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Does economic openness matter in the impact of financial development on income inequality?

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Abstract

Despite increasing academic attention on the income distributional impact of financial development, the debate has remained controversial. Hence, this study argues that economic openness to international trade and capital flows may impact the nexus between financial development and income inequality. Using a panel of 71 developing and developed countries for 1994–2017, we first use split-sampling and interaction analyses to examine the role of the country's level of openness on the relationship between financial development and income inequality. However, these two approaches do not provide specific information on the threshold value, if any, at which the effect changes. For this reason, we also employ the dynamic panel threshold method to investigate whether a financial or trade openness threshold exists beyond which financial development worsens income inequality. We find evidence that financial development generally fosters income inequality, but the level of financial and trade openness impacts the inequality effect of financial development. Our results assert that a higher level of financial and trade openness strengthens the pro-inequality impact of financial development.

Keywords

Financial Development

Income Inequality

Trade Openness

Financial Openness

JEL Classification

D31

D63

F02

O11

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1. Introduction

Although the development of financial systems boosts economic growth over the long run (Levine, 1997, 2005), its impact on the allocation of generated wealth remains under question. The World Income Inequality Report, 2022, highlights that income inequality has increased across most countries since the 1980s (Lucas et al., 2022). Hence, financial development (FD) has been accompanied by the debate of whether it comes at the cost of income inequality. According to the world inequality database and world bank's (2019) report there are stunning cross-country differences in the distribution of income and the level of FD. Among financially developed countries, some face high income inequality (e.g., Brazil, India, and Chile), while others are relatively more equal (e.g., Australia, New Zealand, and Sweden). The same is true among less financially developed countries, with some exhibiting extreme income inequality (e.g., Peru and Indonesia) and some facing moderate to relatively low levels (e.g., Belarus and Kazakhstan). This ambiguity calls for more empirical and theoretical research on the relationship between FD and income inequality.

There is an inclusive literature on the FD and income inequality linkage, which can be classified into four groups. First, it is argued and illustrated that the development of the financial sector benefits upper-income individuals more than lower-income ones and thus widens income inequality (Gimet and Lagoarde-Segot, 2011; Jauch and Watzka, 2016; Rajan and Zingales, 2003a). Second, several studies (Beck et al., 2007; Clarke et al., 2006; Galor and Zeira, 1993; Kim et al., 2021) state that broader FD can help low-income individuals get easier access to external finance and earn more by investing and therefore, mitigate income inequality. Third, the inverted U-shaped hypothesis (Greenwood and Jovanovic, 1990) combines the two preceding outcomes, suggesting that income inequality increases at the early stage of FD and then decreases after a certain level of financial sector development. Fourth, a U-shaped nexus is also found (Park and Shin, 2017; Tan and Law, 2012), which implies that financial deepening can reduce income inequality in the early stages of FD while it increases income inequality after reaching a higher level of FD. Thus, the existing literature has not reached any consensus on the income inequality impact of FD.

However, our study goes beyond the FD-income inequality relationship and argues that a country's openness to trade market and capital flow can impact the nexus between FD and income inequality. The literature has emphasized the role of trade and financial openness in promoting FD and determining income inequality. On the one hand, Rajan and Zingales (2003b) argue that a country's openness can weaken the power of established incumbent industrial and financial interest groups who oppose FD and provide incentives for them to develop the financial market. On the other hand, it is argued that the country's openness is one of the determinants of income inequality (Heimberger, 2020; Mills, 2009). Thus, it may be that a certain level of openness must be attained in a country before FD impacts income inequality. There is little direct evidence to confirm that openness makes a difference in how FD affects income inequality. The most relevant paper is by Kunieda et al. (2014), which argues the importance of financial openness in the FD-income inequality relationship. They theoretically and empirically show that FD increases income inequality under higher financial openness but reduces income inequality under lower financial openness. Despite their theoretical and empirical work, the empirical evidence on the impact of openness on the inequality impact of FD remains thin. Therefore, if we assume that an open economy provides fertile ground for the pro-inequality effects of FD, and the impact of FD on income inequality takes effect only after openness exceeds a threshold level, two questions are raised: (i) Would the relationship between FD and income inequality be uniform with the level of openness, including trade openness and financial openness? and (ii) To what extent can FD contribute to the rises in income inequality? The primary objective of this study is to shed light on these questions. This paper represents a first step in providing such empirical evidence by analysing the impact of both trade and financial openness on FD-income inequality relationship.

For this purpose, first, we split the sample into subgroups by the level of financial openness and trade openness to assess whether the inequality impact of FD varies in different subgroups. In further examination, in the full sample dataset, we interact FD with either financial openness or trade openness to examine whether the impact of FD on income inequality depends on the degree of openness. To capture the persistence of income inequality, we allow for the dynamic behaviour of income inequality, which is estimated by the two-step Generalised Method of Moments (GMM), developed by Arellano and Bond (1991). However, these two approaches do not give us specific information on the threshold value, if any, at which the effect becomes different. For this reason, we employ the dynamic panel threshold method of Kremer et al. (2013), which extends the models of Hansen (1999) and Caner and Hansen (2004) to allow for endogenous regressors in a panel setup. We use a dynamic panel threshold mode with a GMM estimator to investigate whether a financial or trade openness threshold exists beyond which FD worsens income inequality. We find evidence that FD generally fosters income inequality, but the level of financial openness and trade openness influences the inequality effect of FD. This evidence suggests that a higher level of financial openness and trade openness strengthens the pro-inequality impact of FD.

This study's contribution is threefold. First, we depart from previous studies by considering the level of the country's openness to both financial and trade markets in the income inequality impact of FD. As a result, this study adds more dimensions – trade openness and financial openness – to the current literature concerning non-linearity in the link between FD and income inequality. Second, while previous empirical study on the moderation impact of financial openness on the FD-income inequality nexus supports a positive monotonic relationship by increasing financial openness, we examine whether a nonlinear relationship exists with a potential threshold effect, using dynamic panel threshold regression. Third, our findings call attention to the need for policymakers to consider the level of a country's openness when exploring possible outcomes from FD and provide insights into how changes in openness will affect those outcomes.

The paper is organised as follows. Section 2 reviews the relevant literature on this topic. Section 3 describes the dataset and provides some preliminary insights about the income inequality effect of FD at different levels of openness. Our empirical methodology, including two estimation methods, dynamic panel GMM and dynamic panel threshold estimation, are discussed in section 4, followed by our empirical findings and discussion in section 5. Finally, Section 6 provides concluding remarks and policy implications.

2. Related literature review

Economic theories and empirical findings have remained inclusive about the impact of FD on income inequality, as pointed out by Demirgüç-Kunt and Levine (2009). One set of theoretical models predicts a positive linear relationship between FD and income inequality, the inequality-widening hypothesis. According to this view, proposed by Rajan and Zingales (2003a), FD mainly benefits upper-level income individuals who can offer collateral and repay their loans, and excludes low-income individuals with collateral constraints, even when financial markets are well-developed. Thus, improving FD proportionally benefits high-income level individuals and widens income inequality. Various studies have provided empirical support for the inequality-widening hypothesis. For example, using country-level data, de Haan and Sturm (2017), Denk and Cournede (2015), Gimet and Lagoarde-Segot, (2011), Jauch and Watzka (2016), and Seven and Coskun (2016) show the positive linkage between FD, either banking development or financial market development, and income inequality. In addition, further empirical support is provided by studies using regional data. For instance, Rodríguez-Pose and Tselios (2009), using a sample of 102 European regions for 1995–2000, find a positive linkage between the per capita added value of the private financial sector and income inequality.

Another set of theoretical models suggests that development in financial markets can mitigate income inequality (Banerjee and Newman, 1993; Galor and Zeira, 1993), inequality-narrowing hypothesis. From this perspective, capital market imperfections (e.g., information and transaction costs) may be especially binding on low-income individuals who lack collateral and credit histories, and any improvement on the imperfection (e.g., abating credit constraints) disproportionally benefits them (Beck et al., 2007). Furthermore, capital constraint reduces the efficiency of capital allocation and prevents the flow of capital to less privileged and low-income individuals, worsening income inequality (Aghion and Bolton, 1997; Galor and Moav, 2004; Galor and Zeira, 1993). Thus, FD can reduce income inequality by alleviating credit constraints and improving capital allocation efficiency. Various cross-country and country-level studies have uncovered evidence favouring the inequality-narrowing hypothesis. For instance, using cross-country data, Beck et al. (2007), Clarke et al. (2006), Hamori and Hashiguchi (2012), Li et al. (1998), Mookerjee and Kalipioni (2010), and Naceur and Zhang (2016) report that FD benefits the poor and reduces income inequality. In addition, some country-level studies have found that FD is negatively associated with income inequality in India (Ang, 2010), Brazil (Meyer Bittencourt, 2006), China (Liang, 2006, 2008), Pakistan (Shahbaz and Islam, 2011), and Vietnam (Hoi and Hoi, 2012).

In addition to the linear relationship, recent theoretical and empirical studies reveal a nonlinear relationship between FD and income inequality, depending on the level of FD.

