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Autoregressive Distributed Lag Approach to the Income Inequality and Financial Liberalization Nexus: Empirical Evidence from Turkey

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ABSTRACT

This paper critically examines the validity of orthodox assumptions about the positive effects of financial liberalization on income inequality by employing the autoregressive distributed lag (ARDL) Bound test by Pesaran et al. (2001) for the yearly data of Turkey over the 1987-2016 period. The benchmark results suggest that the income distribution worsens by the implementation of more liberalization in the financial sector for a long-run equilibrium relationship. Further, the study attempts to investigate an inverted U-shaped hypothesis of financial development. The findings indicate that income inequality improves at the initial stages of financial development but worsens over time. Therefore, it rejects the income-narrowing hypothesis for the latter stages of financial development. With the advantage that ARDL approach incorporates both $I(0)$ and $I(1)$ series, the study concludes that the positive relationship between financial liberalization and income inequality is prevalent both in the short- and the long-run in control of other variables.

Keywords: Income Inequality, Financial Liberalization, Financial Development, Income Distribution, Autoregressive Distributed Lag Bound Test
JEL Classifications: F65, D31, C22

1. INTRODUCTION

The degree of financial liberalization may have possible effects on different kinds of indicators which are crucial for the overall health of countries' long-run economic performance. In that sense, any kind of interruption for a higher level of financial liberalization can damp many of the critical topics for future economic development. This pro-liberalization perspective is the common vision of orthodox perspective in the economic discipline in which the less liberalized financial systems may intensify economic downturns by way of reducing the production level, slowing down the capital accumulation, increasing income inequality and causing poverty since these and the other indicators are highly correlated to the degree of economic performance, especially on the basis of economic growth. Indeed, this topic and its correlates have been examined by a large number of studies over the past decades. Many of these studies, depending on the orthodox arguments, primarily emphasized on various sub-components and different discussions of financial liberalization. One of the most important correlates of financial liberalization that has been investigated explicitly is the extent of distributional/allocational relations. However, even in the orthodox wisdom, there is no consensus for the case of the relationship between a higher degree of financial liberalization and income distribution. According to the proponents of more liberalized financial sector, the correlation between a higher degree of financial liberalization and the level of income inequality is significant, negative and robust which also leads to higher rates of economic growth through a higher level of productivity, capital accumulation, and economic efficiency. This point is what the paper focuses on where the orthodox arguments on the basis of this negative correlation for financial liberalization-income distribution nexus, and argues that contrary to the conventional wisdom this relationship can have different dynamics and thus have a positive link among these two variables along with socio-economic and political components.

One of the traditional argument for pro-liberal policy framework is that a higher level of financial openness in capital account leads to an efficient allocation of resources and hence alleviates the distributional/allocational problems. The main reason is based on the flows of financial resources which improve and then provide economic needs for future investments. In particular, if the country has not a necessary amount of financial resources for making new investments, a higher degree of openness in the financial account can fill this gap by way of providing flows of money into the economic system, which of those are used in productive investments. It is also argued that financial liberalization is seen as a popular policy choice for removing reserve requirement (McKinnon, 1973; Shaw, 1973) which drives total savings and thereby the level of investment. This traditional model assumes that financial repression creates a high cost in production due to a high level of borrowing cost. However, rising interest rates through the elimination of reserve requirements create incentives for savers to save more and thus enhance resource allocation and credit availability by way of organizing the banks for the intermediation process. In that vein, these policy tools improve the current and future rates of economic growth and physical investment. For instance, Bumann and Lensink (2013) specify the common vision of the standard model on which the proceeds of financial liberalization are equally distributed. According to this argument, the fruits of financial liberalization are equally shared due to the fact that each economic agent has a pearl of rational wisdom who allocates the economic resources for whole life period at an optimal level. In other words, the growth of economic output through financial liberalization is equally distributed among economic agents and hence creates a positive pressure on the decrease of income inequality. Related to the distributional/allocation problems, the standard model further assumes that the marginal productivity of labor increases as the economic growth stimulates, so do their wages or salaries, across different economies where the financial liberalization is adopted and functions well. However, the link between financial liberalization and income inequality is not explicit in the standard model and hence necessitates to deal with its relations to the economic growth. It means that any positive implications in economic growth bring further improvements in capital accumulation, economic efficiency, and productive capacity and therefore leads to a lower level of income inequality. The ways that lead to a positive shift in income distribution can be reflected by way other types of policy reforms as well: (i) The abolishment of entry barriers which includes banks, non-financial and financial institutions; (ii) privatization of financial sector; (iii) the removal of restrictions on capital accounts; (iv) supervision of financial institutions and markets; and (v) the policy reform to lead further development in security markets (Abiad et al., 2010; Agénor and Montiel, 2015).

However, this implicit framework for financial liberalization-income distribution nexus in the standard model does not quite capture the whole story. Some of the studies instead argue that the reduction in the level of income inequality is done by way of conducting the interaction between financial liberalization and financial development. For instance, the liberalized financial sector leads to the removal of constraints on credits which are restricted to the poor agents for their funding of business projects since they had no collateral to avoid from going bankrupt (Galor and Zeira, 1993; Aghion and Bolton, 1997; Aghion et al., 1999). Therefore, a developed financial system as a whole, including both financial markets and financial institutions, stimulates the access of the poor agents to monetary sources which were previously impeded because of repressed financial relations. In that vein, the market economies where their financial sector is intended to be liberalized by the legal authorities should also consider providing a developed financial system. If this is the case, the distributional/allocation problems will be narrowed in the future period and the economic pie will be shared in equal portions among individuals.

Various studies have discussed the assumptions which are related to the positive effects of financial liberalization on economic ingredients but some of them challenged with focal views adherent to the standard model. For instance, traditional wisdom does not significantly notice the importance of the effects of price-setting behavior of monopolistically competitive firms on economic efficiency and thus ignore the role of informal markets in financial relations and transactions. According to the new structuralist theory, informal markets can be necessarily much more competitive than that of what the traditional view argues basically through affecting the process of financial intermediation (Taylor, 1983; van Wijnbergen, 1982). If these conditions are valid then financial liberalization policies invite reallocation of funds but mostly from the informal sector which does not affect the current investment level (Bumann and Lensink, 2013). The only way to provide the optimal level for financial liberalization is to share risks among economic agents (Bencivenga and Smith, 1992). Furthermore, more open financial systems possibly face with the market failures because of asymmetric information in which the number of financial institutions are encountered over time and hence leads to the emergence of moral hazard and adverse selection problems (Hellmann et al., 1996). While sound financial sector needs strong check and balance system in financial relations and transactions, any kind of possibility disrupting the mutual correlations between financial liberalization and financial development on the basis of information asymmetries might lead banks to refrain from lending to their customers. Alternatively, new information asymmetries might negatively affect banking behavior by way of switching some banks to accept higher risk exposure levels in order to make more profits or to encourage them towards speculative strategies in their financial decisions (Stulz, 1999; Boot, 2000; Hellmann et al., 2000). In addition, financial liberalization can exacerbate the negative conditions which lead to the economic downturns such as real and/or financial crises, mostly due to fierce competition and thereby excessive risk-taking (Arestis, 2005). According to this argument, information asymmetries should be controlled through government restrictions on economic conditions. Therefore, governments should provide fair positions for economic agents through creating different opportunities to avoid social, economic and political problems.

While financial liberalization induces some economic problems as discussed by the critical thoughts, the impact on distributional/allocational components is ambiguous even though the development channel of finance is operational. First of all, the equal distribution of income can be significantly affected by financial liberalization in the presence of the financial sector which is well-developed and perfectly functioning. If this is the case, the distributional effects of finance can be determined by other sub-factors which are dependent on the changes in the financial sector as a whole. For instance, Demirguc-Kunt and Levine (2009) argue that there two channels in which financial development can affect income distribution: (i) Extensive margin and (ii) intensive margin. On the one hand, at the extensive margin, financial services are used by the economic agents who were previously neglected from the financial system, such that it narrows the income inequality. On the other hand, at the intensive margin, the short-run dynamics capture the financial sector by which the available funds are substantially controlled under the upper-income segments of society and thus lead to higher inequality. Even though the explanations towards the margins are clear cut to show the direction of distributional phenomena, there are also other studies which document that the changes in income distribution depend on the threshold level of economic development. By considering this argument, Greenwood and Jovanovic (1990) note that the level of inequality becomes lower along with financial development if the countries pass beyond a threshold level of economic development. Kunieda et al. (2014) also indicate that financial development is only beneficial for income distribution in closed economic systems.

Despite the development of a financial sector, the strands of criticism for changing dynamics of finance are also canalized into the role of financial liberalization in distributional problems through the way of institutional differences. As one of the most important policy component of economic globalization, financial liberalization can intensify or mitigate its effect on income distribution through changing income shares accruing to capital and to labor in an aggregate economy. The heterodox thoughts principally argue that globalization-led policy usage in favor of a higher degree of financial liberalization have negative impact on the labor share of income but positively affects the capital share, due to several reasons such as changing institutions, weakening bargaining power of labor or increasing threat option to locate capital abroad (Cornia, 2005; Jayadev, 2007; Checchi and Garcia-Peñalosa, 2010). These factors ultimately worsen income inequality in concordance with a higher level of capital share. Therefore, heterodox arguments note that the traditional outcomes do not totally reflect the changing dynamics of income shares between capital and labor since they neglect the class analysis and evaluate the economic actors as separate agents from each other.

The financial liberalization-income distribution nexus has a mutual framework. While it is widely known as a correlation, the causal role of this link should be also mentioned within the scope of various regressors. As a matter of fact, income inequality across different economies covering both developed and developing markets have substantially increased over the post-1980s. Recent literature, especially the heterodox wisdom, has increasingly focused on the causes of exacerbated problems in income distribution and hence has mostly paid attention to the financial liberalization policies. According to the so-called literature, financial liberalization is expounded as having a less authority of government in the economy to a large extent and having an active role of financial markets (Abiad et al., 2008). In this context, various studies indicate that there is a positive relationship between financial liberalization and income inequality which means that higher degree of financial liberalization leads to lower inequality level in income (Agnello et al., 2012; Delis et al., 2014; Li and Yu, 2014). However, the others find that this positive link turns into negative by which financial liberalization intensifies income inequality (Jaumotte and Buitron, 2015; de Haan and Sturm; 2017; Zhang and Naceur, 2019).

