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The income inequality, financial depth and economic growth nexus in China

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KEYWORDS

ARDL Bounds, Granger causality, Income inequality, Financial Depth, Economic Growth

1 | INTRODUCTION

China has rapidly opened to the world since the onset of its economic reforms, resulting in accelerated economic growth and development (Siebert, 2007). As the country experiences almost a double-digit growth rate, there is a noticeable, corresponding increase in the level of income inequality. Many studies find rising inequality to impede future development and possibly a precursor to social tension or political instability (Gu, Dong, & Huang, 2015; Jian, Sachs, & Warner, 1996; Kanbur & Zhang, 2005). In this view, the Chinese government has set out to reverse the rising inequality. The government aims to build a more "harmonious socialist society" (11th five-year plan) and reinforces the concept of inclusive growth (12th and 13th five-year plans) as one of its key development goals (Kanbur, Rhee, & Zhuang, 2014). A wide range of government policies reflects this commitment, such as the *dibao*[1] system, subsidy to support compulsory education and elimination of agricultural taxes to help rural farmers.

Additionally, greater financial reforms[2] are observed in China. Regulation is passed to make the financial system more competitive and stable. The increase in the availability of credit to public enterprises, firms and households further confirms the development of the financial sector in China. Low-interest loans available from local banks or financial institutions have reduced poverty levels amongst urban and rural households (Zhang & Loubere, 2015). While the exceptional growth rate for the past

[1]"*Dibao*" or urban minimum living standard guarantee programme is an initiative by the Chinese government to help the poor come out of poverty (Wang, 2007).

[2]China also plays a pivotal role in developing the financial sector in the region and many countries have been benefitting from China's economic growth (Koh & Kwok, 2017; Kwok & Koh, 2017). For instance, China continues to play a crucial role in the Asian Infrastructure Investment Bank (AIIB), a multilateral development bank which aims to support infrastructure development in the Asia–Pacific region (Stiglitz, 2015). The Yuan has also been recognised as a major reserve currency after it was added to the International Monetary Fund (IMF) basket of currencies which represents China's increasing status in global financial markets (Tobin, 2013).

few decades is evidence that China is moderately successful in its policies, it also prompts us to question its impact on inequality. President Xi Jinping recently declared that the country needs a "crucial rebalancing to embrace a *new normal* growth phrase" (Hu, 2015). This will ensure that benefits from economic development are more evenly distributed, which is necessary as the country integrates further with the world's economy.[3]

Our study differs from the existing literature in two ways. First, we expand the discussion of inequality-growth-financial depth using theoretical considerations from Kuznets (1955) and Greenwood and Jovanovic (1990) using a two-step procedure of the autoregressive distributed lag (ARDL) bounds and Granger causality. In many instances, a time series analysis will provide deeper insights (Ang & McKibbin, 2007; Arestis & Demetriades, 1997) into the relationship between the three variables compared to cross-country regressions in terms of policy implication. The ARDL is suitable here since macroeconomic variables often reflect its past behaviour and should be seen as a dynamic and autoregressive process (Narayan & Smyth, 2006).

Conceptually, a well-developed financial sector leads to long-run economic growth by easing the ability of firms to access capital. Since capital is an essential input in the production function, higher capital accessibility will lead to higher productivity and growth. As a result, economic growth and financial depth may have an inequality-widening effect temporarily since credit is often provided on condition of available collateral. When the country achieves high economic growth, the benefits are trickled down to other individuals in the society leading to an inequality-narrowing effect (Kuznets, 1955). This argument is also in line with the theoretical stipulations of Greenwood and Jovanovic (1990). Greenwood and Jovanovic (1990) hypothesised an inverted U-shaped relationship between financial sector development and inequality (resembling Kuznets' hypothesis). According to the authors, the country's financial markets formalise with greater economic growth as the income gap widens. The maturity stage is characterised by a well-structured financial market and financial intermediaries. Finally, the income gap will stabilise as the country's growth reaches a higher level. There is a possibility that this experience may not be evident in China and the country may move away from the inequality-narrowing effect as predicted by the inverted U-shape hypothesis (Kotarski, 2015). Earlier papers such as Liu, Liu, and Zhang (2017) look at the different aspects of financial development to examine the inequality-financial structure relationship using a dynamic GMM approach. Nonetheless, the financial system often helps accelerate economic growth through the expansion of economic opportunities (Beck, Demirgüç-Kunt, & Levine, 2007). Furthermore, the literature on the direction of causality looking at either income inequality-finance, finance-growth or inequality-growth has so far provided mixed evidence (Chang, 2002; Jalil & Feridun, 2011; Khalifa Al-Yousif, 2002; Wan, Lu, & Chen, 2006).

Second, the paper contributes in terms of econometric strategy. The bounds approach developed in Pesaran and Pesaran (1997) and subsequently expanded in Pesaran, Shin, and Smith (2001) offers several advantages in comparison with other conventional cointegration techniques.[4] The restrictive assumption that all variables must be integrated in the same order is relaxed here. The bounds test can be used irrespective of whether the variables are $I(0)$ or $I(1)$. This technique provides unbiased estimates as it simultaneously corrects for residual serial correlation and problem of endogenous variables (Pesaran & Shin, 1999). Additionally, while the conventional cointegration techniques estimate the long-run relationships within the context of a system of equations, the ARDL method employs a single reduced form equation.