According to this view, both the inequality-widening and inequality-narrowing hypotheses can be supported. This category can be divided into two. First, an inverted U-shaped relationship between finance and inequality, developed by Greenwood and Jovanovic (1990), suggests that at the early stages of economic development, access to financial services is costly, and only high-income individuals can join financial intermediaries and profit from better financial markets. However, at a higher level of economic development and after a certain level of FD, a more significant proportion of society has access to financial services, which leads to a decrease in income inequality. By using a large sample of countries, some empirical studies (e.g., Kim and Lin, 2011; Nikoloski, 2013) have supported an inverted U shape relationship hypothesis. Second, the U-shaped relationship between FD and income inequality has been reported in some research. For example, Tan and Law (2012) and Park and Shin (2017) show that FD reduces income inequality in the early stages of FD. However, if FD reaches a certain threshold, it will increase income inequality.

Some studies go a step further and argue that institutional quality is the main factor responsible for the nonlinear relationship. Rajan and Zingales (2003b) argue that de jure political representation is dominated by de facto political influence in the presence of weak institutional environments, which allows established interests (incumbent industrial and financial interest groups) to have privileged access to finance so that FD induced by captured direct controls is likely to hurt the poor. In contrast, in the presence of strong institutions, FD may reduce inequality, allowing the poor to invest in building their human and physical capitals (Law et al., 2014). Several studies have attempted to provide direct empirical evidence for the idea that the quality of institutions conditions the link between FD and inequality. Using the aggregate institutional quality measurement, Law et al. (2014) find that FD reduces income inequality only after achieving a certain threshold level of institutional quality. Until then, the effect of FD on income inequality is non-existent. By focusing on single components of governance indicators, some studies find that the positive link between FD and income inequality can be mitigated by low crisis frequency and good governance in the short run (Chen and Kinkyo, 2016), stricter control of corruption (Adams and Klobodu, 2016), and higher democratization (Kim et al., 2021).

In addition to institutional quality, the earlier studies provide a basis for the possible role of the country's openness to trade and financial market in capturing the nonlinear relationship between FD and income inequality. First, the literature has emphasized the role of trade and financial openness in promoting FD. The openness theory of FD, proposed by Rajan and Zingales (2003b), argues that the degree of openness to both international trade and financial flows can boost FD by reducing the power of interest groups and altering their hostile stance toward FD. To further explain this, Rajan and Zingales (2003b) argue that established incumbent industrial and financial interest groups oppose FD because it eases the entry of new firms into the market, increases competition and erodes the monopolistic rents of incumbent groups. However, trade and financial openness can weaken the incumbents' opposition to FD and limit the power of incumbents who oppose FD by introducing foreign competition outside the incumbents' control. It can also create incentives for them to promote FD, which will help them to face competition by providing sufficient finance. A handful of studies have empirically examined the arguments of openness theory (Baltagi et al., 2009; Law, 2009). For example, Baltagi et al. (2009) find that trade and financial openness individually have a significant effect on banking sector development.

Second, many scholars have reached a consensus that there is a relationship between a country's openness and income inequality (Heimberger (2020), and references cited therein). However, despite a wave of research, the sign of the relationship remains ambiguous. Regarding trade openness, the well-known Stolper–Samuelson theorem predicts that the inequality effect of trade openness varies depending on the relative factor abundance. This means that in advanced industrial countries, with an abundance of skilled labour, trade openness increases income inequality by raising the real return to abundant skilled labour and lowering the real rate of return to relatively scarce unskilled labour. The opposite is expected to happen in developing countries, with abundant unskilled labour. International trade will increase demand for unskilled workers, which will push up their real wages and lead to a decrease in income inequality. On an empirical level, the literature has not provided a general conclusion regarding the effect of trade globalisation: while several papers (Goldberg and Pavcnik, 2007; Meschi and Vivarelli, 2009) find a positive impact on income inequality, others (Asteriou et al., 2014; Furceri and Ostry, 2019; Gimet and Lagoarde-Segot, 2011; Jaumotte et al., 2013; Kim et al., 2021) conclude a negative relationship. Regarding financial openness, capital account openness may positively or negatively affect income inequality by fostering international risk-sharing and domestic-consumption smoothing (Kose et al., 2009), increasing the likelihood of financial crises (Furceri and Loungani, 2018; Ghosh et al., 2016) and affecting the bargaining power of labour (Harrison, 2002).[1] A robust positive link between financial openness and income inequality can be found (Asteriou et al., 2014; de Haan and Sturm, 2017; Furceri and Loungani, 2018; Furceri and Ostry, 2019; Jaumotte et al, 2013). In contrast, Kim et al. (2021) and Kunieda et al. (2014) show that financial openness is associated with a reduction in income inequality. All these imply that openness may matter for the nexus between FD and income inequality.

Existing research offers little clear guidance about the role of openness in FD-income inequality. There are two relevant studies in this area of research. Focusing on financial openness, Kunieda et al. (2014) investigate whether financial openness changes the income inequality effect of FD within an economy. Their theoretical model shows that in a financially closed economy, talented agents can borrow financial capital from less talented agents so that less talented agents can utilise the abilities of the talented agents and receive a higher interest rate as credit constraints relax. As a result, FD can narrow income inequality. In contrast, in a financially open economy, talented agents can borrow financial capital in the world market at a low interest rate relative to their abilities. Therefore, the less talented agents cannot utilise the abilities of the talented agents even though credit constraints relax. Thus, inequality increases as the financial market matures. Their empirical results show that financial

[1] See Furceri & Loungani, (2018) for further details. (Mills 2009) (Harrison 2011 for trade openness)

development increases income inequality under higher financial openness, whereas it reduces income inequality under lower financial openness. Focusing on trade openness, Ehrlich and Seidel (2019) theoretically show that the impact of FD on income inequality depends on the size distribution of firms by building a heterogeneous firms model. They argue that FD reduces wage inequality when there are many non-exporting firms and increases wage inequality when there are many large exporting firms. However, their study does not include any empirical investigation. The result of these two studies emphasises the importance of openness in the income inequality impact of FD.

This paper extends the limited literature on FD, openness, and income inequality in three ways. First, this study extends Kunieda et al.'s (2014) work by focusing on the moderation impact of both dimensions of openness, trade integration and financial integration, in the FD-income inequality relationship using a large country-level dataset. Second, this study broadly examines theoretical models developed by Ehrlich and Seidel (2019) at the macro-level by considering trade openness. Third, this study for the first time investigates the threshold effect of openness at which the relationship between FD and income inequality changes.

3. Data and preliminary analysis

The dataset consists of a balanced panel of 71 countries (49 high and upper-middle-income and 22 low and lower-middle-income countries) [2] for which data is available from 1994 to 2017. Our primary dependent variable is income inequality measured using the Gini coefficient from Solt's (2020) Standardized World Income Inequality Database (*SWIID*). The Gini index is derived from the Lorenz curve and ranges between 0 (perfect equality) and 100 (perfect inequality). This index is the most widely used measure of inequality in the literature (Delis et al., 2014; Hasan et al., 2021; Kim et al., 2021). Our preferred income inequality measure is the log of the Gross Gini index (*GrossGini*), which represents household income before taxes since it shows inequality exclusive of the impact of redistribution via taxes and transfers (de Haan and Sturm, 2017; Hasan et al., 2021).

Regarding FD, our main focus is on banking development due to three main reasons explained by Law et al. (2014). First, bank credits are the only possible financing source for most developing countries in our sample. Second, the number of available observations for stock market indicators is insufficient to conduct sample-splitting regression. Third, some empirical studies (e.g., Gimet and Lagoarde-Segot, 2011; Naceur and Zhang, 2016) show that

[2] **High and upper-middle income** countries include Albania, Armenia, Australia, Austria, Barbados, Belgium, Botswana, Brazil, Bulgaria, Chile, China, Colombia, Costa Rica, Czech Republic, Dominican Rep, Ecuador, Finland France, Georgia, Germany, Greece, Hong Kong, , Hungary, Ireland, Italy, Jamaica, Japan, Korea Rep., Malaysia, Mexico, Netherlands, New Zealand, Norway, Panama, Paraguay, Peru, Poland, Portugal , Romania, Singapore, South Africa, Spain, Sweden, Switzerland, Thailand, Turkey, United Kingdom, United States, and Uruguay. **Low and lower-middle income** countries include Bangladesh, Bolivia, Cote d'Ivoire, Egypt, El Salvador, Eswatini, Ghana, Honduras, India, Indonesia, Iran, Kenya, Lesotho, Mongolia, Nigeria, Pakistan, Philippines, Sri Lanka, Tanzania, Tunisia, Uganda, and Ukraine.

the banking sector exerts a more substantial influence on income inequality than the stock market. This may be because the poor have easier access to financial intermediaries (such as banks) than to stock markets, which mostly have stringent participation requirements (Isah, 2016). Following common practice, we used two proxies to measure banking development. Our first and preferred set is the log of private credit by deposit money banks and other financial institutions as a share of GDP (*PrivateCredit*). As robustness checks, we also use the log of liquid liabilities, also known as broad money, over GDP (*LiquidLiabilities*) to measure the respective size of the banking sector. All these data are sourced from the World Bank's Global Financial Development Database (*GFDD*).

