in the dependent variable is probably already captured by the time fixed effects.

As regards the focus of the paper, results in Table 2 place our work on the side of the literature maintaining that financial development increases income inequality (De Haan and Sturm 2017; Chiu and Lee 2019; Jaumotte et al. 2013; Gimet and Lagoarde-Segot 2011), as the coefficient of *cred* (private credit on GDP) always turns out positive and significant[2]. The sign of the aggregate governance indicator (GOV) has a positive sign, indicating that a progress in institutional quality is associated to higher inequality. However, this effect is due to some specific dimensions of

[2] The same outcome holds if we split the sample into two subsamples that include lower-middle and upper-middle income countries on one side (21 countries) and high income on the other (27 countries), based on the World Bank classification described in Table A2 in the Appendix. Results, not reported here but available upon request, indicate that financial development plays a statistically significant pro-inequality effect in both samples, but its magnitude is larger for middle income countries (coefficient of *cred* is 0.008) compared to high income countries (0.002).

political governance. This is confirmed by the evidence presented in the following tables 3 and A3 in the Appendix, in which we also augment the model with the interaction between financial development and governance indicators.

Table 2. Inequality and financial development, baseline model (48 countries, 1996-2014)

	(1)	(2)	(3)
VARIABLES	FE	FE, lagged Gini	IV (GMM)
L.gini		0.948***	0.947***
		(0.009)	(0.010)
cred	0.004	0.003***	0.004***
	(0.003)	(0.001)	(0.001)
INST (GOV)	-0.541	0.249**	0.401***
	(0.423)	(0.112)	(0.147)
lninc	1.662***	-0.390***	-0.445***
	(0.554)	(0.147)	(0.157)
growth	-0.032**	-0.009**	-0.008*
	(0.016)	(0.004)	(0.005)
kaopen	1.501***	0.013	-0.024
	(0.297)	(0.079)	(0.089)
trade	0.018***	0.001	0.001
	(0.004)	(0.001)	(0.001)
govt	-0.063*	-0.028***	-0.027***
	(0.037)	(0.010)	(0.010)
hc	-4.921***	-0.428*	-0.418**
	(0.857)	(0.227)	(0.195)
inf	0.000	-0.000	-0.000*
	(0.001)	(0.000)	(0.000)
indust	-0.155***	-0.035***	-0.033***
	(0.022)	(0.006)	(0.006)
tfp	-3.477***	-0.173	-0.192
	(0.777)	(0.210)	(0.197)
labormark	0.374***	0.092***	0.083***
	(0.079)	(0.021)	(0.020)
crisis	-0.909***	-0.121***	-0.128***
	(0.174)	(0.046)	(0.045)
trend	-0.014	-0.009	-0.006
	(0.029)	(0.008)	(0.007)
Constant	39.603***	7.705***	
	(5.448)	(1.487)	
Observations	912	864	864
R-squared	0.305	0.956	0.956
Adj. R-squared	0.255	0.953	0.953
F-test	26.62	1173.00	916.8
Underid. (Kleibergen_Paap) LM stat			190.000 [0.000]
Weak id. (Kleibergen_Paap) F stat			497.900
Hansen J stat			2.352 [0.125]
Endog. Chi2 stat			10.690 [0.005]

Robust standard errors in round parentheses; p-values in squared parentheses; year effects included; *** p<0.01, ** p<0.05, * p<0.1