In case of financial liberalization channels, we encounter a growing scale of openness in the capital account. According to Furceri and Loungani (2015. p. 4-5), there are different ways in which capital account openness can affect the level of income inequality such as risk-sharing, financial crises, foreign direct investment in the host economy, and the change in the labor share of income. Whereas the empirical findings show that there is a positive relationship among two indicators, the impact of financial liberalization on income distribution is conditioned to the level of financial depth. In such a case, a well-functioning financial system prevents further negative impacts of financial liberalization on income inequality. In addition, Bumann and Lensink (2016) scrutinize the impact of having a sufficient level of financial depth, which is higher than 25%, on income inequality and note that if this condition is prepared, higher degree of financial liberalization creates downward pressure on a growing level of income inequality. Therefore, the interaction between financial depth and financial openness should be well-developed in order to make a positive effect on income distribution.

All in all, the research question of this paper is to reveal whether there is a causal relationship between financial liberalization and income distribution. To understand the link between these two indicators, the methodology of this study benefits from the autoregressive distributed lag (ARDL) method for Turkey over the period of 1987-2016. Although the details of an ARDL framework will be discussed in the following part, it should be also noted that the main objective of this endeavor is to analyze the long- and short-run correlations for financial liberalization-income distribution nexus. To best of our knowledge, the research topic in this study with regard to the current nexus has not been profoundly studied yet for Turkey at a given time period. By the consideration of this detail, the second part will be devoted to the explanation of data sample and methodology which covers the ARDL framework. The third part will be based on the interpretation of empirical evidence which will contain both short- and long-run versions of ARDL bound test. The fourth part will be devoted to the sensitivity tests to determine whether the long-run coefficients are stable. The last part concludes.

2. DATA AND METHODOLOGY

As it was mentioned in the introduction part, the time span of this study contains the 1987-2016 period. One of the reasons for choosing this time interval depends on the data limitation in case of income inequality which is proxied by GINI coefficient developed by Solt (2019) and is available between 1987-2016 period for Turkey. The other reason is also resulted from data selection procedure, especially for financial liberalization indicator. While there are different types of methodologies in calculating the financial liberalization in the empirical literature, we follow the data set of so-called "The Trilemma Indexes" which is constructed by Aizenman (2010; 2018) and includes three policy choices: monetary independence, exchange rate stability, and financial openness. By ignoring the other two indicators, we use financial openness index as a proxy for financial liberalization[1].

Before the explanation of the methodological framework, we can also provide some detailed information about these benchmark variables. First, the GINI coefficient is obtained from the Standardized World Income Inequality Database and ranges between 0 and 100 in which higher values indicate more inequality. Further, the GINI coefficient that we use in the empirical estimations is called as disposable GINI since it is adjusted from taxes and transfers. Therefore, the GINI database is measured on a net basis in equivalized household disposable (i.e., post-tax and post-transfer) income. However, one of the critical shortcomings of GINI coefficient is its difference from the variables such as the labor share of income and return on production capital which lead us to investigate the distributional issues on the basis of social classes. However, the GINI coefficient is more restrictions in itself since it is only proper for the analysis to understand the household-based differences in the distribution of total income.

Second, financial openness index is obtained from Aizenman's (2018) database to measure the financial liberalization. This index is used to describe the extent and intensity of capital account controls and thus adopt the same index of capital account openness by Chinn and Ito (2006; 2008). The basis of the capital account openness index can be traced back to the methodological underpinnings of IMF's calculations for the total amount of foreign assets and debts as a percentage of GDP (Lane and Milesi-Ferretti, 2007). In the later period, IMF systematized its measurement on capital account openness in Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER). Chinn and Ito's (2006; 2008) index is hence constructed upon information regarding reported restrictions in the IMF's AREAER. In particular, four components are the leading ones that form the whole structure of capital account index (Aizenman et al., 2008. p. 8): (i) The presence of multiple exchange rates; (ii) restrictions on capital account transactions; (iii) restrictions on current account transactions; and (iv) restrictions on exports revenues. Since the restrictions about different economic components are used to form capital account openness index, it is, therefore, a *de jure* index. The main objective to employ a *de jure* measure of capital account openness relies upon the policy intentions of the countries[2] (Aizenman et al., 2008. p. 8). All in all, financial openness index is normalized between 0 and 1. 0 means that a country is least open to cross-border capital transactions and 1 means that a country is more open to cross-border capital transactions (Chinn and Ito, 2006; 2008).

Furthermore, the data set of this study contains additional macroeconomic variables which have been obtained from different data sources as follows: (i) Financial development proxied by private credit by deposit money banks and other financial institutions to GDP (%); (ii) square term of financial development (iii) logarithm of employment ratio (% of total population); and (iv) logarithm of human capital index. Table 1 describes the whole data set of this paper.

In the empirical framework, we use a log-linear model since there are some basic studies arguing that it provides more robust estimates than the simple-linear form (Bowers and Pierce, 1975; Ehrlich, 1975; Ehrlich, 1977; Layson, 1983; Cameron, 1994; Ehrlich, 1996). So, to analyze the financial liberalization-income distribution nexus, Eq. (1) is estimated as follows:

$$LGINI = \alpha_0 + \alpha_1 FL + \alpha_2 CV + \varepsilon_t \qquad (1)$$

where income inequality is represented by the logarithm of GINI coefficient (*LGINI*), financial liberalization (*FL*) is proxied by the financial openness index, and *CV* refers to the control variables in the regression such as private credit by deposit money banks and other financial institutions (% of GDP) (*PRV*), logarithm of total employment, including male and female, to total population ratio (%) (*LEMP*), and logarithm of human capital index (*LHC*). Table 2 shows descriptive statistics.

As we discussed in the introduction part, financial development is one of the other leading determinants of the relationship between financial liberalization and income inequality. Therefore, we also add a squared term of *PRV* to investigate its earlier and former effects on income inequality on the basis of studies done by Greenwood and Jovanovic (1990) and thus the nonlinear functional form is assumed as Wahid et al. (2012. p. 93) in Eq. (2):

$$\alpha_{11} PRV + \alpha_{12} PRV^2 \qquad (2)$$

[1] For more information about the theoretical context please see Aizenman (2008).

[2] For a short review of the limitations about *de jure* measures of financial openness please see Edwards (1999).

Table 1: Summary of the data sources

Code	Variable	Source	Period available
LGINI	Logarithm of GINI coefficient	Solt (2019)	1987-2016
FL	Financial liberalization index normalized by [0,1]	Aizenman et al. (2010); Aizenman et al. (2013)	1987-2016
PRV	Private credit by deposit money banks and other financial institutions (% of GDP)	Financial structure database	1987-2016
PRVSQ	Square term of private credit by deposit money banks and other financial institutions (% of GDP)	Financial structure database; Author's calculation	1987-2016
LEMP	Logarithm of employment ratio (% of total population)	Penn world Table 9.1; Author's calculation	1987-2016
LHC	Logarithm of human capital index	Penn world Table 9.1; Author's calculation	1987-2016

Table 2: Descriptive statistics

Summary	*LGINI*	*FL*	*PRV*	*LEMP*	*LHC*
Mean	1.6237	0.3090	0.2503	−0.5403	0.3066
Median	1.6299	0.4156	0.1541	−0.5374	0.3071
Maximum	1.6415	0.4489	0.6474	−0.4878	0.3820
Minimum	1.5999	0.1658	0.1223	−0.5733	0.2411
Std. dev.	−0.0144	0.1368	0.1692	0.0229	0.0427
Skewness	−0.5064	−0.1098	1.2380	0.5986	0.1329
Kurtosis	1.6829	1.0426	3.0578	2.8418	1.7867
Jarque-Bera stat. (P-value)	3.4507 (0.178)	4.8494 (0.088)	7.6676 (0.0216)	1.8231 (0.4019)	1.9284 (0.3813)
Sum	48.710	9.271	7.5083	−16.2093	9.1985
Sum sq. dev.	0.0061	0.5429	0.8305	0.0151	0.0528
Observations	30	30	30	30	30

The signs of the coefficients also reflect the different point of views on this nexus: (i) if α_{11}<0 and α_{12}<0, this means that income inequality narrows in accordance with a higher level of financial development over time; (ii) if α_{11}>0 and α_{12}=0, this means that income inequality widens in accordance with a higher level of financial development over time; (iii) if α_{11}>0 and α_{12}<0, this means that inverted U-shaped hypothesis is valid over time; and (iv) if α_{11}<0 and α_{12}>0, this means that U-shaped hypothesis is valid over time.

