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The relationship between financial development and income inequality in Turkey

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Abstract

Reducing income inequality and enhancing access to financial institutions play a vital role in achieving the Sustainable Development Goal 10, especially in developing countries. While there are studies on the nexus between financial development and economic growth, literature on financial development–income inequality nexus is limited and lack consensus. This paper examined the different dimensions of financial development on income inequality for the period 1990 to 2015 in Turkey. For this purpose, four financial development indices (i.e., overall financial development index, banking sector development index, stock market development index, and bond market development index) were constructed with principal component analysis (PCA). In addition to financial development indicators, the impact of real income, government expenditures, and inflation on income inequality were investigated using the ARDL bound testing procedure. The results showed that increasing real income and government expenditures reduce income inequality. We found a positive impact of inflation on income inequality in the short run, while the reverse holds in the long run. In the case of financial development, an inverted U-shaped relationship with income inequality for overall financial development and banking sector development was confirmed. A monotonically decreasing relationship between stock market development and income inequality was found in Turkey. The study highlights that the implementation of policies that eradicate discriminatory laws and facilitate equal access to financial privileges will promote equity and decline the outcomes of income inequality.

Keywords: Financial development, Income inequality, Financial markets, Sustainable Development Goal 10, Turkey

1 Introduction

The association between financial development and income inequality has been the interest of researchers for a stint period. The earliest studies in this area started with analyzing the association between economic development and income inequality. The seminal work by Simon Kuznets (1955) focused on the association between economic growth and income inequality, and the study showed that income inequality rises during the agrarian phase of economic development, slows down during the industrial development, and declines during the rise of the service sector. Therefore, the study by Kuznets (1955) hypothesized an inverted U-shaped association between economic development and income inequality. Taking a cue from this study, a number of researchers extended this study by analyzing the association between financial development and income inequality. This extension can be explained logically. When the financial sector of any nation starts developing through several channels, namely banking and financial services sector, it directly affects the economic growth pattern and, subsequently, the distribution of income. Now, the distribution of income, which is characterized by disproportionate economic growth, is directly impacted by the development of the financial sector. On the other hand, the financial development of any nation also explains the allocation of monetary resources towards enhancing the quality of life, which is largely the basic premise of economic development. In view of this, it can be said that economic development catalyzed by financial development might have a significant impact on the distribution of income, disproportion in which might lead to income inequality. Researchers have focused on analyzing the association between financial development and income inequality.

Now, this association might prove to be beneficial for a nation which is recognized by industrial growth. Owing to this reason, the present study analyzes the long-run association between financial development and income inequality in Turkey over the period 1990–2015. According to World Bank (2016a), Turkey is categorized as one of the newly industrialized countries under the next 11 categories. During 1990–2015, the global per capita income grew by nearly 1.34 times, whereas the per capita income of Turkey grew by nearly 2.27 times (World Bank 2016b). As of 2015, the global GDP growth rate was approximately 2.47%, whereas the GDP growth rate of Turkey was approximately 3.98% (World Bank 2016b). This shows the growth potential of Turkey in recent years, and owing to the pattern of this economic growth, the income inequality has been reduced by 12.19% during 1990–2015 while demonstrating dimensions of volatility (Fig. 1). It shows that the economic growth fueled by the financial development in Turkey is yet to achieve a proportionate trajectory, where the income distribution can be stabilized. The present study sheds light on this area, which has been overlooked in the existing literature.