[3]"The Chinese century is not at the beginning of the end; it is at the end of the beginning" (Hu, 2015).

[4]The Engle–Granger (1987) single equation model may be problematic when there are more than two variables, and the number of cointegration vectors is unknown (Asteriou & Hall, 2007; Harris, 1995). The multivariate approach developed by Johansen (1988), though widely used, is sensitive to lag length (Hjalmarsson & Österholm, 2010).

Since the robustness of standard unit root test is often questioned in the possibility of structural breaks in the series, we utilise the Narayan and Popp (2010; NP) unit root statistical test to detect the presence of structural breaks to account for structural changes that might have taken place during our period of study. The NP (2010) is a new augmented Dickey-Fuller (ADF) test for unit roots, which introduces two structural breaks compared to the popular Zivot and Andrews (1992) method, which accounts for one structural break. The NP test has an advantage over competing unit root tests since it does not require a priori specification of the possible timing of structural breaks as it is endogenously determined within the model (Narayan & Popp, 2010). Since information on the break dates is crucial in correctly specifying our model, the break dates are included as dummy variables in the cointegration model.

From a policy viewpoint, China's rapid increase in inequality is an issue the government is looking into (Yang & Greaney, 2017) and modelling the short-run and long-run adjustment of the variables provides an additional layer of pertinent information. This information may provide an impetus for the government to increase effort to redistribute income via welfare spending. China has developed and transformed from an equal society to one of the most unequal society in less than fifty years. Economic reforms and financial liberalisation may have some direct or indirect impact on income distribution. If financial depth or economic growth is found to Granger cause income inequality, the Chinese government may want to relook its current economic stance. Better financial depth which promotes economic growth will eventually offset any increase in income inequality in the long run. Besides, the country's economy has shown positive signs with millions lifted out of poverty. Policy implications will be more forceful if there is no long-term adverse effect of this unrestrained growth.[5]

The rest of the study is organised as follows: Section 2 provides a review of the current literature. Section 3 describes the data and estimation technique followed by empirical analysis and discussion of the results. Finally, Section 4 concludes the study.

2 | LITERATURE REVIEW

2.1 | Gini and economic growth

Theoretically, there are several reasons how economic growth can affect the distribution of income. While the coexistence of rapid growth and high inequality in China is worrying, some scholars link this observation to a possible emergence of a Kuznets curve. The inverted U hypothesis is first observed in a seminal paper by Nobel Laureate Simon Kuznets (1955), who argues that early stages of economic growth are characterised by rising inequality, while later stages are associated with lower levels of inequality. His theoretical prediction assumes a transition from agriculture to high productive industries. Although such shift may lead to a temporal increase in income gaps as most gains only benefit certain segments of the society, over time the rise in inequality will stabilise and narrow in the later phases of development.

According to Yang and Greaney (2017), economic growth can affect inequality through two other channels. First, economic growth benefits the rich more through capital gains and leads to inequality-widening effect. Second, economic growth can help the poor via employment opportunities leading to inequality-narrowing effect. Empirical studies which investigate the relationship between *Gini* and economic growth produced mixed results. Several authors reported a negative relationship between growth and inequality. Wan et al. (2006) used provincial panel data and found that inequality negatively affects growth both in the short run, medium run and long run. According to the authors,

[5]An earlier paper by Gozgor and Ranjan (2017) found that redistribution has increased in tandem with globalisation.

financial reforms could not be fully executed as banks were unable to extend credit in the face of high non-performing loans (NPLs) in some of the provinces in China. This deterred the private sector from carrying out productive investment in physical or human capital due to credit market imperfections (Galor & Zeira, 1993). Nonetheless, their study failed to determine the reverse effect of growth on inequality in their model.

The effect of growth on inequality could also be positive. Chen (2010) used a time series vector autoregressive (VAR) model and reported that growth reduces inequality in the long run. The author posits that economic growth increases the country's tax base, and since the revenue from the tax can be reinvested to reduce regional disparities, levels of inequality will fall. Other studies look at the effect of a public policy or trade openness on regional inequality in China (Tsui, 1991; Wei, 2013; Zhang & Zhang, 2003). Most of these studies agree that openness has contributed to growth at the expense of widening income gaps. High levels of inequality also mean that the bottom 10% of society may not be able to afford financial investment or human capital development (Wang, Wan, & Yang, 2014).

On the other hand, Adelman and Sunding (1987) confirmed the existence of the Kuznets existence whereby inequality initially reduces due to land reforms policies but increases after the capital-intensive industrialisation process. The authors posit that income inequality changes in line with different policy regimes and speculate that as urban industrial sector gain importance, the level of inequality in the country will increase. Chen and Fleisher (1996) found no consistent evidence of a Kuznets pattern. Yang and Greaney (2017) reviewed the inequality-growth-redistribution relationship in the United States, Japan, South Korea and China. Their results supported the S-curve hyphothesis relating growth to inequality with different initial points for the four economies. On the whole, the evidence from the literature on the relationship between Gini-growth is mixed.