2.1. Unit-Root Testing Methods

This sub-section briefly explains the theoretical formation of unit root test developed by Ng and Perron (2001) to investigate the order of integration for series using in the model[3]. One of the most important findings of Ng and Perron (2001) is an introduction of modified AIC (MAIC) to sort out the over-specification of lag-length truncation problem (Zapata et al., 2011. p. 24). The lag length is selected through minimizing the MAIC. In addition, Wahid et al. (2012. p. 94) note that Ng-Perron unit root testing approach "…has good size and explaining power". Further, the matter of sample size is corrected in this recently developed approach as it is supported by the statement of Ng and Perron (2001. p. 1519) as follows: "the majority of tests suffer from severe size distortions when the moving-average polynomial of the first differenced series has a large negative autoregressive negative root". In that sense, the testing method is proper for the case of small sample sizes (e.g., t = 20 and 30). In fact, Ng and Perron (1995; 2001) built on detrended data which is originated in the Augmented Dickey-Fuller (ADF)-GLS test and modified Phillips-Perron (PP) test (Arltová and Fedorová, 2016. p. 51). All in all, to derive the Ng-Perron test, the ADF test is reported in Eq. (3):

$$\Delta y_t = \alpha y_{t-1} + x_t^{'}\delta + \beta_1 \Delta y_{t-1} + \beta_2 \Delta y_{t-2} + \ldots + \beta_p \Delta y_{t-p} + v_t \quad (3)$$

According to Dickey and Fuller (1979), traditional student's t-distribution does not depend on the ADF statistics. Therefore, Dickey and Fuller (1979) reproduce the critical values and derive asymptotic results, which is also further developed by MacKinnon (1996) in the presence of a larger set of simulations. However, Said and Dickey (1984) show that the asymptotic validation of ADF test needs to be the presence of a moving average (MA) component. The ADF test regression features different options for trends such as a constant, or a constant and a linear trend. Related to these options, Elliott et al. (1996) also make some modifications in the ADF test. First, according to Elliott et al. (1996), a quasi-difference of y_t defines the specific point for an alternative hypothesis against the null hypothesis testing in Eq. (4), which depends on the value of a:

$$d(y_t\ /a) = y_t \text{ if } t = 1 \text{ and } (y_t\ /a) = y_t - a y_{t-1} \text{ if } t > 1 \quad (4)$$

Second, Elliott et al. (1996) regress quasi-differenced data $d(y_t/a)$ on quasi-differenced data $d(x_t\ /\ a)$ as in Eq. (5):

$$d(y_t\ /a) = d(x_t\ /a)^{'}\delta(a) + \eta_t \quad (5)$$

Where x_t involves a constant and/or a trend and $\delta(a)$ is the OLS estimates obtained from Eq. (5). The only need in that context is to have a value for a. Elliott et al. (1996) assume that $a = \bar{a}$. On the one hand, $\bar{a}$ is equal to $1 - 7/T$ if $x_t = \{1\}$. On the other hand, $\bar{a}$ is equal to $1 - 13.5/T$ if $x_t = \{1, t\}$. y_t^d denotes the GLS detrended data and associates with the $\bar{a}$ as follows in Eq. (6):

$$y_t^d = y_t - x_t^{'}\delta(\bar{a}) \quad (6)$$

Dickey-Fuller test with GLS detrending also reach a new conclusion by substituting the GLS detrended y_t^d with the traditional y_t as follows in Eq. (7):

$$\Delta y_t^d = \alpha y_{t-1}^d + \beta_1 \Delta y_{t-1}^d + \beta_2 \Delta y_{t-2}^d + \ldots + \beta_p \Delta y_{t-p}^d + v_t \quad (7)$$

Unlike the Eq. (3), this last equation excludes the x_t since the y_t^d is detrended in this recently developed test. Following this context, we reach to the ERS point optimal test[4]. It uses the quasi-differenced regression defining in Eq. (4). The residuals in Eq. (5) is $\hat{\eta}_t(a) = d(y_t\ /a) - d(x_t\ /a)^{'}\delta(a)$. In addition, the sum-of-squared residuals function is defined as $SSR(a) = \sum \hat{\eta}_t^2(a)$. The null hypothesis for the point optimal test developed by Elliott et al. (1996) is formed as the following equation in which α=1 and the alternative depends on the hypothesis that $a = \bar{a}$. So, the optimal test statistic is calculated in Eq. (8):

$$P_T = SSR(a) - \bar{a}SSR(a)/f_0 \quad (8)$$

3 Note that the theoretical formation for Ng-Perron unit root testing procedure is obtained from E-views manual package and is also crosschecked with the other related studies.

4 ERS is the abbreviation of Elliot, Rothenberg, and Stock (1996).

Where f_0 denotes the estimator of the residual spectrum at frequency zero. All in all, the unit root testing procedure developed and modified by Ng and Perron (2001) which depends on above-stated four different unit root tests: Bhargava (1986) R_1 statistic, Phillips and Perron (1988) Z_a and Z_t statistics, and Elliott et al. (1996) P_T statistic. All these modified forms of test statistics are based upon the GLS detrended data y_t^d. First, let us define $k = \sum_{t-2}^{T} (y_{t-1}^d)^2 \ /T^2$. Then, the modified test statistics on the basis of four statistics can be written as follows in Eqs. (9), (10), (11), and (12):

$$MZ_a^d = \left(T^{-1}\left(y_T^d\right)^2 - f_0\right) / (2k) \tag{9}$$

$$MZ_t^d = MZ_\alpha \, x \, MSB \tag{10}$$

$$MSB^d = (k / f_0)^{1/2} \tag{11}$$

$$MP_T^d = \left(\bar{c}^2 k - \bar{c}T^{-1}\right)\left(y_T^d\right)^2 / f_0 \quad \text{if } x_t = \{1\} \text{ and}$$
$$MP_T^d = \left(\bar{c}^2 k + (1-\bar{c})T^{-1}\right)\left(y_T^d\right)^2 / f_0 \quad \text{if } x_t = \{1, t\} \tag{12}$$

Where $\bar{c} = -7$ if $x_t = \{1\}$ and $\bar{c} = -13.5$ if $x_t = \{1, t\}$.

2.2. ARDL Approach

This sub-section explains the ARDL approach to investigate the cointegrating links among different series. ARDL models include several features which differ from previous econometric techniques explaining the cointegration among the variables. First, the lagged values of all variables, including regressand and regressors, are included in the model. Second, there is a combination of endogenous and exogenous variables. Third, the variables should not be integrated of order two in the case of unit root testing. Fourth, the model neglects the integration of different orders for the series irrespective of whether the underlying variables are $I(0)$, $I(1)$, or a combination of both. Fifth, the empirical result of bounds tests establishes the short-run and long-run models if the variables are cointegrated. However, if there is no cointegration among the series, it specifies only the short-run model. Sixth, the ARDL framework also provides efficient estimation results for small sample data. Finally, the ARDL method derives unbiased long-run estimates. So, the traditional ARDL (p, q) model is specified as Eq. (13):

$$Y_t = \alpha_{oi} + \sum_{i=1}^{p} \delta_i Y_{t-i} + \sum_{i=0}^{q} \beta_i^{'} X_{t-i} + u_{it} \tag{13}$$

Where the regressand and regressors are allowed to have different integrated orders; β and δ are coefficients; α is a constant; p and q are optimal lag lengths[5]; u_{it} is a vector of the error terms. More formally, the ARDL (p, q_1, q_2,…, q_k) model specification can be estimated by the Eq. (14):

$$\Phi(L,p) y_t = \alpha_{oi} + \sum_{i=1}^{k} \beta_i (L, q_i) x_{it} + \varphi w_t + u_t \tag{14}$$

Where,

$$\Phi(L,p) = 1 - \Phi_1 L - \Phi_2 L^2 - \ldots - \Phi_p L^p$$
$$\beta(L,q) = 1 - \beta_1 L - \beta_2 L^2 - \ldots - \beta_q L^q$$

For $i = 1, 2, 3 \ldots k$, $u_t \sim iid(0; \delta^2)$.

L is a lag operator such that $L^0 y_t = X_t, L^1 y_t = y_{t-1}$, y_t is a regressand, x_t is the i^{th} regressor, α is a constant, w_t is the $sx1$ vector of deterministic variables such as the intercept term, trends, seasonal dummies, or exogenous variables, and p and q are optimal lag lengths.

Unlike the Johansen and Juselius (1990) cointegration method, each variable has a single long-run relationship equation. If there is one cointegrating vector, the ARDL model of cointegrating vector is transformed into the error correction model (ECM). In that vein, both short-run and long-run links among the series can be obtained for the ARDL model. However, the presence of multiple cointegrating vectors makes these statement invalid in which the ECM approach becomes the leading procedure. As the first step in the ARDL cointegration approach, the long-run coefficients should be determined. Eq. (15) estimates the long-run coefficients for y_t to a unit change in x_{it}:

$$\hat{\theta}_i = \frac{\hat{\beta}_i(1, \hat{q}_i)}{\hat{\Phi}(1, \hat{p})} = \frac{\hat{\beta}_{i0} + \hat{\beta}_{i1} + \ldots + \hat{\beta}_{iq}}{1 - \hat{\Phi}_1 - \hat{\Phi}_2 - \ldots \hat{\Phi}_p} \tag{15}$$

Where $\hat{p}$ and $\hat{q}_i, i = 1, 2, \ldots, k$ are the selected values of p and q. In a similar case, the long-run coefficients of exogenous variables are given in Eq. (16) as follows:

$$\hat{\lambda} = \frac{\hat{\varphi}(\hat{p}, \hat{q}_1, \hat{q}_2, \ldots, \hat{q}_k)}{1 - \hat{\Phi}_1 - \hat{\Phi}_2 - \ldots \hat{\Phi}_p} \tag{16}$$

Where $\hat{\varphi}(\hat{p}, \hat{q}_1, \hat{q}_2, \ldots, \hat{q}_k)$ is the ordinary least squares (OLS) estimates of φ in Eq. (14) for the standard ARDL model. As a second step, the ARDL model is also associated with ECM which can be estimated by the OLS method. Since the short-run relationship between the variables gives a comprehensive understanding of the regression equation, it is not sufficient to obtain more information for a longer time span. Therefore, this problem is resolved by the incorporation of ECM into the cointegration analysis in order to get information for long-run correlations among the series under consideration. The ECM is given in Eq. (17) as follows:

$$\Delta y_t = \Delta \alpha_{0i} - \hat{\Phi}(1, \hat{p}) EC_{t-1} + \sum_{i=1}^{k} \beta_{i0} \Delta x_{it} + \varphi \Delta w_t \tag{17}$$
$$- \sum_{j=1}^{\hat{p}-1} \alpha_j \Delta y_{t-1} - \sum_{i=1}^{k} \sum_{j=1}^{\hat{q}_{t-1}} \beta_{ij} \Delta x_{i,t-j} + u_t$$

5 The optimum lag orders for p (used for the regressand) and q (used for the regressors) may not necessarily be the same.