In Table 3 we present the instrumental variable estimates of the empirical model augmented with the interaction term (equation 2), carried out by means of the Hansen GMM method. We now treat as potentially endogenous variables, besides *cred* and the institutional variables, also their interactions, using again as instruments their first lags, *legalsys* and *credmark*. The tests at the bottom of the Table confirm the validity of the instruments used in the first stage. The first column of Table 3, besides confirming all results of Table 2, shows that the inequality-enhancing effect of financial development (i.e., the positive and significant sign of *cred*) is mitigated by better quality political governance (negative and significant sign of the interaction term). This evidence corroborates the idea that the magnitude of the effects of financial deepening on income distribution depend on the way the institutional setting affects its functioning and accessibility (Rajan and Zingales 2003). However, the aggregate indicator is too general and in columns 2 to 7 we replicate the model using single dimensions of governance. Results indicate that only some specific domains of governance are able to affect the impact of financial development on inequality. Voice and accountability (VA), as well as government effectiveness (GE) do not play any role; this is probably due to the fact that such dimensions pertain more strictly, respectively, to the political sphere (ability to participate in selecting the government, freedom of expression and association, a free media) or to the quality of services provided by the state (independence of the civil service, quality of policy formulation and implementation). As such, their link to the financial system is probably too weak to let them emerge as conditioning factors. Conversely, the remaining four governance indicators play a mitigating role on the inequality-increasing impact of financial development. Control of corruption (CC), in particular, captures the extent to which the State is able to limit public power being exercised for private gain, including various forms of corruption and "capture" of the state by elites and private interests. The coefficient of the interaction term (see column 3 of Table 3) is negative and significant, indicating that better control of corruption (i.e., a higher value of CC) weakens the pro-inequality effect of financial development. This suggests that when the political system is more vulnerable to the influence of elites (low CC) the pro-inequality effect of financialisation is magnified. This evidence is consistent with some influential literature on the interactions between the power of elites and inequality. In particular Acemoglu (2011), with reference to the origins of the global crisis, suggests that a similar mechanism might be in place exactly with reference to the development of the financial sector. Citing evidence from Larry M. Bartels (2008) and Martin Gilens (2005), he argues that the policies over the last decades were in fact more closely aligned to the preferences of a minority of high-income voters. Particularly, politicians implemented financial deregulation measures favouring influential high-income constituents (many of whom worked in, or directly benefited from, the financial sector). Paul Krugman (2012) puts forward a similar argument, maintaining that agents at the top of the distribution are able to exert political influence to promote policies of financial deregulation in the pursuit of their personal interest.

The governance indicator of regulatory quality (RQ) is related to perceptions of the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development. This is the dimension of governance more strictly and directly related to rules and policies favouring the functioning of a full market economy such as competition policies, absence of price controls, low burden of government regulations, investment and financial freedom. It is therefore not surprising that the main effect of the indicator RQ in column 4 is positive and significant (and is one of the main drivers of the same sign of GOV in column 1); higher inequality is a feature inherent to liberal market economic systems where, due to higher decentralisation, free market forces shape incentives, bargaining power and outcomes (see, for example, Peter A. Hall and David Soskice 2001; Bruno Amable 2003; Gosta Esping-Andersen 1990, and more in general the literature dealing with the possible variety of capitalistic forms). In such contexts, it is possible that a more developed financial system is capable of providing resources and opportunities to a larger number of individuals, including those at the bottom of the distribution. Such inclusiveness, coupled with the efficiency of the financial sector in selecting good and long-term investment projects, might explain the weaker pro-inequality effect of financial development in contexts of better regulatory quality, as suggested by the negative sign of the interaction term in column 4.

Table 3. The mediating role of institutions on the link between financial development and inequality (Hansen IV GMM estimation, 48 countries, 1996-2014)