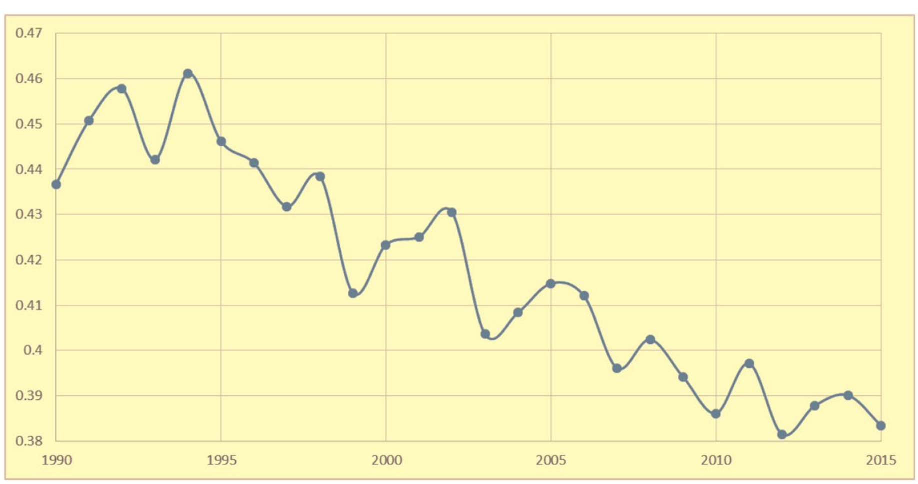

Fig. 1 Movement of income inequality in Turkey, 1990–2015

In line with the Greenwood–Jovanovic (GJ) hypothesis (1990), the present study uses four financial development indicators and assess the impacts of those indicators on income inequality, following an inverted U-shaped framework. This study contributes to the existing literature in several ways: first, we analyze the financial development–income inequality nexus in both integrated and disintegrated ways; second, we analyze the inverted U-shaped association between financial development and income inequality for four different financial development indicators; third, we develop a comprehensive financial development index for Turkey, and analyze the financial development–income inequality nexus for Turkey following the GJ hypothesis. The results of this study confirm the GJ hypothesis for Turkey.

2 Literature review

The seminal work by Kuznets (1955) was the first to identify the link between economic development and income inequality. Subsequent to that study, researchers came up with a new strand of literature, which deals with the association between financial development and income inequality. Over the years, researchers have identified several channels through which financial development affects income inequality. In assessing existing literature, we examine studies on the nexus between financial development and income inequality in different contexts and identify the mentioned channels.

While considering the association between financial development and income inequality, the study by Greenwood and Jovanovic (1990) needs to be mentioned. They were one of the proponents of the financial Kuznets curve hypothesis, which is commonly referred to as Greenwood–Jovanovic (GJ) hypothesis. According to this hypothesis, income inequality rises at the initial phase of financial development, slows down with growth in financial development, and falls during maturity. Therefore, the association between financial development and income inequality follows an inverted U-shaped form, which is also known as financial Kuznets curve. In the study, the researchers focused on the role of intermediaries in collecting and analyzing information, and how it catalyzes the allocation of funds in the economy for achieving a highest social return, in terms of reduced income inequality. This was the first study in the literature to give the financial development–income inequality association a formalized shape.

The relationship between financial development and income inequality was investigated for Pakistan over the period 1971–2005 (Shahbaz and Islam 2011). By adopting the GJ hypothesis and following the Autoregressive Distributed Lag (ARDL) bound testing approach, they found no evidence in support of the GJ hypothesis. In the study, financial development was measured by domestic credit distributed to the private sector as a share of GDP. In a subsequent study and following the similar theoretical framework, Shahbaz et al. (2015) analyzed this association in Iran over the period 1965–2011. Using the ARDL approach, they found evidence in favor of GJ hypothesis. Tiwari et al. (2013) analyzed the impact of financial development on rural–urban income inequality in India over the period 1965–2008. Using the ARDL approach, the researchers found financial development to aggravate rural–urban income inequality in the long run. Law et al. (2014) investigated this association for a total of 81 countries over the period 1985–2010. Using threshold cointegration approach, results of the study indicate that the association between financial development and income inequality is significantly moderated by institutional quality, and better institutional quality helps the channels of financial development to reduce income inequality. The long-run and the short-run heterogeneous association between financial development and income inequality were analyzed in 88 countries over the period 1961–2012 (Chen and Kinkyo, 2016). Using the pooled mean group (PMG) estimation approach, the researchers found that financial development reduces inequality in the long-run, while it increases inequality in the short run. Jauch and Watzka (2016) analyzed a similar association in 138 developed and developing countries over the period 1960–2008. Taking private credit to GDP as the measure of financial development and using fixed-effect two-stage least-squares (2SLS) estimation, the researchers found financial development to have a positive effect on income inequality. Seven and Coskun (2016) analyzed this association in 45 emerging economies over the period 1987–2011 and using the generalized method of moments (GMM), the researchers found that financial development impacts income inequality majorly in low-income emerging economies. In the study, the researchers used a total of eight indicators for financial development. De Haan and Sturm (2017) analyzed the association between financial development, financial liberalization, banking crises, and income inequality in 121 countries over the period 1975–2005. The study found that financial development conditions the impact of financial liberalization on income inequality. Park and Shin (2017) analyzed the association in 162 countries over the period 1960–2011. By following pooled and panel modeling approach, the researchers found that financial development contributes to reducing inequality up to a point, but as financial development proceeds further, it contributes to greater inequality. However, in the presence of positive social indicators, financial development becomes more effective in reducing inequality. In the study, financial development was measured by liquid liabilities as a percentage of GDP, private credit by deposit money banks as a percentage of GDP, and stock market capitalization as a percentage of GDP. Liu et al. (2017) investigated this association for the case of 23 Chinese provinces over the period 1996–2012. Using GMM, the researchers found the association to be linear and inverted U-shaped, thereby, validating the evidence of financial Kuznets curve. Income inequality was segregated for rural and urban area districts. Azam and Raza (2018) analyzed the influence of financial sector development on income inequality in ASEAN-5 countries over the period 1989–2013, and using fixed-effect model, they found the evidence of financial Kuznets curve. Financial development was measured by domestic credit by the banking sector, domestic credit to the private sector, money supply, and stock market capitalization.