And the error correction (EC) term to identify the speed of adjustment parameter is obtained by the Eq. (18):

$$EC_t = y_t - \sum_{i=1}^{k} \hat{\theta}_i x_{it} - \lambda^{'} w_t \tag{18}$$

All in all, as Pesaran et al. (2001) argue that these two steps provide the estimation of long-run relationships between the variables under consideration. Briefly, the first step is to find out whether there is a long-run relationship among the series and the second step is to estimate both long- and short-run coefficients. However, the second step depends on the presence of the first step (Narayan, 2005).

Following the estimations of the ARDL model, we are also responsible for two additional issues: (i) Sensitivity analysis and (ii) stability test. First, the sensitivity analysis checks all diagnostic problems such as heteroskedasticity, autocorrelation, functional form, and normality. Second, we conduct the stability test to obtain the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of the squares of recursive residuals (CUSUMSQ). In particular, the CUSUM test investigates whether there is a structural break in the series. However, it does not provide information about the optimal break date if there is a problem in the series. Therefore, we also benefit from the CUSUMSQ test. By using this further test, we get information about the date of a structural break. The line for the residuals should be laid between the given critical lines. If this is the case, we can argue that there is no structural break in the series.

3. THE EMPIRICAL RESULTS FOR ARDL MODEL

One of the most crucial statistical advantages of the ARDL approach is that the series under consideration in models may not necessarily have the same orders of integration. However, one exception is that the series should not be stationary in *I*(2). According to Nkoro and Uko (2016), this procedure will crash in the presence of stochastic trend of *I*(2). Therefore, the ARDL cointegration approach is favorable for the series which are integrated of different orders such as *I*(0), *I*(1) or mixture of the both. In that vein, this technique is robust only when the long-run relationship among the series is unique with small sample size. Furthermore, Ouattara (2007) notes that the computed F-statistics provided by Pesaran et al. (2001) does not being statistically reliable when the variables become stationary in *I*(2). The bound testing method is conducted on the assumption that the series should be stationary with the integrated of order, *I*(0), *I*(1) or mixture of the both. Therefore, we strictly need to detect the stationary characteristics of the variables in the ARDL technique.

While there are different kinds of unit root tests to find out the integration of orders for variables, some of them are relevant to the large sample data sets such as ADF (Dickey and Fuller, 1979) and PP (Phillips and Perron, 1988) tests. In such cases, the variables are likely to have size distortions in which the confidence limits may not be credible. However, there are also counter-arguments which basically state that sample size and time span do not have a considerable effect on the significance of unit root testing methods. For instance, Pierse and Snell (1995. p. 336) argue that "any test that is asymptotically independent of nuisance parameters under both H_0 and H_A has a limiting distribution under both H_0 and H_A that is independent of m[6]". This actually means that, asymptotically, selective sampling has no considerable results with regards to size distortion (poor size), or power properties, for ADF and PP tests. But the recent literature on the basis of ADF and PP unit root tests also indicates that each of them suffer from severe finite sample power and size problems. First, DeJong et al. (1992) note that the ADF and PP tests have low power properties against the alternative hypothesis. Second, Schwert (1989) states that if the series have a large negative MA root, both of these tests may potentially be met with severe size distortion by incorrectly over-rejecting the null hypothesis. Therefore, Table 3 presents the unit root estimations for KPSS, ERS and Ng-Perron tests to keep away to such discussions having in the recent literature. In other words, these latter tests are utilized to cope with these issues about the size distortions, power properties and also the order of integration of series in the models.

Unit root estimations in Table 3 reveals the fact that the logarithm of GINI coefficient(*LGINI*), financial liberalization index (*FL*), logarithm of employment ratio (% of total population) (*LEMP*) and logarithm of human capital index (*LHC*) are stationary at *I*(0) in all unit root tests while private capital by deposit money banks to GDP (%) (*PRV*) as a proxy for financial development is integrated of order one, i.e., *I*(1), in KPSS test but *I*(0) in ERS and Ng-Perron tests. Although the series are largely integrated of order zero, i.e., *I*(0), this still allows us to estimate the ARDL method with cointegrating bounds in case of Turkey over the 1987-2016 period.

6 *m* is the number of periods that are became together on a flow variable.

Table 3: Results of the unit root tests

Variables	KPSS	ERS	Ng-Perron			
			MZ_a	MZ_t	MSB	MPT
LGINI	0.1649 *I* (0)	3.4731 *I* (0)	−30.6245	−3.9115	0.1277	2.9840
FL	0.1079 *I* (1)	4.6570 *I* (0)	−18.7685	−3.0561	0.1628	4.8982
PRV	0.2100 *I* (1)	2.4547 *I* (0)	−46.2290	−4.7317	0.1023	2.3492
LEMP	0.0026 *I* (0)	0.0355 *I* (0)	−3104.39	−39.387	0.0126	0.0369
LHC	0.1209 *I* (0)	6.6947 *I* (0)	−15.7304	−2.6895	0.1709	6.4534

KPSS unit root testing uses Bartlett kernel spectral estimation method for *LGINI* and *LHC*. The other variables, i.e., *FL*, *PRV*, and *LEMP*, use AR spectral-GLS detrended estimation method to determine whether the series are stationary. In addition, the lag length is manually specified on the basis of the VAR lag selection procedure. For more information for the critical values please see Table 1 in Kwiatkowski et al. (1992) for KPSS, Table 1 in Elliott et al. (1996) for ERS, Table 1 in Ng and Perron (2001) for Ng-Perron unit root testing methods. VAR: Vector autoregression

In addition to unit root estimations, Table 4 shows the lag structure for the unconditional ARDL approach. The lag order selection criteria are basically obtained through the vector autoregression (VAR) method in which both of them are determined in the lag length of three[7]. Since the sample data set does not exceed 30, we restrict the selection of lag length as of three.

In the recent literature related to co-integration testing, there are different methods to determine whether the series move together in the long- and/or short-run. However, these methods have various assumptions and thus crucial to select the correct technique in specifying the co-integrating relations among the series. Since the series are integrated of different orders and the time span is limited, we lead to select the bound testing approach developed by Pesaran et al. (2001). In the presence of bound testing approach and the co-integration among the series, the long- and short-run analyses are done by the ARDL technique. In other words, the co-existing of mixture in the integration of orders may cause the cointegration results to be false if the conventional cointegration tests are used. Therefore, the ARDL model and thus the bound test solves not only the problem of series having different orders but also tackles with small-sample data set problem to accurately examine the long-run equilibrium relationship among variables. By considering these issues, Eq. (19) indicates the model (unrestricted constant and no trend) which we follow to estimate the ARDL long-run form and bounds test:

$$\Delta y_t = c_0 + \alpha^{'}\Delta X_t + \sum_{j=1}^{p-1}\beta_j^{'}\Delta Z_{t-j} + \pi_{yy} y_{t-1} + \pi_{yX.X} X_{t-1} + \varepsilon_t \quad (19)$$

This study aims to examine whether implementing a higher degree of financial liberalization has a moderating effect on income inequality as the neoclassical school artily argues and to analyze whether there exist long and short-run equilibrium links between the income inequality, financial liberalization, and other control variables. The standard model in estimating these conditions as represented in Eq. (1) is the basis of our analysis, which can be rewritten as Eq. (20):

$$LGINI_t = \beta_0 + \beta_1 FL + \beta_2 PRV + \beta_3 LEMP + \beta_4 LHC + \varepsilon_t \quad (20)$$

To construct the ARDL model, we integrate the Eq. (20) into the Eq. (19), presenting unrestricted constant and no trend, to estimate the short- and long-run ARDL model formulating in Eq. (21) as follows:

$$\begin{aligned}\Delta LGINI_t &= c_0 + \mu_{LGINI} LGINI_{t-1} + \mu_{FL} FL_{t-1} + \\ &\mu_{PRV} PRV_{t-1} + \mu_{LEMP} LEMP_{t-1} + \mu_{LHCL} LHC_{t-1} + \\ &\sum_{i=1}^{p-1}\tau_{LGINI,i}\Delta LGINI_{t-i} + \sum_{i=0}^{p-1}\tau_{FL,i}\Delta FL_{t-i} + \sum_{i=0}^{p-1}\tau_{PRV,i}\Delta PRV_{t-i} + \\ &\sum_{i=0}^{p-1}\tau_{LEMP,i}\Delta LEMP_{t-i} + \sum_{i=0}^{p-1}\tau_{LHC,i}\Delta LHC_{t-i} + \varepsilon_t\end{aligned} \quad (21)$$

The null hypothesis is tested by Wald test and states that there is no cointegration among the variables under consideration, $H_0 : \mu_{LGINI} = \mu_{FL} = \mu_{PRV} = \mu_{LEMP} = \mu_{LHC} = 0$ and the alternative hypothesis denotes that the series are cointegrated, $H_1 : \mu_{LGINI} \neq \mu_{FL} \neq \mu_{PRV} \neq \mu_{LEMP} \neq \mu_{LHC} \neq 0$. Two sets of asymptotic critical values are used in the estimation, which are provided by Pesaran and Pesaran (1997), due to the fact that the null hypothesis is not coherent with the asymptotic distribution of the F-statistic. While the first set of critical values indicate that the variables are stationary in $I(0)$, they are stationary in $I(1)$ under the second set of critical values. The computed F-statistic defines that if it is greater than the upper bound critical value, the null hypothesis is rejected. The former situation specifies that there is a steady-state equilibrium relationship between the variables. However, if this is not the case, the steady-state equilibrium relationship among the variables becomes invalid. Further, if the computed F-statistic ranges between lower and upper bounds, $I(0)$ and $I(1)$, respectively, one should use EC term to establish cointegration among the variables (Kremers et al., 1992; Banerjee et al., 1998). The following step is to compute the long-run marginal effects of regressors on regressand. So, the short-run version of ARDL model is estimated.

According to Table 5, the calculated F-statistic is 10.4017, which is far above the insignificance of upper bound $I(1)$ and hence it is highly significant at 1% level of significance. This means that the alternative hypothesis of suggesting a long-run equilibrium relationship between income inequality and financial liberalization in control of other variables is accepted. In other words, the F-bounds statistic results indicate that it is possible to determine the long-run effects of financial liberalization on income inequality in the presence of other macroeconomic and structural variables. This seems to imply that income inequality is particularly affected by the changes in the degree of financial liberalization in Turkey, along with the other determinants. Given the above acceptance through the existence of a long-run co-movement among the variables, we can apply bound test of the ARDL approach, which is well-designed for the case of different order of integration (Pesaran and Shin, 1999; Pesaran et al., 2001).