2.2 | Growth and financial depth

The theoretical foundation between income inequality and finance nexus was laid out in Schumpeter (1911). In his argument, the financial intermediary will mobilise savings and facilitate investments which in turn will spur economic growth. Additionally, Greenwood and Jovanovic (1990) build upon earlier views provided by Goldsmith–McKinnon–Shaw and posit the interconnectedness between financial intermediation and economic growth. Several studies find a link between financial system-growth-income inequality (Demirguc-Kunt & Levine, 2008; King & Levine, 1993). The discussion often revolves around the ability of firms to access credit which leads to higher investment in physical capital which in turn leads to an increase in productivity and eventually higher economic growth.

Literature provides inconclusive results with regard to the relationship between financial depth and economic growth. A well-functioning financial system is linked to economic growth since financial systems develop in anticipation of future economic growth (Calderón & Liu, 2003; Hassan, Sanchez, & Yu, 2011; Levine, 1997). As the financial sector develops, there is a possibility that the poor will have better chances of obtaining credit as it improves access for the rich and poor alike. Additionally, Habibullah and Eng (2006) use a panel of 13 Asian countries to investigate the relationship between financial development and economic growth for the period 1990 to 1998. Using the ratio of domestic credit to GDP as a proxy to measure financial development, the authors find that financial development leads to higher growth. Similarly, such findings are reported in an earlier study by Christopoulos and Tsionas (2004).

A few studies did not find evidence or limited support to establish the relationship between these two variables. For instance, Shan (2005) used variance decomposition and impulse response function to investigate the relationship and found little evidence to support the hypothesis that financial development impacts growth. There was no substantial difference between the Western countries and Asian countries in the sample. Chang (2002) uses a multivariate VAR model for China and find a single cointegrating vector amongst GDP, financial development and openness. However, the results from Granger causality found independence between financial development and economic growth. Murinde and Eng (1994) studied the relationship between financial development and economic growth for Singapore and found support for supply lending hypothesis, whereby financial development spurs growth (limited to the use of M1 and some of the monetisation variables—v1 and v3 as a proxy for financial development). Otherwise, the authors found inconclusive evidence to support the hypothesis. Petkovski and Kjosevski (2014) investigate the relationship between financial development and economic development using the transitional economies in Central, South and Eastern Europe. The authors did not find evidence that banking innovation increases economic growth. The authors posit the possibilities of large NPLs and financial crises which affected these economies during their period of study.

Liang and Jian-Zhou (2006) employed a multivariate vector autoregressive (VAR) framework to evaluate the long-run relationship between financial development and growth in China from 1952 to 2001. Their empirical results suggest a unidirectional causality from economic growth to financial development. This observed relationship may be due to the concentration of credit to state-owned enterprises under the government's guidance. Since the government may use the availability of credit to reduce regional inequality, it may lead to large amounts of NPLs. If so, most credit is not being issued to productive investment and may not spur economic growth.

2.3 | Gini and financial depth

According to Greenwood and Jovanovic (1990), financial depth can either increase inequality (inequality widening effect) or reduce inequality (inequality narrowing effect). In their version of the Kuznets hypothesis, a country experiences an income widening effect as it moves to a developed economy as financial intermediaries promote growth through the availability of credit.

Access to credit can have an inequality-widening effect or an inequality-narrowing effect. Low-interest loans available from local banks or financial institutions may reduce poverty levels and ensure that benefits from economic development are more evenly distributed leading to an inequality-narrowing effect. Financial reforms will also lead to better-organised allocation of financial resources followed by a reduction in inequality. Liang (2006) and Clarke, Xu, and Zou (2006) found that financial development would improve the income distribution in China as the development of financial markets provides better credit access to the urban poor. Deregulation allows more competition in the lending market and better pricing of loans which leads to a reduction of NPLs (Song, Storesletten, & Zilibotti, 2011).

Additionally, access to credit may also have an inequality-widening effect since access to credit often comes with a hefty collateral requirement, thus limiting access for the poor. Several studies found that inequality worsens as a result of financial development (Wahid, Shahbaz, Shah, & Salahuddin, 2012). Wahid et al. (2012) find that financial sector development in Bangladesh increases inequality, whereby it worsens by 0.17% for every 1% increase in domestic credit. According to the authors, income inequality worsens as financial development relatively benefits the rich more than the poor. The high set-up costs impede poor individuals from securing credit from financial intermediaries, and this "pushes them further into the income inequality trap" (Wahid et al., 2012, p. 90). Furthermore, such credit constraints not only deter the poor from obtaining high returns projects but also reduce the efficiency of capital allocation.