By far, we have found only one study (Yeldan 2000) which has considered the association between some aspects of financial development and income inequality for Turkey. The study was aimed at discovering the impacts of financial liberalization and financial rents on income distribution for Turkey during the post-1980 period. Results of the study showed that political and industrial structure of an economy can have a significant impact on the shape of income distribution and overall economic development. However, the study failed to cover the monetary channels of financial development, which is the focus of this study. As a contribution to the extant literature, this study employs the banking sector, stock market, and bond market as the channels of financial development, and the empirical analysis is carried out within the framework of the GJ hypothesis.

3 Data and methodology

3.1 Empirical model and data

The annual data spanning 1990–2015 were collected to examine the relationship between financial development and income inequality in Turkey. In addition, real income, government expenditures, and inflation were added to the empirical model as explanatory variables to control for the omitted variable bias. The main empirical model is constructed as follows:

$$\ln \mathrm{INE}_t = \beta_0 + \beta_1 \ln Y_t + \beta_2 \mathrm{INF}_t + \beta_3 \ln G_t + \beta_4 \ln \mathrm{FD}_t + \beta_5 \ln \mathrm{FD}_t^2 + \varepsilon_t, \tag{1}$$

where t and ε_t are the time period and residual term, respectively. In addition, lnINE, lnY, lnINF, lnG, and lnFD indicates the natural log of Gini coefficient as a proxy for income inequality, the real gross domestic product (GDP) per capita measured in constant 2010 US dollar as a proxy for economic growth, consumer price index as a proxy for inflation, government expenditures' share in GDP, and financial development indicators. To observe the possible non-linear relationship between financial development and income inequality, we plug in the square of financial development ($\ln\mathrm{FD}^2$) as an explanatory variable. Namely, inequality reducing hypothesis is confirmed in the case of $\beta_4 < 0$ and $\beta_5 = 0$; inequality increasing hypothesis is confirmed if $\beta_4 > 0$ and $\beta_5 = 0$; the GJ hypothesis is confirmed if the estimated parameters follow $\beta_4 > 0$ and $\beta_5 < 0$; and the U-shaped relationship is accepted if $\beta_4 > 0$ and $\beta_5 < 0$.

Using different variables to indicate financial development such as financial system deposit to GDP, liquid liabilities to GDP and private credit by deposit money banks to GDP, stock market capitalization to GDP and stock market turnover ratio, are inappropriate for capturing financial development—as all these variables are highly correlated. In addition, separating the development of the financial sector into sub-segments such as the banking sector, the stock market, and the bond market, and examining the effectiveness of these sectors separately allows for more consistent policy implications. Based on these reasons, we used the principal component analysis (PCA) to construct a financial development index (FD) which includes three sub-indices. The first sub-index of financial development index is banking sector development index (BAD), and this index is constructed using deposit money bank assets to GDP, financial system deposit to GDP, liquid liabilities to GDP, and private credit by deposit money banks to GDP. The second sub-index is stock market development index which is computed using the stock market capitalization to GDP, stock market turnover ratio, and stock market total value traded to GDP. The third sub-index is bond market development index which covers the outstanding domestic private debt securities to GDP, the outstanding domestic public debt securities to GDP, the outstanding international private debt securities to GDP, and the outstanding international public debt securities to GDP. The annual data of real GDP per capita, consumer price index, and government expenditures were downloaded from the World Development Indicators—under the auspices of the World Bank. The variables used for constructing the financial development index were obtained from Global Financial Development database of World Bank. In addition, the Gini coefficient data were retrieved from the Standardized World Income Inequality Database (SWIID 6.1) (Solt 2016).