Numerous empirical studies, mostly originated by the neoclassical school of thought, have argued that financial liberalization has

7 By following the traditional arguments on the lag length selection, as the sample size of this study is 30 we do not exceed this given lag number of three.

Table 4: VAR lag order selection criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	314.2764		7.73e-17	−22.90937	−22.66940	−22.83801
1	518.2598	317.2966	1.40e-22	−36.16688	−34.72706	−35.73874
2	554.4999	42.95953	7.55e-23	−37.00000	−34.36033	−36.21508
3	629.1720	60.84386*	3.61e-24*	−40.67940*	−36.83989*	−39.53771*

VAR: Vector autoregression. *indicates lag order selected by the criterion. LR: Sequential modified LR test statistic, FPE: Final prediction error, AIC: Akaike information criterion, SC: Schwarz information criterion, HQ: Hannan-Quinn information criterion a positive effect on income distribution both at personal and aggregate levels. However, they neglect any kind of interruption which may take its source from specific characteristics of the countries around the world economy. Besides, they substantially ignore the contradicting economic values which are started in labor markets and are determined at the end of the production process.

On the basis of the neoclassical structure for income inequality-financial liberalization nexus, the application of traditional ARDL approach to test the co-integrating relationship for this nexus provides conflicting results against the conventional wisdom. The long-run bounds test results are summarized in Table 6.

The long-run coefficients are presented in Table 7, which is determined according to the ARDL (3, 3, 3, 3, 3) model. In case of Turkey, the coefficient of financial liberalization suggests that high degree of financial liberalization (more open to the financial capital from abroad) increases income inequality, which contradicts with the mainstream arguments theoretically stated above. There is the same direction between income inequality and financial liberalization. This means that the level of income inequality increases by two units for each one more unit increase in the openness degree of finance to the foreign capital (i.e., a higher degree of financial liberalization). Furthermore, there is a negative and significant relationship between income inequality and financial development (proxied by private credit by deposit money banks and other financial institutions to GDP), implying that more developed financial system distributes income more equally among individuals. The coefficient of financial development shows that a 1% increase in financial development level reduces income inequality by 0.24%. Moreover, the logarithm of employment level and the logarithm of human capital index are positively associated with income inequality, which means that a higher level of these indicators will worsen the income distribution. These empirical results are very interesting since higher employment level or higher level of educational background have theoretically a downward pressure on the level of income inequality.

However, two factors may change the whole story within the frame of the income distribution. First, a higher level of employment in total population may not reduce income inequality if the workers are employed in low-paid sectors where their bargaining powers are limited. Second, the level of income inequality may still increase even in the case of having a high degree of human capital in the country if workers equipped with relatively a high education level are treated as unskilled workers, which is very common in countries where the policy structure of wages and/or salaries are not well-developed compared to the others.

As it can be seen that the long-run link between income inequality and financial liberalization is approved in Table 7 in case of Turkey. Therefore, we can now depict the short-run movements of income inequality caused by the changes in financial liberalization which is reported in Table 8. The empirical tests of short-run coefficient estimates are obtained by the EC form of ARDL (3, 3,

Table 5: ARDL bound testing for cointegration analysis

Computed F-statistic: 10.4017 (selected model: ARDL 3, 3, 3, 3, 3)
Critical bounds value at 1% (Lower: 4.77 and Upper: 6.67)
Unrestricted constant and no trend in the model
Pesaran et al. (2001. p. 301), Table 1

ARDL: Autoregressive distributed lag

Table 6: Benchmark long-run estimates of the ARDL (3, 3, 3, 3, 3) model

Variables	Lag order			
	0	1	2	3
LGINI		0.8476 (4.22)***	0.0101 (0.04)	−0.6033 (−2.68)**
FL	0.0139 (3.21)**	0.0091 (1.78)	0.0024 (1.10)	−0.0057 (−2.38)**
PRV	−0.0839 (−3.53)***	0.0099 (0.81)	−0.0249 (−1.80)	−0.0824 (−3.54)***
LEMP	0.1515 (3.30)**	0.1037 (2.88)**	0.1758 (4.61)***	0.1681 (5.31)***
LHC	1.2105 (3.28)**	−1.0986 (−2.84)**	−1.6135 (−2.84)**	1.7881 (3.60)***
$LGINI_{t-1}$		−0.7455 (−5.07)***		
LFL_{t-1}		0.0198 (3.06)**		
$LPRV_{t-1}$		−0.1814 (−3.88)***		
$LEMP_{t-1}$		0.5993 (4.38)***		
LHC_{t-1}		0.2865 (3.10)**		
Constant	1.4878 (5.09)***			
R-squared=0.9996				
Adj-R-squared=0.9987				
F-statistic=1129.036				
Prob (F-statistic)=0.0000				
Durbin-Watson statistic=2.938				
AIC=−12.129				
SC=−11.169				
Sensitivity analysis: Normality J-B Value: F=1.2690 (0.53)				
Serial correlation LM, F=2.3514 (0.21)				
Heteroskedasticity test, F=0.4879 (0.90)				
ARCH Test, F=0.7561 (0.53)				
Ramsey reset test, F=1.5439 (0.26)				

t-values are in the parentheses. *, **, and *** denote the significance levels at 10%, 5%, and 1%, respectively. The appropriate lag lengths are determined by the VAR lag order selection criteria. The lag orders are selected by the AIC and the maximum lags for regressand and the regressors are automatically determined at three. ARDL: Autoregressive distributed lag, VAR: Vector autoregression

3, 3, 3) model. The Eq. (22) represents the unrestricted ECM to estimate the short-run relationship for income inequality-financial liberalization nexus.

$$\Delta LGINI = \alpha_0 + \sum_{i=0}^{m}\beta_{1i}\Delta LGINI_{t-i} + \sum_{i=0}^{m}\beta_{2i}\Delta FL_{t-i} + \sum_{i=0}^{m}\beta_{3i}\Delta PRV_{t-i} + \sum_{i=0}^{m}\beta_{4i}\Delta LEMP_{t-i} + \sum_{i=0}^{m}\beta_{5i}\Delta LHC_{t-i} + \beta_6 ECM_{t-1} + \varepsilon_t \quad (22)$$

The EC_{t-1} coefficient depicts the speed of adjustment term meaning that "…how quickly/slowly variables return to long-run equilibrium from short-run changes in income inequality…" (Wahid et al. 2012. p. 101). This EC_{t-1} term should mostly be lie between 0 and −1. The highly significant EC term indicates that there is a stable long-run relationship between the variables.

The use of the lagged of EC term represents the speed of adjustment of income inequality which depicts the movement level to bring back the equilibrium in the short-run model. The estimation results show that EC_{t-1} is highly significant at 1% level of significance. In other words, EC_{t-1} implies that the deviation from the long-run equilibrium level of income inequality of the current period is corrected by 74.6% in the following period to bring back the equilibrium. The lag length of the short-run model is selected relying on the AIC. The results are presented in Table 8.

The short-run estimations also support the initial findings obtained by the long-run regression that financial liberalization worsens income distribution in the short-run as well. It can be argued that in the short-run, an increase in the degree of financial liberalization by one unit leads to an increase in the level of income inequality by 1.4 unit for each year, which is statistically significant at 1%. In addition, it should be noted that this result is also valid for further orders and the effect is still statistically significant. Moreover, the coefficients of differenced variables have the same signs as in the long-run equilibrium estimations and their effects are still statistically significant at different levels of significance.

To develop these benchmark estimations, we also discuss the non-linearity in the relationship of financial development with income inequality. On the one hand, Banerjee and Newman (1993) and Galor and Zeira (1993) argue that making financial sector more developed leads to lower inequality levels in case of income distribution among different social categories. On the other hand, Greenwood and Jovanovic (1990) state that there is an inverted U-shaped relationship between financial development and income inequality in which the initial stages of financial development process are witnessed that the level of inequality is high at the peak but reduces at the latter stages of financial development. This further empirical investigation by including the square term of financial development also provides the robustness checks of the benchmark results.

According to Table 9, the computed F-statistic is 65.5094, which is higher than the upper bound $I(1)$ implying that even in case of an extended ARDL model with the square term of financial development, the long-run equilibrium relationship between income inequality and financial liberalization still prevails at the highest level of significance. Therefore, the existence of this long-run co-movement among the variables leads us to apply the bound test of ARDL approach. The results are presented in Table 10 and the long-run coefficients of an extended ARDL (1, 2, 3, 1, 3, 3) model are presented in Table 11.