VARIABLES	(1) GOV	(2) VA	(3) CC	(4) RQ	(5) PV	(6) GE	(7) RL
L.gini	0.953***	0.945***	0.948***	0.943***	0.950***	0.948***	0.948***
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
cred	0.007***	0.005***	0.006***	0.007***	0.005***	0.004*	0.006***
	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)
INST	0.571***	-0.009	0.055	0.717***	0.110	-0.281	0.607***
	(0.163)	(0.125)	(0.123)	(0.107)	(0.099)	(0.352)	(0.131)
cred*INST	-0.003***	-0.001	-0.002***	-0.003***	-0.003***	-0.000	-0.002***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
lninc	-0.607***	-0.398**	-0.428***	-0.748***	-0.455***	-0.273	-0.619***
	(0.162)	(0.168)	(0.162)	(0.159)	(0.158)	(0.226)	(0.159)
growth	-0.007	-0.006	-0.005	-0.006	-0.006	-0.005	-0.006
	(0.005)	(0.005)	(0.005)	(0.004)	(0.005)	(0.005)	(0.004)
kaopen	-0.065	0.057	0.020	-0.139	0.060	0.109	-0.072
	(0.089)	(0.086)	(0.085)	(0.086)	(0.086)	(0.107)	(0.085)
trade	0.001	0.001	0.001	0.002	0.001	0.001	0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
govt	-0.024**	-0.027***	-0.024**	-0.019*	-0.030***	-0.028**	-0.022**
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.011)	(0.010)
hc	-0.362*	-0.484**	-0.453**	-0.333*	-0.368*	-0.417**	-0.418**
	(0.198)	(0.199)	(0.200)	(0.197)	(0.201)	(0.200)	(0.194)
inf	-0.000	-0.000**	-0.000	0.000	-0.000**	-0.000*	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
indust	-0.028***	-0.035***	-0.034***	-0.023***	-0.032***	-0.038***	-0.023***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.009)	(0.006)
tfp	-0.202	-0.127	-0.145	-0.148	-0.095	-0.026	-0.188
	(0.197)	(0.192)	(0.194)	(0.190)	(0.199)	(0.213)	(0.194)
labormark	0.087***	0.100***	0.101***	0.058***	0.091***	0.102***	0.078***
	(0.020)	(0.020)	(0.020)	(0.020)	(0.020)	(0.020)	(0.020)
crisis	-0.139***	-0.115***	-0.119***	-0.165***	-0.103**	-0.093*	-0.139***
	(0.045)	(0.044)	(0.044)	(0.043)	(0.046)	(0.049)	(0.043)
trend	-0.003	-0.012*	-0.012*	0.005	-0.014*	-0.018**	-0.001
	(0.007)	(0.007)	(0.007)	(0.006)	(0.007)	(0.008)	(0.007)
Observations	864	864	864	864	864	864	864
R-squared	0.957	0.956	0.957	0.959	0.956	0.955	0.958
F-test	885.5	867	854.2	867.4	852.5	808.6	884.7
Adj. R-squared	0.954	0.953	0.953	0.956	0.953	0.951	0.954
Underid. (Kl.Paap) LM	208.1	148.8	200.6	195.1	133.1	60.32	201.6
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Weak id. (Kl Paap) F	357.4	217.9	250.2	119.7	77.08	11.95	446.8
Hansen J stat	0.336	1.196	1.114	4.005	2.998	4.738	1.588
	[0.562]	[0.274]	[0.291]	[0.135]	[0.223]	[0.192]	[0.208]
Endog. Chi2 stat	10.490	8.325	4.718	16.03	7.250	10.490	7.116
	[0.015]	[0.040]	[0.194]	[0.001]	[0.064]	[0.015]	0.069]

Robust standard errors in round parentheses; p-values in squared parentheses; year effects included; *** p<0.01, ** p<0.05, * p<0.1

The last two dimensions able to mitigate the inequality-enhancing effect of financial development pertain to political stability (PV) and rule of law (RL). The first one measures the perception of the likelihood of political instability, violence and terrorism; this is a rather general indicator of, among other things, the perception about security of property rights. The second is more directly focused on the aspects of interest here, being a measure of the perception of the extent to which agents have confidence in and abide by the rules of society (likelihood of crime and violence), and in particular the quality of contract enforcement, property rights and of arrangements aimed at administering justice (police and courts). Both indicators are therefore related to the level of perceived uncertainty, but the second one is more narrowly referred to the economic sphere. Their mitigating role on the pro-inequality effects of financial development might be explained by the fact that in conditions of lower uncertainty investment plans tend to be channelled into long-term projects (like accumulation of physical and human capital) able to guarantee stable returns in the future. When credit access is large enough to reach the worse-off segments of the population, this might trigger a convergence of incomes able to lead to a more equitable distribution (see also Cyn-Young Park and Rogelio Mercado 2018, on a similar mechanism linking better rule of law, financial inclusiveness and lower inequality). The positive and significant sign of the main effect of RL might again be related to the fact that institutional settings able to reduce transaction costs address the economic systems towards increasingly decentralised, market-based models, inherently characterised by higher disparities.