Two main variables to measure economic openness (*Openness*) are trade openness and financial openness. Trade openness is measured as the total volume of imports and exports over the annual GDP of a country (*TO*) from World Bank's World Development Indicators (*WDI*) database. Kim et al. (2010) argue that this measure can provide an unambiguous quantification of trade openness. Financial openness is measured based on a de facto index developed by Lane and Milesi-Ferretti (2007). This variable is defined as the volume of a country's foreign assets and liabilities as a percentage of GDP (*FO*). These foreign assets and liabilities include portfolio debt, foreign direct investment, and foreign indirect investment (portfolio investment). There are also other measurements of financial openness, such as the "de jure" index of Chinn and Ito (2006). However, they are not used in this study because they are noisy indicators of capital account openness (Bui and Bui, 2020).

Finally, we consider the various control variables used to explain income inequality by following the extensive literature on the determinants of income inequality (Beck et al., 2007; Delis et al., 2014). We use the changes in the log of GDP per capita (*Growth*) to account for the impact of economic growth on income distribution, the CPI-based inflation rate (*Inf*) to control for monetary condition, and the log of the population size (*Pop*) to control for the demographics in each country. To account for the activity and growth of government over the sample period, we include the ratio of central government expenditures as a share of GDP (*GovExp*). Higher government spending may disproportionately help the poor if used efficiently, but it may be wasteful when institutions are weak (Delis et al., 2014). In addition, some measure of human capital is also included as a control variable in most inequality equations. Typically, education proxies for human capital. However, because there is an ongoing debate on quantity vs quality in education, the standard education data available via Barro and Lee (2001) may not be a good control (Hasan et al., 2021). Since Castelló-Climent and Doménech (2008) document that life expectancy is strongly linked to human capital accumulation, we use life expectancy (*LifExp*) as a proxy for human capital in our estimations. Because we use estimators based on fixed effects, we do not control for time-invariant variables. All control variables are obtained from the *WDI* database. Detailed variable descriptions are provided in Appendix A.

We use data from 1994 to 2017 for 71 countries because of data availability for important variables such as FD and economic openness. Then, as is standard in the literature (de Haan and Sturm, 2017; Delis et al., 2014; Kim et al., 2021), we create 3-year averaged data for eight non-overlapping periods: 1994–1996, 1997–1999, 2000–2002, 2003–2005, 2006-2008, 2009-2011, 2012-2014, and 2015-2017, to mitigate any noise associated with short-run economic fluctuations. To maintain a balanced dataset, variables with missing values are imputed. Table 1 provides descriptive statistics for all variables in the sample of 568 observations. Our dependent variable is the GrossGini which has an average of 46.68 and ranged from a minimum of 24.23 (Ukraine for the period of 2000-2002) to 69.03 (South Africa for the period of 2000-2002). Turning to our main independent variable, private credit by deposit money banks and other financial institutions (*PrivateCredit*) and broad money (*LiquidLiabilities*) are, on average, 41.26 (3.7201) and 50.51 (3.9222) per cent of GDP, respectively. In addition, Table 2 reports the correlation coefficients among all dependent, independent, and control variables. Table 2 shows that none of the variables are highly correlated, with the largest correlation coefficient being 0.7065 between *LifeExp* and *PrivateCredit*.

Table 1. Summary statistic

Variable	**N**	**mean**	**S.d.**	**min**	**p25**	**p50**	**p75**	**max**
GrossGini	568	3.8434	0.1451	3.1877	3.7861	3.8498	3.9170	4.2346
PrivateCredit	568	3.7201	0.8745	1.3424	3.1444	3.8384	4.4302	5.1476
LiquidLiabilities	568	3.9222	0.6637	2.2696	3.5047	3.9207	4.3786	5.4277
FO	568	0.5057	0.8871	-0.9075	-0.1092	0.3260	0.9355	3.2035
TO	568	4.2374	0.5445	3.0739	3.8958	4.2018	4.5365	5.9108
Growth	568	0.0726	0.0708	-0.1121	0.0315	0.0692	0.1098	0.2985
Inf	568	7.6298	12.9632	-0.4801	2.0283	4.1115	8.2910	93.2399
GovExp	568	2.6595	0.3544	1.5988	2.4222	2.6805	2.9355	3.5168
Pop	568	16.7423	1.5819	12.5499	15.5669	16.6232	17.8993	20.9826
LifeExp	568	71.9344	8.9222	44.7493	69.0885	74.1557	78.2787	83.1480

Table 2. Correlation Matrix

	GrossGini	privateCredit	LiquidLiabilities	FO	TO	Growth	Inf	GovExp	Pop
GrossGini	1								
PrivateCredit	0.0640	1							
LiquidLiabilities	-0.0097	0.8622*	1						
FO	0.1660*	0.6174*	0.5855*	1					
TO	-0.0133	0.2318*	0.2449*	0.5987*	1				
Growth	-0.0898*	-0.0890*	-0.0509	-0.2016*	0.0464	1			
Inf	-0.1454*	-0.4223	-0.3666*	-0.2928*	-0.1073*	-0.1710*	1		
GovExp	0.2505*	0.3472*	0.2948*	0.3782*	0.1681*	-0.2185*	-0.1548*	1	
Pop	-0.1765*	0.0600	0.1104*	-0.3142*	-0.5784	0.0706	0.02910	-0.2216*	1
LifeExp	-0.0730	0.7065*	0.6437*	0.5167*	0.1567*	-0.0491	-0.2388*	0.2345*	-0.0400

To provide further insight for our analysis, Fig. 1 shows the relationship between income inequality and FD controlling for country and year fixed effect. This figure suggests that more FD, measured by *PrivateCredit* or *LiquidLiabilities*, slightly increases income inequality, measured by *GiniGross*. To explore whether the impact of FD on income inequality changes by economic openness, which includes both financial and trade openness, we divide countries into four groups. The first two groups are countries whose average level of financial openness over the sample period is less (more) than the first quantile (equal to 0.97), denoted by LFO (HFO). Fig. 2 implies that the relationship between FD and income inequality varies according to the level of financial openness. There is a positive relationship between FD and income inequality among HFO countries (e.g., Netherlands, France, United Kingdom, etc.), while no relationship is observed among LFO countries (e.g., Brazil, Colombia, Bangladesh, etc.). The second two groups are countries whose average level of trade openness over the sample period is less (more) than the first quantile (equal to 50.25), denoted by LTO (HTO). Fig. 3 implies that the relationship between FD and income inequality varies according to the level of trade openness. There is a positive relationship between FD and income inequality among HTO countries (e.g., Thailand, Malaysia, Singapore, etc.), while this positive relationship is not observed among LTO countries (e.g., Japan, United States, Australia, etc.). This suggests that openness strengthens the pro-inequality impact of FD. Appendix B provides a list of LFO, HFO, LTO and HTO countries in our sample.

Figure 1. Relationship between FD and income inequality

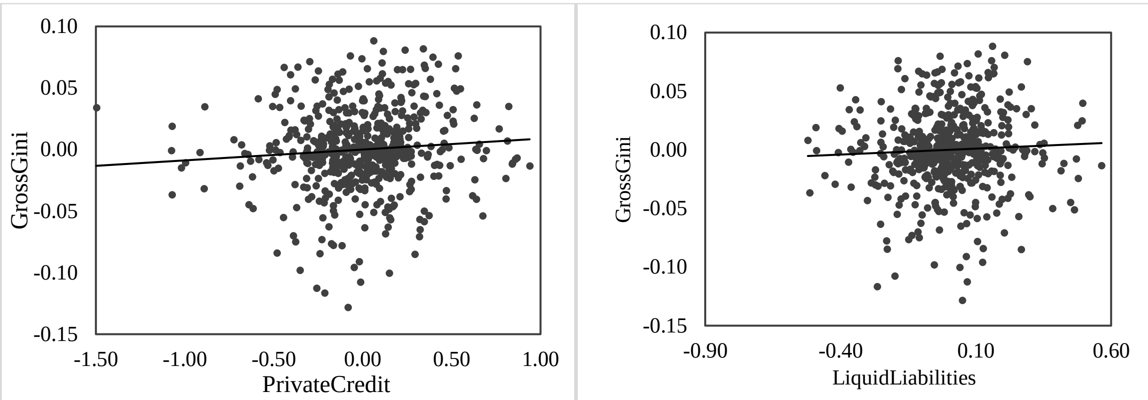


Figure 2. Relationship between FD and income inequality for LFO and HFO countries