According to Claessens and Perotti (2007), unequal access to financial services is observed in countries with a high level of inequality. Such countries may have certain groups of individuals who exert a political influence which in turn distort the institutional environment thereby leading to unequal access to finance and reinforcing any initial income inequality. Similar results were reported in Canavire-Bacarreza and Rioja (2008), whereby financial development in Latin America did not have an impact on the income of the poorest quintile of the population but benefited the middle or higher quintiles. The authors posited that this observed relationship might be due to the poor relying on other forms of lending rather than commercial banks or financial institutions. Besides, Ang (2010) posited that credit constraints might increase income inequality since the poor have difficulty securing financial services or obtain credit as they lack collateral. Jahan and McDonald (2011) explain that if financial services are not accessible to the poor, it may not have much impact on reducing inequality.

3 | METHODOLOGY AND EMPIRICAL FINDINGS

The long-run and causal relationships between income inequality, economic growth and financial depth in China will be performed in two steps. First, we test the long-run relationships amongst the variables by using the ARDL bounds testing approach of integration. Second, we test the causal relationships by using error-correction-based causality models. Following Granger (1988), if the variables in the model are cointegrated, the Granger causality should be estimated using a vector error-correction modelling rather than a VAR.

3.1 | Data description

We rely on the *Gini* index from World Bank PovcalNet dataset. When PovcalNet data are not available, we utilise Milanovic's (2014) dataset since it extends the PovcalNet. If data on the year surveyed are not available, the data prior to the year surveyed are used.

In line with several studies, the paper relies on domestic credit to the private sector (% of GDP) as the preferred indicator for financial depth (Demirguc-Kunt & Levine, 2008; King & Levine, 1993). Recent literature often does not distinguish between financial development and financial depth although they are different concepts. For instance, Loayza, Ouazad, and Rancière (2018) state that financial development is a “general term” and financial depth is more common and popularly measured using a ratio of credit over GDP.[6] Financial depth which is the focus of our study best captures the “financial sector relative to the economy” (World Bank, n.d.) and may impact long-term growth or income inequality.

We obtain the proxy for financial depth from the Global Financial Development Database (World Bank GFDD)[7], whereby higher values indicate higher credit available mainly because banks and financial institutions play a crucial role to provide financial functions (Levine, 1997). Domestic credit to the private sector refers to “financial resources provided to households and businesses by financial corporations in the form of loans, purchases of non-equity securities, trade credits and other accounts receivable. Additionally, credit to the private sector may sometimes include credit to state-owned or partially state-owned enterprises” (Khaltarkhuu, 2014). This variable is chosen over other indicators since it includes credit to government-linked enterprises, which is a crucial feature of the Chinese

[6]Levine, Loayza and Beck (2000) discusses the components of financial development and how it influences economic growth. It was again emphasised in a recently published book on the relationship between financial and real sector development (Beck & Levine, 2018).

[7]A complete description of the database is available from Čihák, Demirgüç-Kunt, Feyen, & Levine (2012). *Benchmarking financial systems around the world*. World Bank Policy Research Working Paper, (6175).

economy. A sound signal of a healthy economy is often indicated by the availability of credit to fund productive investments or spur private consumption.

The growth of real GDP per capita is used to measure economic growth (constant 2005 US$). We use annual data from 1980 to 2013 (34 observations). Average years of schooling is added as a control variable since there is a tendency for it to affect the dependent variable in terms of the human capital endowment. The data on education are obtained from Barro and Lee (2013). The Microfit 5.0 statistical software is utilised to conduct the bounds test and Granger causality test. The data are converted to logarithm before the estimation process.

3.2 | Unit root test with structural breaks

The first step in the cointegration technique is to examine the time series properties of the respective variables. Even though the ARDL bounds test is suitable for time series modelling with $I(0)$ and $I(1)$ variables, it is based on the assumption that none of the variables are integrated to the order of 2 (Pesaran & Shin, 1999).

The standard ADF test does not account for structural breaks and uses a parametric autoregression to approximate the structure of errors (Pesaran & Pesaran, 1997). Since time series variables may be sensitive to structural breaks and ignoring them may result in spurious regression (Berg, Ostry, & Zettelmeyer, 2012; Perron, 1989), the Narayan and Popp (2010) endogenous unit root test with two structural breaks is also utilised to determine the order to integration. The authors use Monte Carlo simulations to confirm that the proposed unit root test has the correct size and stable power and identifies the structural breaks accurately. The results of the ADF test and NP test are presented in Table 1.

The results of the two-break unit root are reported in Table 1. The NP (2010) considers different specifications for the trending data. M1 allows for two structural breaks in the level of the series, while M2 accounts for breaks in the levels and slope. The results from M1 revealed that the unit root null hypothesis for *Gini* and *Growth* is not rejected at the 5% level, while results from M2 revealed that the unit root null hypothesis for *Gini* and *Growth* is not rejected at the 1% and 5% level, respectively. The NP (2010) test also detected structural breaks for *Gini* (1995, 2001 and 2004), growth (1988, 1990 and 1992) and financial depth (1993 and 2005). We include the structural break dates as control variables in our model.