In order to check the robustness of our results, we run the same dynamic model using an alternative estimator (GMM-sys), equally able to address endogeneity and dynamic panel bias. As already emphasised, the estimator is designed for small-T large-N panels. When the time dimension becomes relatively large the number of instruments, which in the GMM-sys is quadratic in T, tend to increase and cause several problems (see Roodman 2009), which include a weakening of the test of over identifying restrictions and the over fit of the endogenous variables. Unfortunately, the literature provides little guidance on the maximum number of instruments to be employed, but as a minimally arbitrary rule of thumb the instruments should not outnumber individual units. Also, implausibly good values of the Hansen/Sargan test (approaching a p-value of 1.000) are a clear signal of a too large instruments collection. Results of the estimation of equation 2 by means of GMM-sys are reported in Table A3 in the Appendix. In our case, the minimum number of instruments for which the over identifying restrictions test is satisfied is in all cases around twenty (the *collapse* option available in the STATA command *xtabond2* has been used), so considerably lower than the individual units (48). The Sargan test never approaches implausibly good values. Nonetheless, results should be considered with caution, as some signs of the control variables tend to be unstable to changes in the number of lags used as instruments; in addition, a few control variables, such as *govt* and *indust*, turn out with different signs compared to the results presented in Table 3, for reasons that are difficult to disentangle. Hence, the evidence proposed in Table A3 should only be considered as corroborative (or not) of the results obtained with our preferred method (the Hansen GMM, Table 3), which on the contrary provides stable outcomes. The coefficients of interest for our aims (i.e., the impact of financial development on inequality and the mediating effects of the various institutional dimension) are all confirmed, indicating that they are robust to the use of econometric techniques that can produce significant alterations in the estimates. The only exception is the non-significance of the interaction effect of the political stability/violence (PV) variable (column 5 in Table A3); we have already discussed the fact that its effects, if any, tend to be rather indirect and better captured by the impact of the rule of law (RL), which is instead confirmed as a mitigating factor.

5. Conclusions

In this paper we use a dataset covering 48 countries for the period 1996-2014 to investigate whether the distributive impact of financial development is influenced by some specific institutional settings. To this aim, we employ an array of econometric approaches able to address the various issues posed by the analysis, namely the potential endogeneity of our key variables and the persistence over time of inequality. Our core results are robust to different specifications and indicate that: (i) financial development tends to increase income inequality; (ii) this impact is mitigated when the quality of institutions related to specific governance domains increases. In particular, the pro-inequality effect of finacialisation is weaker the better the quality of those institutional features able to pose the conditions for an efficient functioning of a market economy, i.e.: lower political uncertainty and higher independence of politics from specific interest groups, better protection of property rights, better competitive environments, more efficient enforcement of contracts. It is plausible that, in such frameworks, the financial system develops as a more inclusive device and is therefore better able to channel resources into investment plans that trigger income growth patterns for lower income individuals. On the contrary, when such institutional features are of a poor quality, the inequality enhancing effect of financial development is accentuated.

Policy implications of our results, if confirmed, are not small. A strong role of the financial sector is one of the defining characteristics of a market economy and the core of its main channels of transmission. However, its overall effects on growth and macroeconomic stability are still an open arena for research and political discussion. According to the evidence provided here, unintended effects of financial development on distributive patterns should be added to the list. Our results indicate that policy makers who are willing to narrow down the pro-inequality effects of financial development should provide an adequate institutional environment in specific domains. They are in particular related to the control of corruption and 'capture' of the state by elites and private interests, to the credibility of the government in promoting private sector development, to political stability and to the effectiveness of those institutions enabling markets to play their allocative role and the poor to access finance. Since distributive dynamics have been shown to be themselves related to the stability of economic systems, policy makers should set a socially sustainable pattern of development of the financial sector among their priorities.