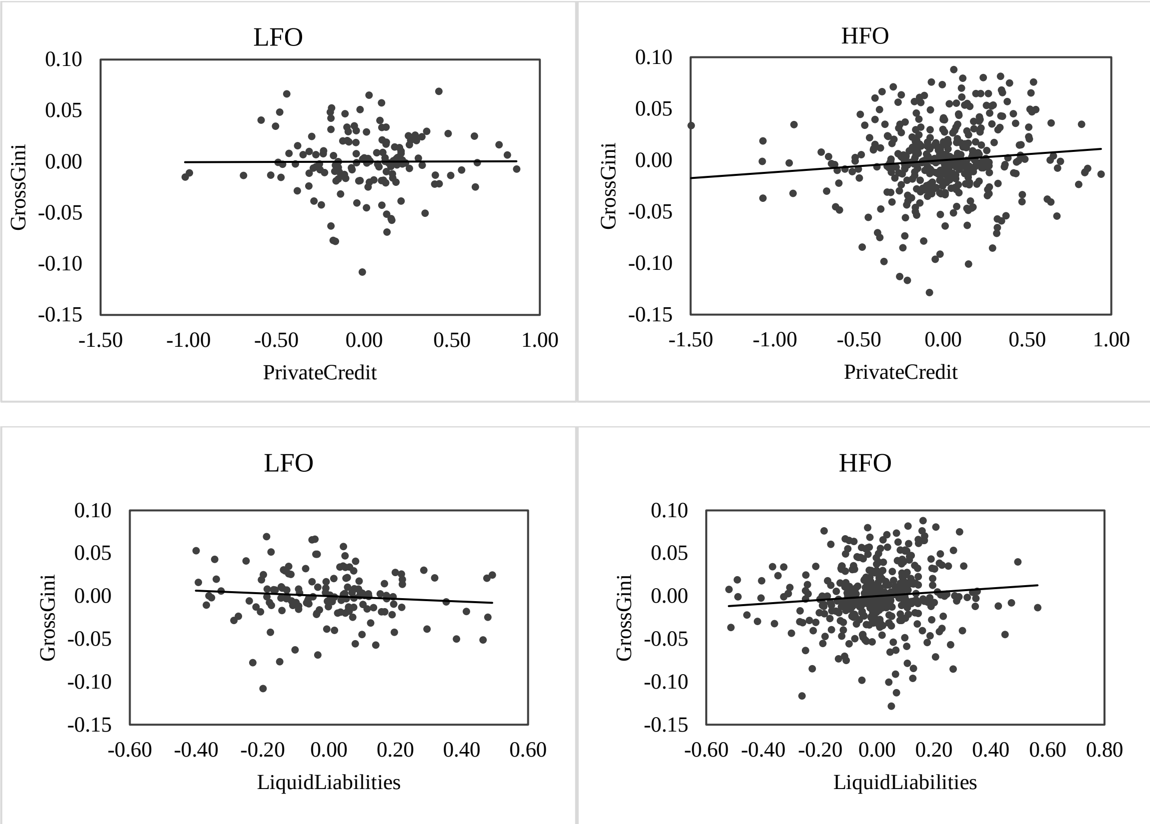
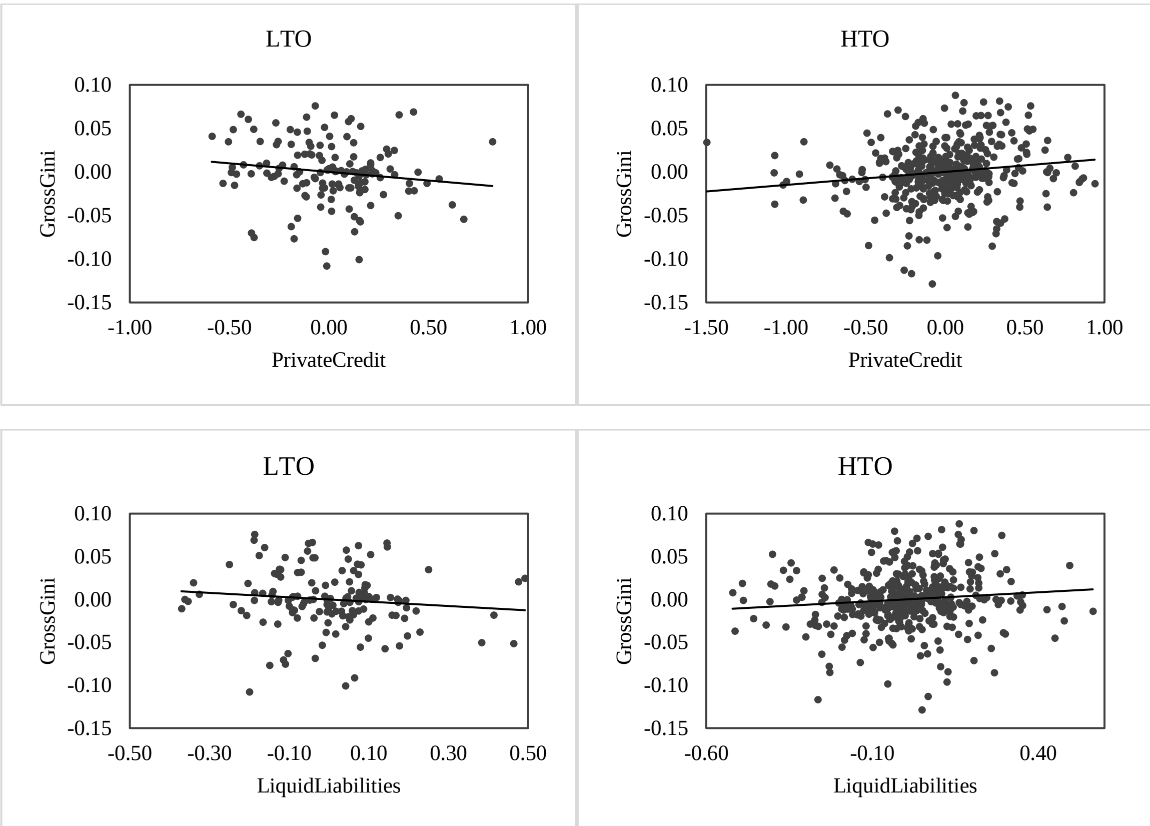


Figure 3. Relationship between FD and income inequality for LTO and HTO countries



4. Methodology

4.1. Dynamic Panel GMM estimation

Our empirical analysis of the FD-income inequality nexus begins by using split sample and interaction analyses to examine the extent to which income inequality and financial development relationship depend on the country's openness level. For split-sampling analysis, we estimate the following dynamic equation for HTO, LTO, HFO and LFO subsamples:

$$II_{i,t} = \beta_1 II_{i,t-1} + \beta_2 FD_{i,t} + \beta_3 FO_{i,t} + \beta_4 TO_{i,t} + \beta_5 Growth_{i,t} + \beta_6 Inf_{i,t} + \beta_7 GovExp_{i,t} + \beta_8 Pop_{i,t} + \beta_9 \, LifeExp_{i,t} + \tau_t + \mu_i + \varepsilon_{i,t} \qquad (1)$$

The subscript *i* is the country indicator, and *t* is the period index. τ indicates time-fixed effects, μ represents the time-invariant country-fixed effect, and ε is an error term assumed to be independent and identically distributed with mean zero and constant variance. II (income inequality) is the dependent variable, which is the indicator of income inequality. The model includes a lagged dependent variable to reflect the persistence of income inequality over time (Delis et al., 2014; Jauch and Watzka, 2016; Kim et al., 2021). FD, FO and TO are financial development, financial openness and trade openness, respectively. To strengthen our empirical results, we also include some other variables, such as the log of GDP per capita growth (*Growth*), inflation rate (*Inf*), the log of government expenditures over GDP (*GovExp*), the log of the population (*Pop*), and the log of life expectancy (*LifeExp*).

To address whether the effect of FD on income inequality differs along with the extent of financial and trade openness, the interaction terms between FD and FO and between FD and TO are included in Equations 2 and 3, respectively.

$$II_{i,t} = \beta_1 II_{i,t-1} + \beta_2 FD_{i,t} + \alpha_1 FD_{i,t} \times FO_{i,t} + \beta_3 FO_{i,t} + \beta_4 TO_{i,t} + \beta_5 Growth_{i,t} + \beta_6 Inf_{i,t} + \beta_7 GovExp_{i,t} + \beta_8 Pop_{i,t} + \beta_9 \, LifeExp_{i,t} + \tau_t + \mu_i + \varepsilon_{i,t} \qquad (2)$$

$$II_{i,t} = \beta_1 II_{i,t-1} + \beta_2 FD_{i,t} + \alpha_2 FD_{i,t} \times TO_{i,t} + \beta_3 FO_{i,t} + \beta_4 TO_{i,t} + \beta_5 Growth_{i,t} + \beta_6 Inf_{i,t} + \beta_7 GovExp_{i,t} + \beta_8 Pop_{i,t} + \beta_9 \, LifeExp_{i,t} + \tau_t + \mu_i + \varepsilon_{i,t} \qquad (3)$$

At the margin, the total effect of increasing FD can be calculated by examining the partial derivatives of income inequality with respect to FD. We expect negative β_2 and positive α_1 and α_2, which implies that FD increases (decreases) inequality in an open (closed) economy to the world financial or trade market.