The overall results suggest that there is still a positive relationship between income inequality and financial liberalization in Turkey over the 1987-2016 period. The empirical evidence in Table 10

Table 7: Long-run coefficients for ARDL (3, 3, 3, 3, 3) model

Variables	Coefficient	Std. errors	t-statistic (P-value)
FL	0.0266	0.0065	4.0941 [0.0046]***
PRV	−0.2433	0.0282	−8.6300 [0.0001]***
LEMP	0.8039	0.0808	9.9445 [0.0000]***
LHC	0.3842	0.0745	5.1554 [0.0013]***

*, **, and *** denote the significance levels at 10%, 5%, and 1%, respectively. ARDL: Autoregressive distributed lag

Table 8: The error correction form on the basis of ARDL (3, 3, 3, 3, 3) model

Variable	Coefficient	Std. error	t-statistic	P-value
Δ*LGINI* (−1)	0.5931	0.0795	7.46	0.0001***
Δ*LGINI* (−2)	0.6033	0.1233	4.93	0.0017***
Δ*FL*	0.0140	0.0017	8.04	0.0001***
Δ*FL* (−1)	0.0032	0.0011	2.83	0.0252**
Δ*FL* (−2)	0.0057	0.0013	4.31	0.0035***
Δ*PRV*	−0.0840	0.0119	−7.05	0.0002***
Δ*PRV* (−1)	0.1074	0.0113	9.50	0.0000***
Δ*PRV* (−2)	0.0824	0.0086	9.63	0.0000***
Δ*LEMP*	0.1515	0.0209	7.26	0.0002***
Δ*LEMP* (−1)	−0.3440	0.0346	−9.95	0.0000***
Δ*LEMP* (−2)	−0.1681	0.0224	−7.50	0.0001***
Δ*LHC*	1.2106	0.1763	6.87	0.0002***
Δ*LHC* (−1)	−0.1745	0.1390	−1.26	0.2496
Δ*LHC* (−2)	−1.7881	0.2482	−7.20	0.0002***
EC_{t-1}	−0.7455	0.0825	−9.04	0.0000***
Constant	1.4878	0.1649	9.02	0.0000***
R-squared=0.9829				
Adj-R-squared=0.9597				
F-statistic=42.3590				
Prob (F-statistic)=0.0000				
Durbin-Watson statistic=2.938				
AIC=−12.425				
SC=−12.197				

*, **, and *** denote the significance levels at 10%, 5%, and 1%, respectively. The appropriate lag lengths are determined by the VAR lag order selection criteria. The lag orders are selected by the AIC and the maximum lags for regressand and the regressors are automatically determined at three. ARDL: Autoregressive distributed lag, VAR: Vector autoregression

Table 9: Extended ARDL bound testing for cointegration analysis

Computed F-statistic: 65.5094 (selected model: ARDL 1, 2, 3, 1, 3, 3)
Critical bounds value at 1% (Lower: 4.53 and Upper: 6.37)
Unrestricted constant and no trend in the model
Pesaran et al. (2001. p. 301), Table 1

ARDL: Autoregressive distributed lag

Table 10: The long-run estimates of an extended ARDL (1, 2, 3, 1, 3, 3) model

Variables	Lag order			
	0	1	2	3
LGINI		0.5529 (6.77)***		
FL	0.0128 (4.52)***	0.0079 (2.51)**	0.0053 (3.04)**	
PRV	−0.0368 (−2.40)**	−0.0547 (−4.57)***	0.0023 (0.28)	−0.0521 (−4.64)***
PRV^2	−0.0245 (−1.31)	0.0956 (5.63)***		
LEMP	0.0496 (1.94)*	0.0293 (1.53)	0.1016 (4.88)***	0.1032 (4.73)***
LHC	0.7420 (4.19)***	−0.5652 (−2.63)**	−0.7413 (−2.25)*	0.6619 (2.34)**
$LGINI_{t-1}$		−0.4471 (−5.47)***		
LFL_{t-1}		0.0259 (5.18)***		
$LPRV_{t-1}$		−0.1414 (−5.39)***		
$LPRV^2_{t-1}$		0.0711 (3.88)***		
$LEMP_{t-1}$		0.2836 (3.76)***		
LHC_{t-1}		0.0973 (2.16)*		
Constant	0.8682 (5.29)***			
R-squared=0.9998				
Adj-R-squared=0.9994				
F-statistic=2636.992				
Prob (F-statistic)=0.0000				
Durbin-Watson statistic=2.958				
AIC=−12.864				
SC=−11.951				
Sensitivity analysis: Normality J-B value: F=0.0209 (0.98)				
Serial correlation LM, F=1.3284 (0.36)				
Heteroskedasticity test, F=1.5548 (0.27)				
ARCH test, F=3.4009 (0.04)				
Ramsey reset test, F=0.0633 (0.81)				

t-values are in the parentheses. *, **, and *** denote the significance levels at 10%, 5%, and 1%, respectively. The appropriate lag lengths are determined by the VAR lag order selection criteria. The lag orders are selected by the AIC and the maximum lags for regressand and the regressors are automatically determined at three. ARDL: Autoregressive distributed lag, VAR: Vector autoregression

shows that there is a significant positive response of financial liberalization to income inequality. Such a response of financial liberalization also maintains in other periods. This finding thus substantially backs up the benchmark results related to the income inequality-financial liberalization nexus. Moreover, the logarithm of employment ratio (% of total population) and the logarithm of human capital index are positively associated with the income inequality in the short-run and their effects are statistically significant at 1% and 5% levels of significance, respectively. However, the striking result can be deduced from the coefficient of square term of financial development, which rejects the inverted U-shaped hypothesis proposed by Greenwood and Jovanovic (1990). According to the empirical evidences in Table 11, the level of income inequality reduces at the initial stages of financial development proxied by private credit by deposit money banks and other financial institutions (% of GDP) remains low and goes up at the latter stages of financial development[8] (For similar evidences please see Tan and Law, 2012; Jauch and Watzka, 2016). In other words, the empirical evidence shows that the linear correlation between income inequality and financial development is a U-shaped and the coefficient of the square term of financial development is statistically significant at 5% level of significance.

Finally, Table 12 presents the short-run evidence for extended ARDL (1, 2, 3, 1, 3, 3) model suggesting that the deviation from long-run equilibrium level of income inequality of the current period is corrected by 44.7% in the next period. While the coefficient of lagged of EC term reduces to the −0.447 compared to that of benchmark point (i.e., −0.745), it is still highly significant. Therefore, the speed of adjustments in the estimated models, covering both the benchmark estimations and robustness checks, are high and have the expected significant and negative sign. The Eq. (23) represents the unrestricted ECM to estimate the short-run relationship for income inequality-financial liberalization nexus.

Table 11: Long-run coefficients for and extended ARDL (1, 2, 3, 1, 2, 3) model

Variables	Coefficient	Std. Errors	t-statistic (P-value)
FL	0.0581	0.0129	4.49 [0.002]***
PRV	−0.3162	0.0522	−6.05 [0.0003]***
PRV^2	0.1591	0.0557	2.85 [0.0213]**
LEMP	0.6344	0.0852	7.44 [0.0001]***
LHC	0.2177	0.0758	2.87 [0.0208]***

*, **, and *** denote the significance levels at 10%, 5%, and 1%, respectively. ARDL: Autoregressive distributed lag

$$\Delta LGINI = \alpha_0 + \sum_{i=0}^{m}\beta_{1i}\Delta LGINI_{t-i} + \sum_{i=0}^{m}\beta_{2i}\Delta FL_{t-i} + \sum_{i=0}^{m}\beta_{3i}\Delta PRV_{t-i} + \sum_{i=0}^{m}\beta_{4i}\Delta PRVSQ_{t-i} + \sum_{i=0}^{m}\beta_{5i}\Delta LEMP_{t-i} + \sum_{i=0}^{m}\beta_{6i}\Delta LHC_{t-i} + \beta_7 ECM_{t-1} + \varepsilon_t \quad (23)$$

[8] For detailed information about the theoretical background and discussions please also see Baiardi and Morana (2018).

4. RESULTS FOR THE STABILITY TESTS

This section applies the CUSUM and CUSUMSQ to test the stability of the ARDL (3, 3, 3, 3, 3) and the ARDL (1, 2, 3, 1, 3, 3) models used. In other words, by practicing the CUSUM and CUSUMSQ, we examine the stability of the long-run parameters together with the short-run dynamics for the equations, which is developed by Borensztein et al. (1998). The stability tests can be applied to the residuals obtained in EC models (Pesaran and Shin, 1999). CUSUM and CUSUMSQ tests do not necessarily need to the determination of breakpoints of the series as in the Chow test. Therefore, it does not need primarily to specify the breakpoints at the beginning of the regressions. The main reason for why the stability tests are applied depends on the fact that the Turkish economy as a whole may confront with structural breaks at a given period and therefore the CUSUM and CUSUMSQ tests are applied to determine the coherence of short- and long-run coefficients in case of stability among the variables, proposed by Brown et al. (1975). Figures 1 and 2 plot the CUSUM and CUSUMSQ statistics for Eq. (22) and Figures 3 and 4 graphically represent the CUSUM and CUSUMSQ for Eq. (23). On the one hand, it can be seen from Figures 1 and 2 that the plot of CUSUM and CUSUMSQ statistics remain within the critical bounds of the %5 significance level that approve the long-run relationship among the variables and hence

Table 12: The error correction form on the basis of an extended ARDL (1, 2, 3, 1, 2, 3) model

Variable	Coefficient	Std. error	t-statistic	P-value
ΔFL	0.0128	0.0011	11.87	0.0000***
$\Delta FL\,(-1)$	−0.0053	0.0007	−7.05	0.0001***
ΔPRV	−0.0368	0.0060	−6.09	0.0003***
$\Delta PRV\,(-1)$	0.0499	0.0045	11.05	0.0000***
$\Delta PRV\,(-2)$	0.0521	0.0041	12.86	0.0000***
ΔPRV^2	−0.0245	0.0097	−2.51	0.0364**
$\Delta LEMP$	0.0496	0.0063	7.90	0.0000***
$\Delta LEMP\,(-1)$	−0.2048	0.0112	−18.34	0.0000***
$\Delta LEMP\,(-2)$	−0.1031	0.0106	−9.71	0.0000***
ΔLHC	0.7420	0.0986	7.52	0.0001***
$\Delta LHC\,(-1)$	0.0794	0.0882	0.90	0.3942
$\Delta LHC\,(-2)$	−0.6619	0.0709	−9.33	0.0000***
EC_{t-1}	−0.4471	0.0178	−25.08	0.0000***
Constant	0.8682	0.0345	25.19	0.0000***
R-squared=0.9912				
Adj-R-squared=0.9824				
F-statistic=111.75				
Prob (F-statistic)=0.0000				
Durbin-Watson statistic=2.958				
AIC=−13.234				
SC=−12.562				

Note: *, **, and *** denote the significance levels at 10%, 5%, and 1%, respectively. The appropriate lag lengths are determined by the VAR lag order selection criteria. The lag orders are selected by the AIC and the maximum lags for regressand and the regressors are automatically determined at three. ARDL: Autoregressive distributed lag, VAR: Vector autoregression

Figure 1: Plot of the cumulative sum of recursive residuals (Autoregressive distributed lag [3, 3, 3, 3, 3] model)