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Appendix

Table A1. Variables abbreviations, definitions and sources

Abbreviation	Variable definition	Unit of measurement	Source
gini	Gini coefficient	Scaled from 0 to 100	SWIID
cred	Domestic credit to private sector	% of GDP	WDI
lninc	GDP per capita (in natural logarithm)	constant 2010 US$	WDI
growth	GDP growth	%	WDI
kaopen	Capital account openness	Scaled from 0 to 1	Chinn & Ito (2008)
trade	Trade openness (sum of exports and imports)	% of GDP	WDI
govt	General government final consumption expenditure	% of GDP	WDI
hc	Human capital (average years of schooling + return to education)	Index	PWT
inf	Annual change of consumer price index	%	WDI
indust	Value added of industry	% of GDP	WDI
tfp	Total factor productivity	current PPPs (USA=1)	PWT
crisis	Dummy variable for the global crisis (1 for years 2008 to 2014, 0 otherwise)	Binary	--
VA	Voice and Accountability	Index (-2.5/2.5)	WGI
CC	Control of Corruption	Index (-2.5/2.5)	WGI
RQ	Regulatory quality	Index (-2.5/2.5)	WGI
PV	Political instability and absence of violence	Index (-2.5/2.5)	WGI
GE	Government effectiveness	Index (-2.5/2.5)	WGI
RL	Rule of law	Index (-2.5/2.5)	WGI
GOV	Average of VA, RL, PV, GE, RQ and CC	Index (-2.5/2.5)	WGI
legalsys	*Legal system and property rights:* Judicial independence, impartial courts, protection of property rights, military interference in rule of law and politics, integrity of the legal system, legal enforcement of contracts, regulatory restrictions on the sale of real property, reliability of police, business costs of crime.	Index (0/10)	EFW
credmark	*Credit market deregulations*: Ownership of banks, private sector credit, interest rate controls/negative real interest rates	Index (0/10)	EFW
labormark	*Labour market deregulations:* Hiring regulations and minimum wage, hiring and firing regulations, centralized collective bargaining, hours regulations, mandated cost of worker dismissal, conscription	Index (0/10)	EFW

Notes: *SWIID: Standardized World Income Inequality Database;*
WDI: World development Indicators, The World Bank;
PWT: Penn World Table 9.0;
WGI; Worldwide Governance Indicators, The World Bank;
EFW: Economic Freedom of the Worlds, The Fraser Institute

Table A2. List of Countries

Lower Middle Income	Upper Middle Income	High Income
Bolivia	Argentina	Austria
Egypt	Bulgaria	Belgium
Honduras	Brazil	Croatia
Morocco	China	Chile
Philippines	Colombia	Cyprus
Ukraine	Costa Rica	Denmark
	Dominican Republic	Finland
	Ecuador	France
	Mexico	Germany
	Paraguay	Greece
	Peru	Ireland
	Romania	Israel
	Russian Federation	Italy
	Turkey	Japan
	Venezuela	Korea, Rep.
		Netherlands
		Norway
		Panama
		Poland
		Portugal
		Singapore
		Spain
		Sweden
		Switzerland
		United Kingdom
		United States
		Uruguay

World Bank Country Classification, 2020, see: https://datahelpdesk.worldbank.org/knowledgebase/articles/906519-world-bank-country-and-lending-groups