Some potential problems arise when estimating the above equations. The primary identification issue is the potential endogeneity of FD. As discussed in the literature (Beck et al., 2007; Jauch and Watzka, 2016; Law et al., 2014), FD is highly likely to be endogenous, possibly due to feedback from income inequality to FD (reverse causality). For example, reductions in income inequality may stimulate demand for financial services (Beck et al., 2007). Previous research used instruments for financial development. These instruments were similar to those in the literature on the FD–growth nexus, usually the origin of a country's legal system, which may not be good instruments for FD when investigating the inequality nexus (Jauch and

Watzka, 2016). In addition, the inclusion of the lagged dependent variable in the empirical model implies a correlation between the regressors and the error term, which could bias the coefficient estimates (Baltagi et al., 2009). Besides these endogeneity considerations, even when using standard two-stage least squares regressions and instruments for financial development, this does not control for the endogeneity of other explanatory variables, which may bias the coefficient estimates on financial development (Beck et al., 2007).

Thus, the preferred estimator, in this case, is the two-step Generalised Method of Moments (*GMM*) suggested by Arellano and Bond (1991) with robust standard errors, which eliminates any endogeneity that may be due to the correlation of country-specific, time-invariant factors and right-hand side regressors[3] (more details in Baltagi et al., 2009; Jauch and Watzka, 2016). There is convincing evidence that too many moment conditions introduce bias while increasing efficiency. It is, therefore, suggested that a subset of these moment conditions be used to take advantage of the trade-off between the reduction in bias and the loss in efficiency (see Baltagi, 2005, and the references cited there). We treat *GrossGini* and all right-hand side variables, except *Inf, GovExp, Pop and LifeExp,* as potentially endogenous variables, and we use their lagged values as instruments.[4] The specification is checked using the Hansen statistic, a test of overidentifying restrictions for the validity of the instrument set. In addition, two diagnostics are computed using the Arellano and Bond GMM procedure to test for first-order and second-order serial correlation in the disturbances. One should reject the null of the absence of first-order serial correlation and not reject the absence of second-order serial correlation.

4.2. Dynamic panel threshold estimation

We continue our empirical analysis by testing the existence of a threshold level of economic openness (*Openness*), either FO or TO, in the relationship between income inequality and FD. Thus, the dynamic panel threshold model of economic openness takes the following form:

$$\begin{aligned} II_{i,t} = {} & \beta_1 II_{i,t-1} + \theta_1 FD_{i,t}\ I\big(Openness_{i,t} \leq \gamma\big) + \theta_2 FD_{i,t}\ I\big(Openness_{i,t} > \gamma\big) + \beta_3 FO_{i,t} \\ & + \beta_4 TO_{i,t} + \beta_5 Growth_{i,t} + \beta_6 Inf_{i,t} + \beta_7 GovExp_{i,t} + \beta_8 Pop_{i,t} \\ & + \beta_9\ LifeExp_{i,t} + \tau_t + \mu_i + \varepsilon_{i,t} \quad (4) \end{aligned}$$

where subscript *i* represents the country and *t* indicates the period. μ and τ are the time-fixed effect and country-fixed effect respectively, the error term $\varepsilon_{i,t} \overset{iid}{\sim} (0, \sigma^2)$. γ is the threshold level, and $I(\cdot)$ is an indicator function taking a value of 1 if the argument in the indicator function holds and 0 otherwise. Openness is the threshold variable, which is measured by either *FO* or *TO*. FD is a regime-dependent variable, measured by *PrivateCredit* or *LiquidLiabilities*, with the slope parameter θ_1 if Openness is less than or equal to γ and θ_2 otherwise. In this model, the explanatory variables are partitioned into a subset of exogenous

[3] An additional advantage of the GMM estimator is that by differencing it helps to ensure that all the regressors are stationary.

[4] To reduce the instrument proliferation problem that can result in biased parameter estimates, we used the "collapse" option in the xtabond2 STATA command to collapse the instrument matrix. We kept the number of instruments below the number of countries. See Roodman (2009) for details.

variables (*TO*, *FO*, *Inf*, *GovExp*, *Pop*, and *LifeExp*) uncorrelated with $\varepsilon_{i,t}$, and a subset of endogenous variables ($II_{i,t-1}$, *FD* and *Growth*) correlated with $\varepsilon_{i,t}$.

To estimate Eq. 6, we employ Kremer et al.'s (2013) estimation method, which allows estimating of threshold effects with panel data in the case of endogenous regressors by extending Hansen's (1999) and Caner and Hansen's (2004) model. Following their method, we consider the forward orthogonal deviation transformation suggested by Arellano and Bover (1995) to eliminate the country-specific fixed effects and avoid the serial correlation of the transformed error terms. According to Kremer et al. (2013), first, we estimate a reduced-form regression of the endogenous variables on their instruments and then replace the endogenous variables in the structural equation with the predicted values. Second, we estimate Eq. 6 using least squares for a fixed threshold γ where the endogenous variables are replaced by their predicted values and define $S(\gamma)$ as sum of squared residuals. We repeat this step for a strict subset of the threshold variable *Openness*. Finally, the estimator of the threshold value γ is selected as the one associated with the smallest sum of squared residuals, i.e. , $\hat{\gamma} = \underset{\gamma}{\arg\min} S_n(\gamma)$. In addition, the 95% confidence interval of the threshold value is calculated by $\Gamma = \{\gamma : LR(\gamma) \leq C(\alpha)\}$, where $C(\alpha)$ is the 95th percentile of the asymptotic distribution of the likelihood ratio statistic $LR(\gamma)$ (Caner and Hansen, 2004; Hansen, 1999). Once the threshold value (γ) is obtained, the slope coefficients are estimated by the GMM for the previously used instruments and estimated threshold. Following Arellano and Bover (1995), we use lags of endogenous variables as instruments.

5. Empirical Results

This section starts by reporting the results of estimating Eq. 1, 2 and 3 on the dataset described before using dynamic GMM estimation and outlines their implications for the hypothesis of interest. It is followed by the result of the dynamic threshold regression that tests the existence of the threshold value of economic openness.

5.1. Dynamic Panel GMM estimation results

Table 3 reports the result of estimating Eq. 1 for the entire sample of 71 countries using dynamic panel GMM as proposed by Arellano and Bond (1991), which presents the impact of FD, measured by either *PrivateCredit* or *LiquidLiabilities* on income inequality. All columns include country and time-fixed effects, and robust standard errors are reported in parentheses. In Columns (1) and (2), in which the primary determinants of income inequality are controlled, the estimated coefficient for both *PrivateCredit* and *LiquidLiabilities* is positive and statistically significant at 5%. The effect of banking development on inequality remains positively significant even if *Pop* and *LifeExp* are controlled for in columns (3) and (4). Thus, the positive coefficient of banking development in all columns suggests that banking development will lead to an increase in income inequality; for example, according to the result in the first column, a 10% increase in *PrivateCredit* increases the gross Gini coefficient by 1.4%. This finding is in line with Gimet and Lagoarde-Segot (2011), Jauch and Watzka (2016), De Hann and Strum (2017), Blau (2018), Hsieh et al. (2019), and Kim (2021). Among the rest of the explanatory variables, column 4 shows that *Growth*, *FO* and *GovExp* negatively impact income inequality. Furthermore, the identification tests show no overidentifying restrictions and no serial correlation between the instruments and the disturbance. A reason for this positive link between FD and income inequality might be, as argued by Rajan and Zingales (2003), that the rich can offer collateral and are more likely to repay their loans. The poor, who do not enjoy this benefit, might find it difficult to obtain loans even in a well-developed banking sector.

Table 3. The impact of financial development on income inequality

	GrossGini			
	(1)	(2)	(3)	(4)
Lag.GrossGini	0.6622***	0.6750***	0.6417***	0.6878***
	(0.1103)	(0.0976)	(0.1009)	(0.0703)
PrivateCredit	0.0149**		0.0095**	
	(0.0059)		(0.0043)	
LiquidLiabilities		0.0313**		0.0618***
		(0.0141)		(-0.023)
FO	0.0058	-0.0012	0.0111	-0.0444*
	(0.0161)	(0.0134)	(0.0196)	(0.0239)
TO	0.0192	0.0239	0.0302	0.0192
	(0.0259)	(0.0244)	(0.0270)	(0.0264)
Growth	-0.0628	-0.0968	-0.0728	-0.1673**
	(0.0619)	(-0.062)	(0.0598)	(0.0667)
Inf	-0.0001	-0.0002	-0.0002	-0.0001
	(0.0002)	(0.0002)	(0.0003)	(0.0003)
GovExp	-0.005	-0.0135	-0.007	-0.0302**
	(0.0116)	(0.0111)	(0.0118)	(0.0141)
Pop			0.0319	-0.0553
			(0.0597)	(0.0618)
LifeExp			-0.0000	-0.0010
			(0.0010)	(0.0012)
Country fixed effect	Yes	Yes	Yes	Yes
Time fixed effect	Yes	Yes	Yes	Yes
Robust Standard Error	Yes	Yes	Yes	Yes
Observation	426	426	426	426
Instrument	35	35	33	32
Country	71	71	71	71
Hansen test of over-identification	0.1467	0.1622	0.1186	0.5457
AR(1)	0.5816	0.6883	0.6053	0.3453
AR(2)	0.1241	0.2375	0.1287	0.5201

Notes: Table 3 reports the impact of financial development on income inequality.
Each regression includes country and time-fixed effects. Robust standard errors clustered at the country level are reported in parentheses.
*, **, *** indicate significance at the 10%, 5% and 1% levels, respectively.
All variables are defined in Appendix A

Financial Openness moderation effect:

In order to examine whether financial openness affects the relationship between financial development and income inequality, as suggested by Kunieda et al. (2014), Table 4 presents the result of split-sampling and interaction analysis. First, we start with split-sampling analysis, and we classify the sample into two groups: countries with strongly low levels of financial openness (FO < 1st Quantile) and Countries with high levels of financial openness (FO>1st Quantile). Then we estimate sub-sample regressions to compare the FD-II relationship among LFO and HFO Countries. Columns (1) and (4) present the GMM estimation result for LFO countries. As shown, the coefficient of *PrivateCredit* and *LiqiuidLiablilities* is not statistically significant, suggesting that banking development has no significant impact on income inequality in financially closed countries. Next, we examine the FD-income inequality relationship for financially open countries (HFO). As shown in columns (2) and (5) of Table 4, the GMM estimation results for these 53 open countries indicate that the coefficients of *PrivateCredit* and *LiqiuidLiablilities* are statistically significant and positive. This result suggests that FD widens income inequality in financially open countries.