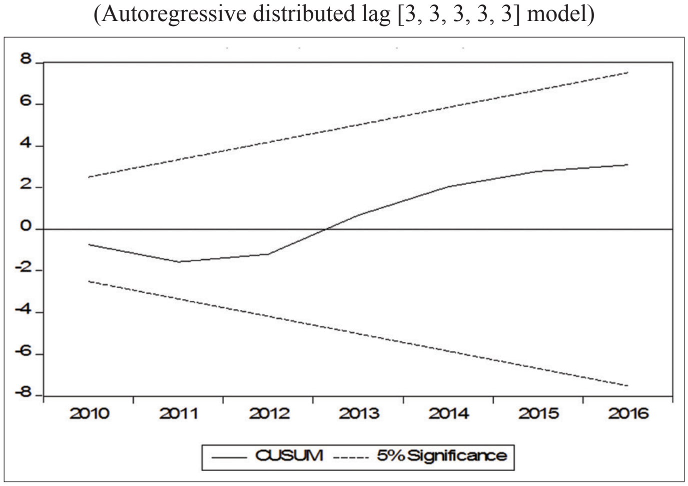


Figure 2: Plot of the cumulative sum of squares of recursive residuals (Autoregressive distributed lag [3, 3, 3, 3, 3] model)

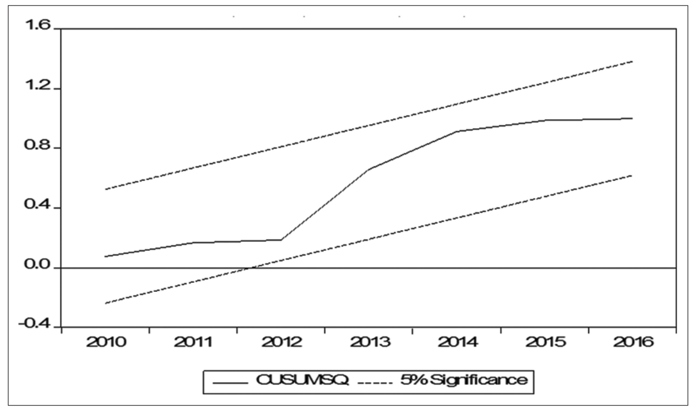


Figure 3: Plot of the cumulative sum of recursive residuals (Autoregressive distributed lag [1, 2, 3, 1, 3, 3] model)

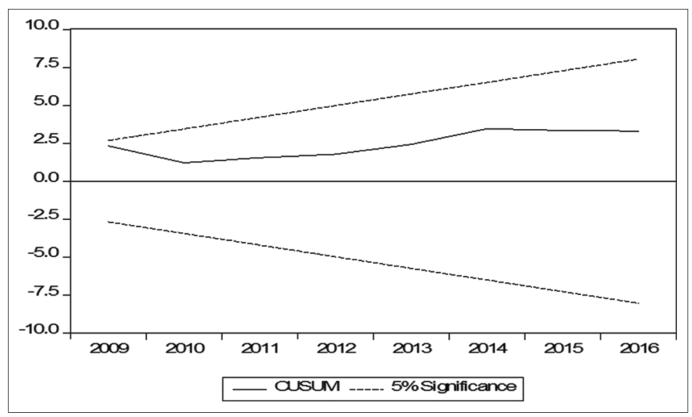
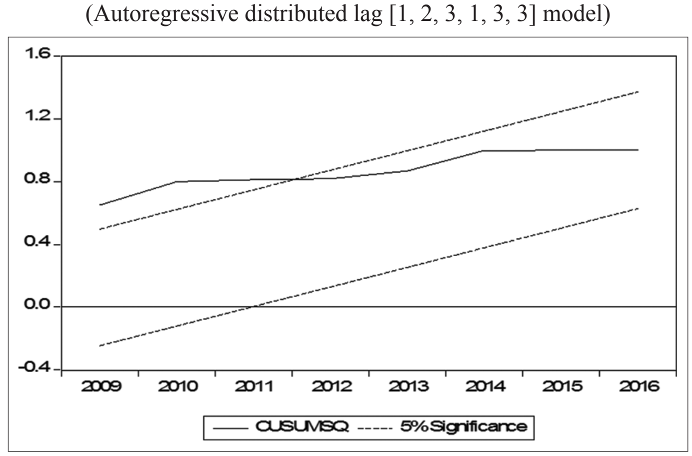


Figure 4: Plot of the cumulative sum of squares of recursive residuals (Autoregressive distributed lag [1, 2, 3, 1, 3, 3] model)



confirm the stability of coefficients. On the other hand, the same results also hold for Figure 3 but not for Figure 4, which represent the CUSUM and CUSUMSQ, respectively, for an extended ARDL model by including the square term of financial development. All in all, the estimated parameters are substantially stable within the given period in the case of Turkish economy.

5. CONCLUDING REMARKS

This paper critically re-investigates the orthodox arguments about the positive effects of implementing a higher degree of financial liberalization on income inequality by employing an ARDL bound test by Pesaran et al. (2001). The empirical evidence is indicated that there is a long-run equilibrium relationship between income inequality and financial liberalization in control of other variables such as financial development, logarithm of employment ratio (% of total population), and logarithm of human capital index. However, contrary to the mainstream outputs, our results show that financial liberalization exacerbates the level of income inequality in Turkey over the 1987-2016 period. These contrary results are also prevailing both in the short- and the long-run. Moreover, this study re-examines the inverted U-shaped relationship between financial development and income inequality, which is supported by Greenwood and Jovanovic (1990) and finds that this relationship is not accepted. In other words, contrary to the inverted U-shaped hypothesis, the empirical results suggest that there is a U-shaped relationship between financial development and income inequality, showing that while at the initial stages of financial development, income inequality decreases but then it increases at the latter stages of financial development, in which it is found to be significant and robust in a linear form.

The empirical evidence primarily shows that the only indicator to dampen the level of inequality is the proxy of financial development for the case of ARDL (3, 3, 3, 3, 3) model. While this coincides with the orthodox wisdom, the square term of financial development in the extended ARDL (1, 2, 3, 1, 3, 3) model supports the fact that this inequality-reducing power of financial development at the initial stages turns into positive, implying that the further periods of financial development witness a worsening distribution of income and the effect is statistically significant. Furthermore, the coefficients of the other regressors contradict with the orthodox assumptions. First, the positive sign of financial liberalization coefficient implies that there is a positive relationship between income inequality and financial liberalization due to several reasons such as the conditions of labor markets, political conjecture, the bargaining position of workers and/or macroeconomic dynamics. Second, contrary to the mainstream arguments, the level of income inequality is positively correlated with the employment ratio (% of total population) and human capital index. While the reasons may change over time depending on the socio-political and economic conditions in Turkey, the demand and supply sides of labor markets in relation to the capitalist production process should be well-examined. All these empirical findings statistically show that the orthodox assumptions are mostly contradictory in the case of Turkey over the 1987-2016 period. Therefore, the policy recommendations provided by the public authorities for implementing a higher degree of financial liberalization in an aggregate economy should be evaluated in caution and thus if this is not the case then its future effects upon the total economy should be well-determined.

REFERENCES

Abiad, A., Oomes, N., Ueda, K. (2008), The quality effect: Does financial liberalization improve the allocation of capital? *Journal of Development Economics*, 87(2), 270-282.

Abiad, A., Detragiache, E., Tressel, T. (2010), A New Database of Financial Reforms. *IMF Staff Papers*, Palgrave Macmillan, 57(2), 281-302, June.

Agénor, P.-R., Montiel, P.J. (2015), *Development Macroeconomics*. Princeton: Princeton University Press.

Aghion, P., Bolton, P. (1997), A Theory of Trickle-Down Growth and Development. *The Review of Economic Studies*, 64(2), 151-172.

Aghion, P., Caroli, E., Garcia-Peñalosa, C. (1999), Inequality and Economic Growth: The Perspective of the New Growth Theories. *Journal of Economic Literature*, 37(4), 1615-1660.

Agnello, L., Mallick, S.K., Sousa, R.M. (2012), Financial reforms and income inequality. *Economics Letters*, 116(3), 583-587.

Aizenman, J. (2010), *The Impossible Trinity (aka The Policy Trilemma)*. Available at: http://web.pdx.edu/~ito/Tril_Aizenman_Dic.pdf, May.

Aizenman, J. (2018), *A Modern Reincarnation of Mundell-Fleming's Trilemma* (March 6, 2018). Available at: https://ssrn.com/abstract=3135405.

Aizenman, J., Chinn, M.D, Ito, H. (2008), Assessing the Emerging Global Financial Architecture: Measuring the Trilemma's Configurations over Time. *NBER Working Papers 14533*, National Bureau of Economic Research, Inc.

Aizenman, J., Chinn, M.D, Ito, H. (2010), The Emerging Global Financial Architecture: Tracing and Evaluating the New Patterns of the Trilemma's Configurations. *Journal of International Money and Finance*, 29(4), 615-641.

Aizenman, J., Chinn, M.D, Ito, H. (2013), The 'Impossible Trinity' Hypothesis in an Era of Global Imbalances: Measurement and Testing. *Review of International Economics*, 21(3), 447-458.

Arestis, P. (2005), Financial Liberalisation and the Relationship Between Finance and Growth. *CEEP Working Paper No. 05/05*, June.

Arltová, M., Fedorová, D. (2016), Selection of Unit Root Test on the Basis of Length of the Time Series and Value of AR(1) Parameter. *Statistika*, 96(3), 47-64.

Baiardi, D., Morana, C. (2018), Financial development and income distribution inequality in the euro area. *Economic Modelling*, 70, 40-55.

Banerjee, A.V., Newman, A.F. (1993), Occupational Choice and the Process of Development. *Journal of Political Economy*, 101(2), 274-298.

Banerjee, A.V, Dolado, J., Mestre, R. (1998), Error-correction Mechanism Tests for Cointegration in a Single-equation Framework. *Journal of Time Series Analysis*, 19(3), 267-283.

Bencivenga, V.R., Smith, B.D. (1992), Deficits, Inflation, and the Banking System in Developing Countries: The Optimal Degree of Financial Repression. *Oxford Economic Papers*, 44(4), 767-790.