The structural breaks detected for *Gini* coincide with the household responsibility system (HRS) implementation which was seen as an extraordinary move in rural poverty reduction (Lin, 1992; Ravallion & Chen, 2007), China's ascension into the World Trade Organization in 2001 and stricter implementation of the minimum wage (2004). The structural break dates for growth are in line with the country's gradual approach to privatisation (1988, 1990) and the restructuring of state-owned enterprises exercise (1992). The structural break for financial depth in 1992–93 coincides with the

TABLE 1 NP (2010) two-break unit root test

	M1				M2			
Series	*t*-stats	*TB*1	*TB*2	*K*	*t*-stats	*TB*1	*TB*2	*K*
Gini	−3.492	2001	2004	0	−6.028*	1995	2004	0
Growth	−3.831	1988	1990	5	−3.227	1988	1992	5
Financial depth	−5.406*	1993	2005	1	−5.373**	1993	2005	1

Notes: Critical values (CV) are obtained from NP (2010). The 1% CV are −5.259 (M1) and −5.949 (M2), the 5% CV are −4.514 (M1) and −5.181 (M2), and the 10% CV are −4.143 (M1) and −4.789 (M2). TB1 and TB2 are the detected dates of the breaks. *k* is the number of optimal lag.

*, **, *** Denotes the rejection of the null hypothesis of a unit root at the 1%, 5% and 10% level.

government's adoption of a 16-point programme credit plan which imposes individual credit ceilings on specialised banks (Montes-Negret, 1995). The structural break date in 2005 could be linked to the start of the property bubble in both commercial and real estate in the country.

3.3 | Cointegration

The bounds test approach estimates the unrestricted error-correction model (UECM) as follows:

$$\begin{aligned}\Delta Gini_t = a_{oGini} + \sum_{i=1}^{n} b_{iGini} \Delta Gini_{t-i} + \sum_{i=1}^{n} c_{iGini} \Delta FD_{t-i} + \sum_{i=1}^{n} d_{iGini} \Delta Growth_{t-i} \\ + \sigma_{1Gini} Gini_{t-1} + \sigma_{2Gini} FD_{t-1} + \sigma_{3Gini} Growth_{t-1} + \sigma_4 Control + \sigma_5 D_{Ginit} + \alpha t + \varepsilon_{1t}\end{aligned}, \quad (1)$$

$$\begin{aligned}\Delta Growth_t = a_{oGrowth} + \sum_{i=1}^{n} b_{iGrowth} \Delta Growth_{t-i} + \sum_{i=1}^{n} c_{iGrowth} \Delta FD_{t-1} + \sum_{i=1}^{n} d_{iGrowth} \Delta Gini_{t-i} \\ + \sigma_{1Growth} Growth_{t-1} + \sigma_{2Growth} FD_{t-1} + \sigma_{3Growth} Gini_{t-1} + \sigma_4 Control + \sigma_5 D_{Growth} + \alpha t + \varepsilon_{2t},\end{aligned} \quad (2)$$

$$\begin{aligned}\Delta FD_t = a_{oFD} + \sum_{i=1}^{n} b_{iFD} \Delta FD_{t-i} + \sum_{i=1}^{n} c_{iFD} \Delta Growth + \sum_{i=1}^{n} d_{iFD} \Delta Gini_{t-i} \\ + \sigma_{1FD} FD_{t-1} + \sigma_{2FD} Growth_{t-1} + \sigma_{3FD} Gini_{t-1} + \sigma_4 Control + \sigma_5 D_{FDt} + \alpha t + \varepsilon_{3t}\end{aligned}. \quad (3)$$

In Equations (1)–(3), Δ reflects the first difference operator, *Gini* is the log of the *Gini* index, Growth is the changes in log of *GDP* per capita, and *FD* is the log of domestic credit to the private sector for our model. $\varepsilon_{1t},\ \varepsilon_{2t},\ \varepsilon_{3t},$ are serially independent random errors with mean zero and finite covariance matrix. Education is added in the model as a control since several studies (Hou, Li, & Wang, 2018; Ning, 2010) found that it may be related to the dependent variable, and αt is the deterministic trend for our model. In the above equations, D_{Gini}, D_{Growth} and D_{FD} represent the structural breaks detected in the preceding section.

In Equation (1), when *Gini* is the dependent variable, the null hypothesis of no cointegration is H_0: $\sigma_{1Gini} = \sigma_{2Gini} = \sigma_{3Gini} = 0$, and the alternative hypothesis of cointegration is $H_1 = \sigma_{1Gini} \neq \sigma_{2Gini} \neq \sigma_{3Gini} \neq 0$. On the other hand, in Equation (2), when Growth is the dependent variable, the null hypothesis of no cointegration is H_0: $\sigma_{1Growth} = \sigma_{2\text{Growth}} = \sigma_{3Growth} = 0$, and the alternative hypothesis of cointegration is $H_1 = \sigma_{1Growth} \neq \sigma_{2Growth} \neq \sigma_{3Growth} \neq 0$. In Equation (3), when FD is the dependent variable, the null hypothesis of no cointegration is H_0: $\sigma_{1\text{FD}} = \sigma_{2FD} = \sigma_{3FD} = 0$ and the alternative hypothesis of cointegration is $H_1 = \sigma_{1FD} \neq \sigma_{2FD} \neq \sigma_{3FD} \neq 0$.