As an alternative to split-sampling, columns 3 and 6 report the estimation result of Eq. 2, in which the interaction of FD and financial openness is included. As shown in both columns, FD and its interaction with financial openness significantly and positively affect income inequality. This finding suggests that financial openness strengthens the positive association between banking development and income inequality. To better analyse the interaction result, Fig. 4 shows the marginal impact of FD on income inequality for different levels of financial openness, based on the estimates reported in columns 3 and 6. The whiskers in Fig. 4 show the 95% confidence band. Fig. 4 shows no significant relationship between FD and income inequality at the lower level of financial openness. However, by increasing financial openness, the impact of FD on the Gini coefficient is higher and more significant. This finding holds for both measures of FD. Considering *PrivateCredit* as a measure of FD, when financial openness increases from the lowest to the highest, FD's marginal impact on income inequality increases from -0.4% to 11.56%. To put these coefficient estimates into perspective, for each country, the average value of FO over 1994-2017 is calculated, and then the marginal impact of FD on income inequality for the top three financially open countries and the bottom three financially closed countries are measured. The result shows that in countries with the highest FO, including Ireland, Hong Kong, and Singapore, 10% increases in *PrivateCredit* lead to 0.99, 0.98 and 0.94% rises in income inequality, respectively. However, in countries with the lowest FO, such as Kenya, Iran, and Bangladesh, the FD-II relationship is insignificant.

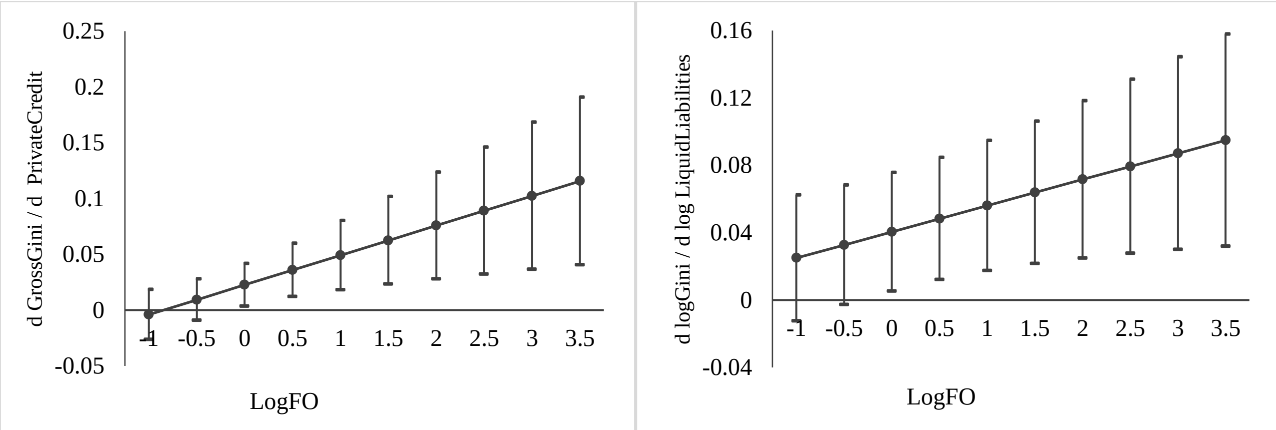
This finding is partially in line with Kunieda et al. (2014), who argue that financial openness changes the FD-income inequality relationship. In countries with higher financial openness, our result is consistent with their study asserting that the development of the financial system leads to a rise in income inequality in financially open countries. According to their theoretical model, when the domestic financial market develops and credit constraints relax in financially open countries, investors borrow financial capital in the world market with a constant interest rate that is low relative to their abilities. On the other hand, the lenders in the country do not benefit from the development of the domestic financial market since the interest rate does not increase. Therefore, inequality increases as the domestic financial market develops. In countries with lower financial openness, our results do not support their prediction asserting a negative relationship between income inequality and FD in financially closed countries. In fact, our findings show that there is no significant inequality impact of FD in countries with lower financial openness. Therefore, our results show that financial openness strengthens the pro-inequality impact of FD.

Table 4. Income inequality, FD and financial openness

	logGrossGini					
	(1) FO<Q1	(2) FO>Q1	(3) Interaction	(4) FO<Q1	(5) FO>Q1	(6) Interaction
Lag.GrossGini	0.7720***	0.6422***	0.6642***	0.6985***	0.6492***	0.7040***
	(0.1018)	(0.0660)	(0.0614)	(0.0822)	(0.0638)	(0.0762)
PrivateCredit	0.066	0.0129**	0.0225**			
	(0.0439)	(0.0062)	(0.0097)			
PrivateCredit * FO			0.0266***			
			(0.0098)			
LiquidLiabilities				-0.0197	0.0699***	0.0405**
				(0.0529)	(0.0251)	(0.0179)
LiquidLiabilities *FO						0.0155**
						(0.0074)
FO	-0.0553	0.0196	-0.0613	-0.0468	-0.0546	-0.0678*
	(0.0357)	(0.0329)	(0.0436)	(0.0293)	(0.0459)	(0.0384)
TO	0.0198	0.0551*	0.0412	-0.0116	0.0113	0.0141
	(0.0491)	(0.0328)	(0.0281)	(0.0577)	(0.0246)	(0.0256)
Growth	0.1028	-0.1287	0.0242	0.1013	-0.1700**	-0.1253**
	(0.2627)	(0.1043)	(0.0989)	(0.2283)	(0.0776)	(-0.049)
Inf	0.0005	-0.0005	-0.0002	0.0004	0.0002	-0.0003
	(0.0005)	(0.0005)	(0.0003)	(0.0005)	(0.0003)	(0.0003)
GovExp	-0.0111	-0.0298	-0.005	-0.0066	-0.0318**	-0.0279**
	(0.0256)	(0.0261)	(0.0169)	(0.0133)	(0.0154)	(0.0132)
Pop	0.0811	0.1276	0.1296	-0.1063	-0.1005	0.0486
	(0.0850)	(0.1205)	(0.0872)	(0.1361)	(0.0746)	(0.0503)
LifeExp	-0.0063	-0.0005	-0.0016	0.0018	-0.0019	-0.0025
	(0.0052)	(0.0019)	(0.0013)	(0.0037)	(-0.002)	(0.0016)
Observation	108	318	426	108	318	426
Instrument	16	28	21	17	32	36
Country	18	53	71	18	53	71
Hansen test of over-identification	0.4774	0.4161	0.7003	0.1186	0.4342	0.1746
AR(1)	0.4595	0.9633	0.386	0.956	0.0643	0.9877
AR(2)	0.8228	0.0588	0.0715	0.2769	0.4476	0.0833

Notes: Table 4 reports the impact of financial openness on the FD and income inequality relationship. Each regression includes country and time-fixed effects. Robust standard errors clustered at the country level are reported in parentheses. *, **, *** indicate significance at the 10%, 5% and 1% levels, respectively.
All variables are defined in Appendix A

Figure 4. Marginal impact of FD on income inequality for different levels of FO



Trade Openness moderation effect:

This section reports how openness to the global trade market can change the relationship between FD and income inequality (Table 5). In the split-sampling analysis, the sample is broken down into two groups: countries with strongly low-level trade openness (TO < 1st Quantile) and countries with high-level trade openness (TO>1st Quantile). Then Eq. 1 is estimated for each sub-sample to compare the FD–income inequality relationship in LTO and HTO Countries. For LTO countries, columns (1) and (4) of Table 5 show that banking development, measured by *PrivateCredit* or *LiquidLiabilities,* does not have a statistically significant impact on the Gini Coefficient. However, turning to HTO countries, columns (2) and (5) indicate that FD worsens income inequality in countries open to the trade market. This finding suggests that, similar to financial openness, trade openness also strengthens the positive relationship between FD and income inequality.