Bhargava, A. (1986), On the Theory of Testing for Unit Roots in Observed Time Series. *The Review of Economic Studies*, 53(3), 369-384.

Boot, A.W.A. (2000), Relationship Banking: What Do We Know? *Journal of Financial Intermediation*, 9(1), 7-25.

Borensztein, E., De Gregorio, J., Lee, J-W. (1998), How does foreign investment affect economic growth? *Journal of International Economics*, 45(1), 115-135.

Bowers, W.J., Pierce, G.L. (1975), The Illusion of Deterrence in Isaac Ehrlich's Research on Capital Punishment. *The Yale Law Journal*, 85(2), 187-208.

Brown, R.L., Durbin, J., Evans, J.M. (1975), Techniques for Testing the Constancy of Regression Relationships over Time. *Journal of the Royal Statistical Society*, Series B (Methodological), 37(2), 149-192.

Bumann, S., Lensink, R. (2013), *Financial Liberalization and Income Inequality: Channels and Cross-Country Evidence*. Working Paper, University of Groningen, Department of Economics, Econometrics and Finance, The Netherlands.

Bumann, S., Lensink, R. (2016), Capital account liberalization and income inequality. *Journal of International Money and Finance*, 61, 143-162.

Cameron, S. (1994), A review of the econometric evidence on the effects of capital punishment. *The Journal of Socio-Economics*, 23(1-2), 197-214.

Checchi, D., Garcia-Peñalosa, C. (2010), Labour Market Institutions and the Personal Distribution of Income in the OECD. *Economica*, 77(307), 413-450.

Chinn, M.D., Ito, H. (2006), What matters for financial development? Capital controls, institutions, and interactions. *Journal of Development Economics*, 81(1), 163-192.

Chinn, M.D., Ito, H. (2008), A New Measure of Financial Openness. *Journal of Comparative Policy Analysis: Research and Practice*, 10(3), 309-322.

Cornia, G.A. (2005), *Policy Reform and Income Distribution*. DESA Working Paper No. 3, October.

de Haan, J., Sturm, J.-E. (2017), Finance and income inequality: A review and new evidence. *European Journal of Political Economy*, 50, 171-195.

DeJong, D.N., Nankervis, J.C., Savin, N.E., Whiteman, C.H. (1992), Integration Versus Trend Stationarity in Time Series. *Econometrica*, 60(2), 423-433.

Delis, M.D., Hasan, I., Kazakis, P. (2014), Bank Regulations and Income Inequality: Empirical Evidence. *Review of Finance*, 18(5), 1811-1846.

Demirguc-Kunt, A., Levine, R. (2009), Finance and Inequality: Theory and Evidence. *Annual Review of Financial Economics*, 1(1), 287-318.

Dickey, D.A., Fuller, W.A. (1979), Distribution of the Estimators for Autoregressive Time Series With a Unit Root. *Journal of the American Statistical Association*, 74(366), 427-431.

Edwards, S. (1999), How Effective are Capital Controls? *The Journal of Economic Perspectives*, 13(4), 65-84.

Ehrlich, I. (1975), The Deterrent Effect of Capital Punishment: A Question of Life and Death. *American Economic Review*, 65(3), 397-417.

Ehrlich, I. (1977), The Deterrent Effect of Capital Punishment: Reply. *American Economic Review*, 67(3), 452-458.

Ehrlich, I. (1996), Crime, Punishment, and the Market for Offenses. *Journal of Economic Perspectives*, 10(1), 43-67.

Elliott, G., Rothenberg, T.J., Stock, J.H. (1996), Efficient Tests for an Autoregressive Unit Root. *Econometrica*, 64(4), 813-836.

Furceri, D., Loungani, P. (2015), *Capital Account Liberalization and Inequality*. IMF Working Paper WP/15/243.

Galor, O., Zeira, J. (1993), Income Distribution and Macroeconomics. *Review of Economic Studies*, 60(1), 35-52.

Greenwood, J., Jovanovic, B. (1990), Financial Development, Growth, and the Distribution of Income. *Journal of Political Economy*, 98(5), 1076-1107.

Hellmann, T.F., Murdock, K.C., Stiglitz, J.E. (1996), *Deposit mobilisation through financial restraint*. in Niels Hermes and Robert Lensink (eds.), Financial Development and Economic Growth: Theory and Experiences from Developing Countries, 219-246, London: Routledge.

Hellmann, T.F., Murdock, K.C., Stiglitz, J.E. (2000), Liberalization, Moral Hazard in Banking, and Prudential Regulation: Are Capital Requirements Enough? *American Economic Review*, 90(1), 147-165.

Jauch, S., Watzka, S. (2016), Financial Development and Income Inequality: A Panel Data Approach. *Empirical Economics*, 51(1), 291-314.

Jaumotte, F., Buitron, C.O. (2015), *Inequality and Labor Market Institutions*. IMF Staff Discussion Note SDN/15/14.

Jayadev, A. (2007), Capital Account Openness and the Labour Share of Income. *Cambridge Journal of Economics*, 31(3), 423-443.

Johansen, S., Juselius, K. (1990), Maximum Likelihood Estimation and Inference on Cointegration-with Applications to the Demand for Money. *Oxford Bulletin of Economics and Statistics*, 52(2), 169-210.

Kremers, J.J.M., Ericsson, N.R., Dolado, J.J. (1992), The Power of Cointegration Tests. *Oxford Bulletin of Economics and Statistics*, 54(3), 325-348.

Kunieda, T., Okada, K, Shibata, A. (2014), Finance and Inequality: How Does Globalization Change Their Relationship? *Macroeconomics Dynamics*, 18(5), 1091-1128.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., Shin, Y. (1992), Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root? *Journal of Econometrics*, 54(1-3), 159-178.

Lane, P.R., Milesi-Ferretti, G.M. (2007), Europe and Global Imbalances. *Economic Policy*, 22(51), 521-573.

Layson, S. (1983), Homicide and Deterrence: Another View of the Canadian Time-Series Evidence. *Canadian Journal of Economics*, 16(1), 52-73.

Li, J., Yu, H. (2014), Income inequality and financial reform in Asia: the role of human capital. *Applied Economics*, 46(24), 2920-2935.

Mackinnon, J.G. (1996), Numerical distribution functions for unit root and cointegration tests. *Journal of Applied Econometrics*, 11(6), 601-618.

McKinnon, R.I. (1973), *Money and Capital in Economic Development*. Washington: The Brookings Institution.

Narayan, P.K. (2005), The saving and investment nexus for China: evidence from cointegration tests. *Applied Economics*, 37(17), 1979-1990.

Ng, S., Perron, P. (1995), Unit Root Tests in ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag. *Journal of the American Statistical Association*, 90(429), 268-281.

Ng, S., Perron, P. (2001), Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power. *Econometrica*, 69(6), 1519-1554.

Nkoro, E., Uko, A.K. (2016), Autoregressive Distributed Lag (ARDL) Cointegration Technique: Application and Interpretation. *Journal of Statistical and Econometric Methods*, 5(4), 63-91.

Ouattara, B. (2007), Foreign Aid, Public Savings Displacement and Aid Dependency in Côte d'Ivoire: An Aid Disaggregation Approach. *Oxford Development Studies*, 35(1), 33-46.

Pesaran, M.H., Pesaran, B. (1997), *Working with Microfit 4.0: Interactive Econometric Analysis*. Oxford: Oxford University Press.

Pesaran, M.H., Shin, Y. (1999), *An Autoregressive Distributed-Lag Modelling Approach to Cointegration Analysis*. In Steinar Strøm (ed.), Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium, 371-413, Cambridge: Cambridge University Press.

Pesaran, H.M., Shin, Y., Smith, R.J. (2001), Bounds testing approaches to the analysis of level relationships. *Journal of Applied Econometrics*, 16(3), 289-326.

Phillips, P.C.B., Perron, P. (1988), Testing for a Unit Root in Time Series Regression. *Biometrica*, 75(2), 335-346.

Pierse, R.G., Snell, A.J. (1995), Temporal aggregation and the power of tests for a unit root. *Journal of Econometrics*, 65(2), 333-345.

Said, S.E., Dickey, D.A. (1984), Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order. *Biometrika*, 71(3), 599-607.

Schwert, G.W. (1989), Tests for Unit Roots: A Monte Carlo Investigation. *Journal of Business and Economic Statistics*, 7(2), 147-160.

Shaw, E.S. (1973), *Financial Deepening in Economic Development*.

Oxford, New York: Oxford University Press.
Solt, F. (2019), *Measuring Income Inequality Across Countries and Over Time: The Standardized World Income Inequality Database*. SWIID Version 8.1, May.
Stulz, R.M. (1999), Globalization of Equity Markets and the Cost of Capital. *NBER Working Paper No.7021*, National Bureau of Economic Research, Inc.
Tan, H-B., Law, S-H. (2012), Nonlinear Dynamics of the Finance-Inequality Nexus in Developing Countries. *Journal of Economic Inequality*, 10(4), 551-563.
Taylor, L. (1983), *Structuralist Macroeconomics: Applicable Models for the Third World*. New York: Basic Books.
van Wijnbergen, S. (1982), Stagflationary effects of monetary stabilization policies: A quantitative analysis of South Korea. *Journal of Development Economics*, 10(2), 133-169.
Wahid, A.N.M., Shahbaz, M., Shah, M., Salahuddin, M. (2012), Does Financial Sector Development Increase Income Inequality? Some Econometric Evidence from Bangladesh. *Indian Economic Review*, 47(1), 89-107.
Zapata, H.O., Maradiaga, D.I., Pujula, A.L., Dicks, M.R. (2011), *Recent Developments in Unit Root Tests and Historical Crop Yields*. Agricultural and Applied Economics Association, 2011 Annual Meeting, July 24-26, Pittsburgh, Pennsylvania 103871.
Zhang, R., Naceur, S.B. (2019), Financial development, inequality, and poverty: some international evidence. *International Review of Economics and Finance*, 61, 1-16.