To identify the evidence of a long-run relationship in Equations (1)–(3), an *F* test for the joint significance of the coefficients of the lagged levels of variables is conducted. Two sets of critical values for a given significance level are determined. The first level is calculated based on the assumption that all variables included in the model are integrated of order 0, while the second one is calculated based on the assumption that the variables are integrated of order 1. The null hypothesis of no cointegration is rejected when the value of *F* test statistic exceeds the upper critical bounds value, and it is accepted when the *F* test is lower than the lower bound value. The bounds test results are reported in Table 2.

The bounds test is applied with an unrestricted intercept and with trend. To determine the lag length, the study uses a priori lag selection procedure (SBC). Since exact critical values are not available for a mix of *I*(0) and *I*(1) variables, Pesaran et al. (2001) reported critical values for asymptotic distribution for F-statistics. Lower bound values are based on the assumption that the variables are *I*(0), while upper bound values are based on the assumption that the variables are *I*(1). The study uses

TABLE 2 Bounds test and Lagrange multiplier (LM) test

	F-stats	Outcome	LM test (3)
Model: *Gini*, *Growth*, *FD*			
Equation (1) (*Gini*|*Growth*, *FD*)	5.832	Cointegrated	4.3694 [0.224]
Equation (2) (*Growth*|*Gini*, *FD*)	2.864	Not cointegrated	
Equation (3) (*FD*|*Gini*, *Growth*)	4.636	Not cointegrated	
F test critical values (Case IV)	*I* (0)	*I* (1)	
Critical values at 1%	6.328	7.408	
Critical values at 5%	4.433	5.245	

Notes: If the estimated *F*-statistic is higher than the upper bound of the critical values, then the null hypothesis of no cointegration is rejected. If the estimated *F*-statistics fall below the upper bound of the critical values, then the null hypothesis of no cointegration cannot be rejected. The critical values for the lower bounds *I*(0) and upper bounds *I*(1) are taken from Case IV Narayan (2005).

the small samples critical values akin to Pesaran et al. (2001) as tabulated by Narayan (2005) due to the small sample size of the study (34 observations).

For the bounds test to be valid (and cointegration to exist), there should be at least one long-run relationship reported, whereby the computed F-stats exceeds the upper bound critical value. For Equation (1) (*Gini*|*Growth*, *FD*), the *F* test = 5.832; for Equation (2) (*Growth*|*Gini*, *FD*), the *F* test is 2.864; and for Equation (3) (*FD*| *Gini*, *Growth*), the *F* test is 4.636.

Since the *F* test for Equation (1) is higher than the upper bound critical value (5.245-Case IV) at the 5% level, there is a long-run cointegrating relationship amongst the variables when Gini is considered as a dependent variable in the model. However, the null hypothesis is not rejected for Equations (2) and (3). The presence of a cointegration relationship here shows that long-run Granger causality exists in at least one direction amongst the variables.

Next, the Lagrange multiplier (LM) test is applied to ensure that the cointegrating relationships are not autocorrelated since it is a key assumption of the bounds test. Furthermore, if serial correlation is detected in the residuals (the difference between the observed and the estimated value), it is expected that e_t is related to e_{t-1} which leads to model misspecification. Therefore, the study does not expect to reject the null (H_0: $\rho = 0$) in the LM test. Table 2 reports that the null of no autocorrelation (up to lag order 3) cannot be rejected at the 1% significant level and the cointegrating relationships (of *Gini*|*Growth*, and *FD*) satisfy the required condition of no autocorrelation.

The bounds test (Table 2) has indicated the presence of a cointegrating relationship amongst the variables but does not indicate the causal direction. Thus, the study relies on the Granger causality test to examine the causal direction between the variables. Granger causality also provides information on the short-run and the long-run relationship of the said variables[8].

3.3.1 | Granger causality

Since the bounds test indicates that cointegration does not exist when growth or financial depth is treated as a dependent variable, the short-run[9] causal relationships between Growth and financial

[8]Assume there are two-time series y_t and x_t. A variable x_t Granger causes y_t if y_t can be better predicted using the lagged values of both x_t and y_t than it can using the lagged values of y_t alone.

[9]An earlier study by Christopoulos and Tsionas (2004) found that ignoring the short-run effects and testing only for the long-run causality may lead to wrong conclusions. The rationale given is that most benefits of higher levels of financial development could be realised in the short run but slowly disappears with higher economic development (Darrat, 1999).