With regard to the interaction analysis, we estimate Eq. 3 in which the interaction of FD and trade openness is included to examine whether trade openness influences the impact of FD on income inequality. Columns (3) and (6) show that the estimated coefficients of banking development and its interaction with trade openness are statistically negative and positive, respectively. This means that banking development, measured by either *PrivateCredit* or *LiquidLiabilities*, tends to significantly increase income inequality in countries with higher trade openness. Corresponding to columns (3) and (6), Fig. 5 shows the marginal impact of FD on income inequality for different levels of financial openness with a 95% confidence band. Fig. 5 suggests that trade openness changes the relationship between FD and income inequality. In fact, the estimated marginal effect of FD on II is non-significantly negative but turns significantly positive when trade openness is increasing. Considering *PrivateCredit* as a measure of FD, by increasing trade openness from the lowest to the highest level, FD's marginal impact on income inequality increases from -3.9% to 8.3%. Focusing on high trade openness countries, based on the average value of trade openness over 1994-2017, 10% increases in

PrivateCredit led to a significant rise in the Gini index of 0.77% in Singapore, 0.74% in Hong Kong and 0.5% in Malaysia.

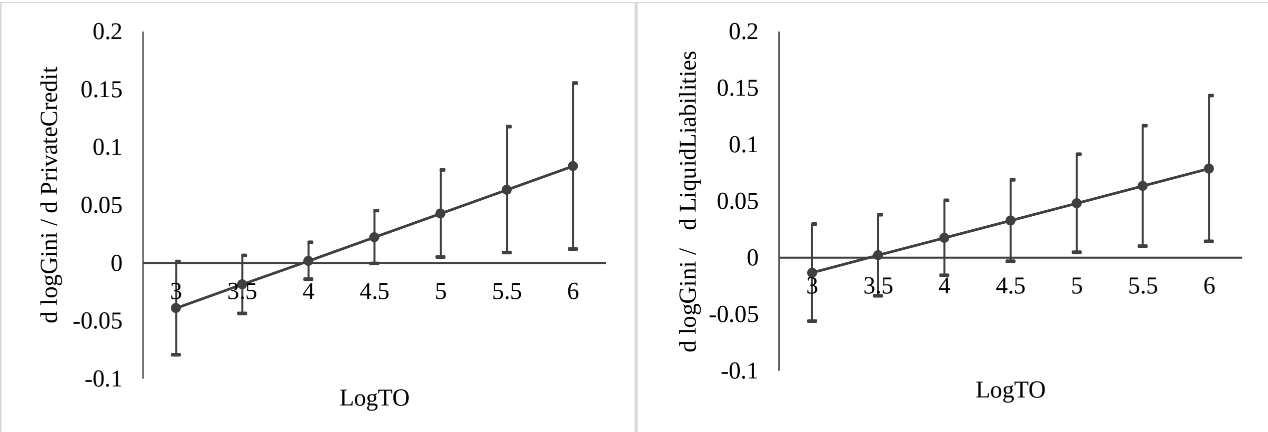
This result provides broad empirical support at the macro level for the prediction of Ehrlich and Seidel (2019), who argue that the FD–income inequality relationship depends on the number of exporting firms. In fact, they assert that FD increases income inequality in countries with a high percentage of exporting firms, while it decreases income inequality in countries with a high percentage of non-exporting firms. Consistent with their argument, at the macro-level, our results show that the development of financial markets significantly increases income inequality in countries with higher trade openness. However, we do not find a negative relationship between FD and income inequality in countries with lower trade openness.

Table 5. Income inequality, FD, and trade openness

	logGrossGini					
	(1) TO<Q1	(2) TO>Q1	(3) Interaction	(4) TO<Q1	(5) TO>Q1	(6) Interaction
Lag.GrossGini	0.8980***	0.6320***	0.6527***	0.9784***	0.6480***	0.6456***
	(0.1235)	(0.0707)	(-0.067)	(0.2391)	(0.0679)	(0.0658)
PrivateCredit	0.0178	0.0166**	-0.1618**			
	(0.0159)	(0.0079)	(-0.074)			
PrivateCredit * TO			0.0409**			
			(0.0182)			
LiquidLiabilities				0.1031	0.0259**	-0.1053*
				(0.0988)	(0.0117)	(0.0584)
LiquidLiabilities *TO						0.0307**
						(-0.014)
TO	0.0277	0.0401	-0.0922*	0.0558	0.0161	-0.0745
	(0.0351)	(0.0336)	(0.0468)	(0.0596)	(0.0253)	(0.0457)
FO	-0.0204	0.0076	-0.0087	-0.077	0.007	0.0046
	(-0.041)	(0.0158)	(0.0316)	(-0.066)	(0.0165)	(0.0188)
Growth	-0.0458	-0.0036	-0.092	0.0629	-0.0302	-0.0433
	(0.1381)	(0.0508)	(0.1122)	(0.2263)	(0.0464)	(0.1005)
Inf	-0.0002	-0.0002	-0.0003	-0.0006	-0.0003	-0.0002
	(0.0002)	(0.0003)	(0.0003)	(0.0004)	(0.0003)	(0.0003)
GovExp	-0.0054	0.0064	-0.0012	-0.0695	0.0045	-0.0041
	(0.0262)	(0.0213)	(0.0215)	(0.0554)	(0.0174)	(0.0141)
Pop	0.0969	0.0546	0.0813	0.0313	0.0169	0.051
	(0.0646)	(0.0699)	(0.0727)	(0.1223)	(0.0539)	(0.0439)
LifeExp	-0.0057*	-0.001	-0.0015	-0.0081	-0.0005	-0.0011
	(-0.003)	(0.0013)	(0.0011)	(0.0053)	(-0.001)	(0.0009)
Observation	108	318	426	108	318	426
Instrument	16	26	24	16	29	24
Country	18	53	71	18	53	71
Hansen test of over-identification	0.657	0.7072	0.4025	0.9084	0.673	0.3166
AR(1)	0.3823	0.1117	0.4999	0.7276	0.1958	0.512
AR(2)	0.1179	0.2846	0.0512	0.3752	0.2923	0.0905

Notes: Table 5 reports the impact of financial openness on the FD and income inequality relationship.
Each regression includes country and time-fixed effects. Robust standard errors clustered at the country level are reported in parentheses.
*, **, *** indicate significance at the 10%, 5% and 1% levels, respectively.
All variables are defined in Appendix A

Figure 5. Marginal impact of FD on income inequality for different levels of TO



5.2. Dynamic Threshold Regression

Although the split-sampling and interaction analysis provide informative results, each approach has its limitations. In the split-sampling regression, the result may be sensitive to the arbitrary cut-off value. With regard to the interaction analysis, it is assumed that the impact of banking development on the Gini coefficient grows monotonically with the increase in financial or trade openness. To overcome these limitations, we perform a dynamic panel threshold model to test the existence of a threshold level of financial or trade openness in the relationship between FD and income inequality. Table 6 presents the estimation result of the dynamic panel threshold model in Eq. 6, where the threshold variable is either financial openness (columns 1 and 3) or trade openness (columns 2 and 4) and *PrivateCredit* or *LiquidLiabilities* measure FD. The first and second rows display the estimated FD threshold values and the 95% confidence intervals, respectively. The below-threshold and above-threshold slope parameter estimates illustrate the regime-dependent marginal effects of FD on income inequality.

Columns (1) and (3) of Table 6 present the threshold impact of financial openness in the FD-income inequality relationship. Regardless of the FD measurement, the threshold value (financial openness) point estimate is 1.58 with a corresponding 95% confidence interval [-0.1, 1.75], which means that observations with financial openness of less than 1.58 are classified into the low financial openness regime, while those with greater values are classified into the high financial openness regime, which includes approximately 13% of observations. With regard to the regime-dependent marginal effects, FD appears to have a significantly positive effect on income inequality if financial openness is greater than the threshold value of 1.58. In contrast, the coefficient estimate of FD is insignificant when observations fall below the threshold level. Regarding the control variables, in both columns, the estimated coefficient of *FO* is significant and negative, suggesting that an increase in financial openness reduces income inequality. In addition, consistent with the Kuznets hypothesis, economic growth alleviates income inequality. However, the effect of *growth* on income inequality diminishes when a country's financial openness increases. Furthermore, *GovExp* increases income inequality. A previous study by Kunieda et al. (2014) has examined the impact of financial openness on FD-income inequality, although, to the best of our knowledge, the dynamic panel threshold method has not been used to examine the FD-FO-income inequality nexus.