Table A3. Institutions, financial development and inequality, robustness check (GMM-sys estimation, 48 countries, 1996-2014)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	GOV	VA	CC	RQ	PV	GE	RL
L.gini	0.958***	0.994***	0.996***	0.980***	0.994***	0.993***	1.033***
	(0.014)	(0.005)	(0.007)	(0.005)	(0.002)	(0.002)	(0.015)
cred	0.006*	0.005***	0.006***	0.008***	0.004***	0.004***	0.016***
	(0.003)	(0.001)	(0.002)	(0.002)	(0.001)	(0.002)	(0.005)
INST	0.375***	0.065	-0.091	0.310**	0.094	0.014	0.767***
	(0.144)	(0.234)	(0.106)	(0.134)	(0.085)	(0.078)	(0.237)
cred*INST	-0.007***	-0.001	-0.002**	-0.007***	-0.001	-0.001	-0.010***
	(0.003)	(0.002)	(0.001)	(0.002)	(0.001)	(0.001)	(0.003)
lninc	0.340***	-0.066	-0.264*	0.015	-0.026	-0.026*	-1.087***
	(0.121)	(0.068)	(0.140)	(0.026)	(0.018)	(0.016)	(0.369)
growth	-0.013***	0.005	-0.009**	-0.004	-0.004	-0.002	0.028
	(0.005)	(0.007)	(0.005)	(0.004)	(0.003)	(0.003)	(0.033)
kaopen	-0.013	0.128	-0.180***	0.003	-0.064*	-0.051	1.621**
	(0.065)	(0.079)	(0.069)	(0.070)	(0.038)	(0.034)	(0.807)
trade	-0.002***	0.000	0.001	-0.000	0.000	0.000*	0.001**
	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
govt	0.010*	0.027**	0.023***	0.005	0.013***	0.014***	0.099***
	(0.005)	(0.011)	(0.005)	(0.005)	(0.003)	(0.004)	(0.032)
hc	-1.461***	-0.135*	0.047	-0.233*	-0.070**	-0.063**	1.061***
	(0.453)	(0.073)	(0.185)	(0.125)	(0.027)	(0.029)	(0.390)
inf	0.003**	0.010*	0.000	0.005***	0.000*	0.000*	0.002***
	(0.001)	(0.006)	(0.000)	(0.002)	(0.000)	(0.000)	(0.001)
indust	0.019***	0.007	0.004	0.003	0.006***	0.005***	0.042***
	(0.005)	(0.007)	(0.003)	(0.003)	(0.002)	(0.001)	(0.013)
tfp	-0.634***	-0.218**	1.881*	-0.138*	-0.176***	-0.168***	1.315**
	(0.221)	(0.099)	(1.101)	(0.077)	(0.050)	(0.056)	(0.568)
labormark	0.386***	0.032**	0.090	0.174**	0.032***	0.032***	-0.009
	(0.135)	(0.014)	(0.067)	(0.077)	(0.010)	(0.009)	(0.023)
crisis	-0.727***	-0.213**	-0.363***	-0.266***	-0.243***	-0.229***	-0.348***
	(0.236)	(0.091)	(0.098)	(0.065)	(0.055)	(0.060)	(0.085)
trend	0.032	0.003	0.013	-0.000	0.002	-0.000	0.014
	(0.022)	(0.011)	(0.015)	(0.010)	(0.007)	(0.007)	(0.010)
Observations	864	864	864	864	864	864	864
Sargan Chi2 stat	4.483	6.972	3.471	6.212	7.332	5.510	6.232
	[0.214]	[0.137]	[0.176]	[0.102]	[0.291]	[0.239]	[0.101]
N. Instruments	19	20	18	19	22	20	19
AB AR(1) test	-4.092	-2.718	-4.575	-3.914	-4.737	-4.624	-1.992
	[0.000]	[0.006]	[0.000]	[0.000]	[0.000]	[0.000]	[0.046]
AB AR(2) test	-0.835	0.0276	-1.454	-0.576	-1.540	-1.553	-0.0637
	[0.404]	[0.987]	[0.146]	[0.565]	[0.124]	[0.120]	[0.542]

Robust standard errors in round parentheses; p-values in squared parentheses; year effects included; *** p<0.01, ** p<0.05, * p<0.1