TABLE 3 Results of short-run and long-run Granger causality test

	Wald tests χ^2				
	Δ*GINI* (prob. values)	**Δ*Growth* (prob. values)**	**Δ*FD* (prob. values)**	**ECT_{t-1}**	***t*-stats**
Dependent variable					
Δ*Gini*	–	5.972	10.52	–	−7.00
		[0.02]**	[0.00]***	1.940***	[0.00]
Δ*Growth*	9.063	–	4.2881	–	
	[0.00]***		[0.04]**		
Δ*FD*	0.175	5.226	–	–	
	[0.675]	[0.02]**			

Notes: ***Significance at 1%, **Significance at 5%, *Significance at 10%. In the equation when *Gini* is the dependent variable, the *t*-statistic on the coefficient of the lagged error-correction term (*ECT*) indicates the statistical significance of the long- run causal effect. Figures in parenthesis are the *p*-values. Estimated long-run coefficients using ARDL (4,1,4) are selected using SBC.

depth are modelled using an unrestricted vector autoregressive framework (VAR). At the same time in the case of cointegration, when *Gini* is treated as the dependent variable (Table 2), the error-correction term must be included in the VAR model, and the model now becomes a vector error-correction model (VECM) or a restricted VAR. A lagged error-correction term (ECT_{t-1}) is included since it examines the short-run deviations from their long-run equilibrium path. The *ECT* is included to capture the short-run deviations of *Growth*, *FD* and *Gini* from their long-run equilibrium path:

$$\Delta Growth_t = v + \sum_{i=1}^{m} \vartheta_i \Delta Growth + \sum_{i=1}^{o} \varphi \Delta FD_{t-i} + \sum_{i=1}^{p} \psi_i \Delta Gini_{t-i} + \varepsilon_{1t}, \tag{4}$$

$$\Delta FD_t = v + \sum_{i=1}^{m} \varphi_i \Delta FD_{t-i} + \sum_{i=1}^{o} \vartheta_i \Delta Growth_{t-i} + \sum_{i=1}^{p} \psi_i \Delta Gini_{t-i} + \varepsilon_{2t} \tag{5}$$

$$\Delta Gini_t = v + \sum_{i=1}^{p} \psi_i \Delta Gini_{t-i} + \sum_{i=1}^{m} \vartheta_i \Delta Growth_{t-i} + \sum_{i=1}^{o} \varphi \Delta FD_{t-i} + \pi ECT_{t-1} + \varepsilon_{3t}, \tag{6}$$

where ε_1, ε_2 and ε_3 are serially uncorrelated error terms. The ECT_{t-1} is the lagged error-correction term obtained from the long-run cointegrating relationship, and other variables are as defined earlier. For Equations (4)–(6), to model the short-run structure of the variables, the dependent variable is regressed against its lagged values and lagged values of the independent variables.

The VEC modelling approach allows us to differentiate between short-run and long-run Granger causality. The Wald test chi-square of the differenced explanatory variables indicates the significance of the short-run causal effects, whereas the long-run causality is implied through the significance of the *t* test of the lagged ECT. The ECT contains long-term information since it is derived from the long-run cointegrating relationship.

Table 3 examines the short-run and long-run Granger causality results. The optimal lag length used in the VAR model and VECM were automatically selected by SBC. To test the hypothesis of interest, it is convenient to use the Wald form of the F-statistics for testing a set of linear restrictions. The study follows Narayan and Smyth (2006) and reports the Wald[2] chi-square test of the lagged explanatory

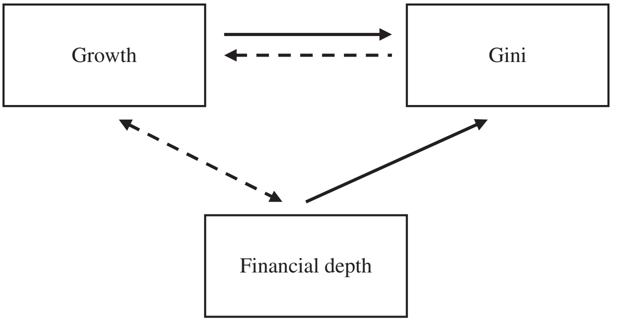

FIGURE 1 Short-run and long-run Granger causality: *Gini*, *Growth* and Financial depth. A ⇢ B indicates that Granger causality runs from A to B in the short run (unidirectional). A → B indicates that Granger causality runs from A to B in the long run (unidirectional). A ↔ B indicates Granger causality runs from A to B and B to A in the short run (bidirectional)

variables. The Wald test chi-square indicates the significance of the short-run causal effects. The *t*-statistics on the coefficients of the lagged ECT (Equation (6)) specifies the significance of the long-run causal effects.

In the short run, a change in *Gini* is associated with a change in growth at a 1% significance level. Consistent with Yang and Greaney (2017), we find that increased inequality increases economic growth. The authors posit that through capital accumulation, an increase in inequality will spur growth. Findings also revealed that a change in financial depth is associated with a change in growth at the 5% significance level, which is in line with Beck et al. (2007). Economic reforms may lead to an unequal economy since the government only started implementing inclusive growth policies in the 11th five-year plan (2006–2010) by giving priority to equitable income distribution (Lee, Lee, & Park, 2014). At the same time, we observe a bidirectional effect between financial depth and growth in the short run. The bidirectional effect between financial depth and growth can be explained by the country's massive infrastructure development to lift the economy. The country has embarked on infrastructure development and expansion through the availability of credit. To put this result into perspective, China was an egalitarian society before 1978 (Mah, 2013). China's development strategy from agriculture reforms to industrial and financial development may have caused an increase in income inequality over these few decades. These reforms were meant to spur economic growth and came with the acceptance that wealth will have a trickle-down effect (Hong, 2016).