Table 6. Income inequality, FD and Openness (Dynamic panel threshold model)

	PrivateCredit		LiquidLiabilities	
	(1) FO	(2) TO	(3) FO	(4) FO
Threshold	1.5798	3.9382	1.5798	3.9382
95% confidence interval	[-0.0979,1.7466]	[3.6001,4.9366]	[-0.0979,1.7466]	[3.6001,4.9366]
Impact of FD:				
Below threshold	-0.0014	0.0058	0.0056	0.017
	(0.0067)	(0.0054)	(0.0131)	(-0.011)
Above threshold	0.0396***	0.0110**	0.0326**	0.0227**
	(0.0115)	(0.0055)	(0.0146)	(0.0106)
Impact of covariates:				
Lag.GrossGini	0.9100***	0.8859***	0.9559***	0.9068***
	(0.0393)	(-0.056)	(0.0337)	(0.0598)
FO	-0.0581***	0.0067	-0.0380**	0.0019
	(0.0161)	(0.0041)	(0.0149)	(0.0051)
TO	-0.0058	-0.0173*	-0.0142	-0.0224*
	(-0.012)	(0.0098)	(0.0102)	(0.0124)
Growth	-0.0868*	-0.0293	-0.0051	-0.0726
	(-0.049)	(0.0401)	(0.0811)	(0.0701)
Inf	-0.0004*	0.0001	0	0.0001
	(0.0002)	(0.0002)	(0.0004)	(0.0002)
GovExp	0.0444***	0.0079	0.0241*	0.0058
	(0.0143)	(0.0085)	(0.0125)	(0.0079)
Pop	-0.0042	-0.0003	-0.0050*	-0.0014
	(0.0032)	(0.0025)	(0.0028)	(0.0024)
LifeExp	0.0003	-0.0012***	-0.0002	-0.0013***
	(0.0007)	(0.0004)	(0.0005)	(0.0005)
Gamma	0.3282	0.5442**	0.2506	0.4771**
	(0.2132)	(0.2207)	(0.1846)	(0.2419)
Observation	497	497	497	497
Country	71	71	71	71
Instrument	64	67	50	57

Notes: Table 6 reports the impact of financial openness on the FD and income inequality relationship using Dynamic Panel Threshold Regression. Private credit and liquid liabilities are used as the financial development (FD) and FO and TO are used as threshold variables. The point estimates of the thresholds and the corresponding 95% confidence intervals are reported in the first two rows respectively. The regime dependent marginal effects of FD on income inequality are shown by "Below threshold" and "Above threshold". Gamma is the intercept.
*, **, *** indicate significance at the 10%, 5% and 1% levels, respectively.
All variables are defined in Appendix A

Turning to the threshold effect of trade openness, columns (2) and (4) illustrate the estimates from the dynamic panel threshold model where trade openness is used as the threshold variable and FD is measured by *PrivateCredit* and *LiquidLiabilities*, respectively. The point estimate of the threshold value in both columns is 3.938, and approximately 73% of the observations in the sample are above this threshold value. The result shows that FD appears to have a significantly positive effect on income inequality if FD is greater than the threshold. Below the threshold, however, the effect of FD is insignificant. This result is consistent with the GMM analysis, suggesting that improving FD increases income inequality at a higher level of trade openness. With regard to the control variable, greater trade openness corresponds to lower income inequality. According to standard Heckscher–Ohlin trade theory, the inequality effect of openness varies depending on relative factor abundance and productivity differences as well as on the extent to which individuals earn income from wages or capital. In addition, the coefficient of *LifeExp*, as a proxy for human capital, is negative and significant in both columns, consistent with the previous research (Furceri and Ostry, 2019).

Overall, our findings, based on all three approaches – split-sampling, interaction and threshold analysis – emphasise the role of openness in the FD and income inequality linkage. The data suggest that with increased financial openness and trade openness, banking development benefits the richer segments of society more than the poorer ones and hence significantly increases income inequality. In other words, the widening income inequality effect of FD tends to become more significant as a country becomes more open to financial or trade markets. This finding supports the theoretical models of Kunieda et al. (2014) and Ehrlich and Seidel (2019) in countries open to trade and financial markets. In contrast to their prediction, it does not provide any evidence of a negative relationship between FD and income inequality in countries with low financial openness or trade openness.

6. Conclusion

This study examines whether the combination of a country's degree of integration into the world economy, trade integration or financial integration, and FD increase or decrease inequality within a country. If economic openness is an important determinant of income inequality and if it leads to higher FD, it raises the question of to what extent it can moderate the relationship between FD and income inequality. Accordingly, we investigate whether the income inequality effect of FD is monotonic with the level of economic openness and whether a certain economic openness threshold exists beyond which FD increases income inequality.

For this purpose, we employ two different methodologies to investigate the possible non-linearities. Initially, we split the sample into different subgroups by the level of either financial openness or trade openness and use GMM to estimate the effect of FD on income inequality for each group. While GMM helps us use the lagged dependent variable and regressors to address potential endogeneity issues, this method does not give us specific information on the threshold value at which the effect changes, if at all. For this reason, we use the dynamic panel threshold method of Kremer et al. (2013) with a GMM estimator to investigate whether there is an economic openness threshold. Using a panel of developing and developed countries for

1994–2017, our empirical result shows that in an economy closed to the world financial or trade market, growing FD does not significantly impact income inequality. However, if an economy is open to the world financial and trade market, inequality within the economy increases as its financial market develops. This finding is consistent across different econometric methods, subsamples and interaction analyses, and distinct FD indicators. In general, this study extends the current literature by providing empirical evidence on the role of openness in the FD-income inequality relationship covering both developed and developing countries. In addition, evidence of the pro-inequality impact of FD in open countries informs policymakers about the importance of redistribution policies. In fact, open countries that desire to decrease inequality resulting from having open markets should consider implementing redistribution policies to mitigate the inequality-increasing effect of FD.

Clearly, more research is necessary. First and foremost, if data quality concerns are dropped, researchers should study the effect of FD on the incomes of individuals at the top and bottom of the income distribution at different levels of economic openness. This is important because the Gini coefficient measures deviations from perfect income equality regardless of where in the distribution these deviations arise. In particular, the finding that finance increases inequality does not necessarily imply that finance ignores the poor. Another exciting extension relates to the question of whether the threshold effect of economic openness on the FD–income inequality relationship varies across different levels of institutional quality, and how and what level of institutional quality can change the nonlinear income inequality effect of FD. We leave these avenues of exploration for future research.

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Appendix

Appendix A: Descriptions and Data Sources of the variables used in this study.

Variable Name	Definition	Source
Income Inequality:		
GrossGiniCoefficient	The Standardized World Income Inequality Database	SWIID
Financial Development:		
PrivateCredit/GDP	The financial resources provided to the private sector by domestic money banks as a share of GDP. Domestic money banks comprise commercial banks and other financial institutions that accept transferable deposits, such as demand deposits.	GFDD
BroadMoney/GDP	Ratio of liquid liabilities to GDP. Liquid liabilities are also known as broad money, or M3.	GFDD
Economic Openness:		
TradeOpenness	Sum of exports and imports of goods and services measured as a share of GDP.	WDI
FinancialOpenness	the sum of external assets and liabilities as a share of GDP (These foreign assets and liabilities include foreign debt, foreign direct investment, and foreign indirect investment (portfolio investment))	Author calculation based on Lane and Milesi-Ferretti (2007)
Controls:		
Growth GDPPerCapita	Changes in Log GDP per capita (constant 2015 US$)	WDI
GovExp/GDP	General government final consumption expenditure (% of GDP)	WDI
Inf	Inflation, consumer prices (annual %)	WDI
Population	Population, total	WDI
LifeExp	Life expectancy at birth, total (years)	WDI

Appendix B: Country list based on LFO, HFO, LTO and HTO.

LFO	HFO		LTO	HTO	
Albania	Armenia	Jamaica	Australia	Albania	Germany
Bangladesh	Bolivia	Japan	Bangladesh	Bolivia	Ghana
Brazil	Costa Rica	Lesotho	Brazil	Chile	Jamaica
China	Cote d'Ivoire	Malaysia	China	Cote d'Ivoire	Korea, Rep.
Colombia	Ecuador	Mongolia	Colombia	Dominican Rep	Norway
Dominican Rep	El Salvador	New Zealand	Egypt	Ecuador	Paraguay
Egypt	Eswatini	South Africa	India	France	Philippines
Ghana	Georgia	Ukraine	Iran	Greece	Poland
India	Honduras	United States	Italy	Indonesia	Sweden
Indonesia	Korea, Rep.	Uruguay	Japan	Kenya	Tunisia
Iran	Mexico	Austria	Nigeria	Mexico	Belgium
Kenya	Paraguay	Belgium	Pakistan	New Zealand	Bulgaria
Nigeria	Peru	Finland	Peru	Portugal	Czech Republic
Pakistan	Philippines	France	Tanzania	Romania	Eswatini
Sri Lanka	Poland	Germany	Turkey	South Africa	Honduras
Tanzania	Romania	Hong Kong, China	Uganda	Spain	Hong Kong, China
Turkey	Thailand	Hungary	United States	Sri Lanka	Hungary
Uganda	Tunisia	Ireland	Uruguay	United Kingdom	Ireland
	Australia	Netherlands		Armenia	Lesotho
	Barbados	Norway		Austria	Malaysia
	Botswana	Panama		Barbados	Mongolia
	Bulgaria	Portugal		Botswana	Netherlands
	Chile	Singapore		Costa Rica	Panama
	Czech Republic	Spain		El Salvador	Singapore
	Greece	Sweden		Finland	Switzerland
	Italy	Switzerland		Georgia	Thailand
		United Kingdom			Ukraine