In the long run, the coefficient of the lagged ECT_{t-1} (lagged error-correction term) is negative and statistically significant when *Gini* is treated as the dependent variable. It demonstrates that there is a long-run relationship between the variables. This result implies that causality runs through ECT_{t-1} from growth and financial depth to *Gini* at 5% and 1% level, respectively. Since the *ECM* result shows the speed of adjustment back to the long-run equilibrium after a short-run shock, the coefficient value of −1.940 indicates a rapid adjustment process. This means that if *Gini* deviates from its long-run equilibrium in the current period, convergence will be very rapid as the disequilibria will be corrected in the subsequent period. The relationship between *Gini*, Growth and Financial depth is illustrated in Figure 1.

3.3.2 | Autoregressive distributed lag (ARDL)

The long-run elasticities of *Gini*, *Growth* and financial depth are computed using the ARDL approach, and the results are reported in Table 4. In order to investigate the long-run properties of *Gini*, *Growth* and financial depth, we apply the following equation:

TABLE 4 ARDL long-run coefficient/elasticity treating *Gini* as the dependent variable

Variable	Coefficient	*t*-stats [prob]
Growth	0.033***	3.067 [0.01]
Financial depth	0.515***	5.449 [0.00]

Notes: ***Significance at 1%; **Significance at 5%; *Significance at 10%.

$$Gini = k_0 + \sum_{i=1}^{p} a_i Gini_{t-i} + \sum_{j=0}^{m} b_j Growth_{t-j} + \sum_{k=0}^{o} c_k FD_{t-i} + \varepsilon_{1t} \tag{7}$$

The long-run statistics reveal that growth and financial depth determines *Gini*. The coefficient of economic growth at 0.033 is positive and statistically significant at the 1% level. Similarly, the coefficient of financial depth at 0.515 is positive and statistically significant at the 1% level. Our results suggest that in the long run, an increase in both growth and financial depth leads to a higher level of inequality.

Our findings confirm Kuznets's (1955) postulation that as a higher proportion of population moves towards urban areas or into modern sectors in the course of development, income inequality will widen. Similarly, our findings confirm Greenwood and Jovanovich's (1990, p. 1076) posit that "in the transition from a primitive slow-growing economy to a developed fast-growing one, a nation passes through a stage in which the distribution of wealth across the rich and poor widens" since investments may be well taken by individuals or firms owing to capital accessibility. These investments lead to increased productivity and promote further growth, such that those with access to credit will benefit first (creating a widening gap); though eventually, everyone in the society will benefit due to the trickle-down effect discussed earlier. Hence, Lim and McNelis (2016) predicted that the income gap will improve as the country reaches a critical threshold in capital intensity. However, we were not able to empirically test Lim and McNelis (2016) since the thirty-four years of macroeconomic data (1980–2013) available for utilisation in this study represents a reasonably short time frame to capture the inverted U-shape income equalising effect.[10]

4 | CONCLUSION

This paper investigated the short-run and long-run equilibrium relationship and causality between income inequality-growth-financial depth in China. The Bounds test results reveal that a long-run relationship exists when income inequality is treated as the dependent variable. The long-term effects of growth and financial depth are statistically significant, and the coefficient value of the ECT shows a rapid adjustment process. Granger causality results indicate bidirectional causality between financial depth-growth and unidirectional causality between Gini-growth. Both growth and financial depth have an inequality-widening effect suggesting that China may still be on the upward slope of the Kuznets' inverted U-curve. Despite China's redistributive policies that started in the mid-2000s, the parallel increase in inequality observed in our study may be due to the policies have yet to take effect considering the country's large population.

Our results demonstrate that growth and financial depth will increase inequality in the long run. The implication suggests a trade-off of growth enhancing policies on income equality. Cognisant

[10]Studies analysing the Kuznets's theory of long-run distribution in income suggest the inclusion of historical data coverage, that is, since the start of industrialisation period or early twentieth century (Roine & Waldenström, 2015).

of the rising inequality, the government is endeavouring to move away from "short-term growth to prioritizing policies and strategies to ensure long-term prosperity for the entire nation" (Lee et al., 2014, p. 233). The earlier growth trajectory benefitted more than 500–600 million living in poverty but led to rising inequality between rural–urban areas (Lee et al., 2014). Several measures to address geographical disparity to reduce inequality such as personal tax reform, social security programmes, regional development strategy and poverty alleviation policies were in favour of reducing poverty levels in the country (Li, Wan, & Zhuang, 2014). However, whether the reduction of poverty levels will narrow inequality in the long-term warrants future studies. Given the massive growth of credit and inequalities observed in China, we call for an extended study in the context of rural–urban inequality.

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