3.2 Empirical methodology

3.2.1 Integration process

Ignoring the possibility of structural breaks will exhibit size distortions that "over reject" the null hypotheses of a unit root. Based on this reason, this study employed the unit root test by Lee and Strazicich (2003, 2004) which allows a maximum of two breaks in the testing procedure to examine the stationary properties of variables with endogenous breaks. The data-generating process of Lagrange Multiplier (LM) statistic of Lee and Strazicich (2003, 2004) is as follows:

$$\Delta y_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + u_t, \tag{2}$$

where $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t \tilde{\delta} (t = 2 \ldots T)$, Z_t is a vector of exogenous variables defined by data generation process, $\tilde{\delta}$ is the vector of coefficients in the regression of Δy_t and ΔZ_t, and $\tilde{\psi}_x = y_1 - Z_1 \tilde{\delta}$. The null hypothesis is described as $\phi = 0$, and the augmented terms $\Delta \tilde{S}_{t-j}, j = 1, \ldots, k$ are included to correct the serial correlation. The general-to-specific search procedure is used to determine k value. The LM unit root procedure searches for possible breakpoints for the minimum t-statistics to endogenously determine the location of breaks (T_{B}) as follows:

$$\mathrm{Inf}\tilde{\tau}(\tilde{\lambda}) = \mathrm{Inf}_{\lambda} \tilde{\tau}(\lambda); \quad \text{where } \lambda = T_{\mathrm{B}}/T. \tag{3}$$

The critical values for the two-break case are tabulated in Lee and Strazicich (2003) and the critical values for the one-break case are tabulated in Lee and Strazicich (2004).

3.2.2 Cointegration process

This study used the autoregressive distributed lag (ARDL) approach of Pesaran et al. (2001) to examine the validity of the cointegration between variables and to determine both the long-run and the short-run coefficients of the variables. The main advantage of the ARDL estimation approach is the non-binding pre-test requirements. Namely, the ARDL method allows variables that are stationary in levels [I(0)] or first-differenced form [I(1)]. Because of this feature, the ARDL method has been used in many studies. In this procedure, computed F-statistics are compared with two bounds of critical values. The null hypothesis that there is no cointegration is rejected if computed F-statistic exceeds the upper critical value. The null hypothesis is accepted if the computed F-statistic smaller than lower critical bound. The relevant ARDL procedure of Eq. 1 can be written as follows:

$$\begin{aligned}\Delta \ln \mathrm{INE}_t = \beta_0 &+ \sum_{i=1}^{n} \beta_{1,i} \Delta \ln \mathrm{INE}_{t-i} + \sum_{i=0}^{n} \beta_{2,i} \Delta \ln Y_{t-i} + \sum_{i=0}^{n} \beta_{3,i} \Delta \ln \mathrm{INF}_{t-i} \\ &+ \sum_{i=0}^{n} \beta_{4,i} \Delta \ln G_{t-i} + \sum_{i=0}^{n} \beta_{5,i} \Delta \ln \mathrm{FD}_{t-i} + \sum_{i=0}^{n} \beta_{6,i} \Delta \ln \mathrm{FD}^2_{t-i} \\ &+ \delta_1 \ln \mathrm{INE}_{t-1} + \delta_2 \ln Y_{t-1} + \delta_3 \mathrm{INF}_{t-1} + \delta_4 \ln G_{t-1} \\ &+ \delta_5 \ln F_{t-1} + \delta_6 \ln F^2_{t-1} + \mu_t,\end{aligned} \tag{4}$$

where Δ and n indicate the difference operator and lag length, respectively. According to Eq. 4, the null hypothesis of no cointegration between variables $H_0: \delta_1 = \delta_2 = \delta_3 = \delta_4 = \delta_5 = \delta_6 = 0$ is tested against the alternative hypothesis $H_I: \delta_1 \neq \delta_2 \neq \delta_3 \neq \delta_4 \neq \delta_5 \neq \delta_6 \neq 0$. The optimal lag length (n) in Eq. 4 is chosen with Schwarz information criteria (SIC). If there is cointegration among variables, the long-run ARDL equation is estimated as follows:

$$\ln \mathrm{INE}_t = \beta_0 + \sum_{i=1}^{p} \beta_{1,i} \ln \mathrm{INE}_{t-i} + \sum_{i=0}^{q} \beta_{2,i} \ln Y_{t-i} + \sum_{i=0}^{r} \beta_{3,i} \ln \mathrm{INF}_{t-i} + \sum_{i=0}^{s} \beta_{4,i} \ln G_{t-i} + \sum_{i=0}^{t} \beta_{5,i} \ln \mathrm{FD}_{t-i} + \sum_{i=0}^{v} \beta_{6,i} \ln \mathrm{FD}^2_{t-i} + u_t, \tag{5}$$

where the optimum lag values of *p, q, r, s, t,* and *v* are also selected using with SIC. Finally, short-run coefficients of the variables are estimated with error-correction model as follows:

$$\ln \mathrm{INE}_t = \beta_0 + \sum_{i=1}^{p} \beta_{1,i} \Delta \ln \mathrm{INE}_{t-i} + \sum_{i=0}^{q} \beta_{2,i} \Delta \ln Y_{t-i} + \sum_{i=0}^{r} \beta_{3,i} \Delta \ln \mathrm{INF}_{t-i} + \sum_{i=0}^{s} \beta_{4,i} \Delta \ln G_{t-i} + \sum_{i=0}^{t} \beta_{5,i} \Delta \ln \mathrm{FD}_{t-i} + \sum_{i=0}^{v} \beta_{6,i} \Delta \ln \mathrm{FD}^2_{t-i} + \gamma \mathrm{ECM}_{t-1} + \vartheta_t, \tag{6}$$

where the coefficient (γ) of an error-correction term (ECM_{t-1}) is the speed of adjustment parameter and the expected sign of this coefficient should be negative with statistical significance.

4 Empirical results

In the first step of the analysis, we constructed the sub-indices of financial development index and overall financial development index using PCA analysis, as shown in Table 1. In the case of banking development index, the eigenvalues reveal that the first principal component (PCA1) is the best principal component, because it explains about the 94.7% of the standardized variance. The individual contributions of deposit money bank assets to GDP (DMB), financial system deposit to GDP (FSD), liquid liabilities to GDP (LL), and private credit by deposit money banks to GDP (PC) to standardized variance of PCA1 (i.e., 25.59, 25.34, 23.98, and 25.04%) were used as the weights to obtain the banking sector development index (BAD). In the case of stock market development, the first principal component (PCA1) explains about the 71.2% of the standardized variance and the individual contributions of the stock market capitalization to GDP (SC), stock market turnover ratio (ST), and stock market total value traded to GDP (STR) to standardized variance of PCA1 are 30.89%, 20.72%, and 33.40%, respectively. In case of bond market development, the first principal component analysis (PCA1) explains about the 56.3% of the standardized variance and the weights of outstanding domestic private debt securities to GDP, the outstanding domestic public debt securities to GDP, the outstanding international private debt securities to GDP, and the outstanding international public debt securities to GDP are 24.76%, 16.78%, 29.51, and 27.12%, respectively.

After constructing the financial development indicators, we examined the stationary properties of variables using with unit root test developed by Lee and Strazicich (2004) which allows one endogenous structural break. As shown in Table 2, all variables are

Table 1 Principal component analysis

Index: BAD	PCA1	PCA2	PCA3	Pca4
Eigenvalues	3.7890	0.2008	0.0078	0.0022
Proportion	0.9473	0.0502	0.0020	0.0006
Cumulative proportion	0.9473	0.9975	0.9994	1.0000
Variables	**Vector 1**	**Vector 2**	**Vector 3**	**Vector 4**
DMB	0.5119	0.0783	− 0.8553	0.0130
FSD	0.5069	− 0.349	0.2601	− 0.744
LL	0.4797	0.7957	0.3612	0.0799
PC	0.5008	-0.4888	0.2651	0.6633
Index:SMD	**PCA1**	**PCA2**	**PCA3**	
Eigenvalues	2.1359	0.8051	0.0589	–
Proportion	0.7120	0.2684	0.0197	–
Cumulative proportion	0.7120	0.9803	1.0000	–
Variables	**Vector 1**	**Vector 2**	**Vector 3**	
SC	0.6179	− 0.4454	0.6479	–
ST	0.4145	0.8848	0.2129	–
STR	0.6681	− 0.1370	− 0.7314	–
Index: BND	**PCA1**	**PCA2**	**PCA3**	**PCA4**
Eigenvalues	2.2539	1.4405	0.1940	0.1113
Proportion	0.5635	0.3601	0.0485	0.0278
Cumulative proportion	0.5635	0.9236	0.9722	1.0000
Variables	**Vector 1**	**Vector 2**	**Vector 3**	**Vector 4**
DPRI	0.4952	− 0.5227	0.0200	0.6936
DPUB	0.3355	0.6819	0.5969	0.2572
IPRI	0.5901	− 0.3176	0.3198	− 0.6698
IPUB	0.5423	0.4011	− 0.7355	− 0.0636

Table 2 Lee and Strazicich's (2004) test with one endogenous structural break

Variables	*t*-statistics	TB	Variables	*t*-statistics	TB
lnINE	− 3.750	2001	ΔlnINE	− 5.304***	2001
lnY	− 4.112	1999	ΔlnY	− 5.167***	1996
lnINF	− 4.314*	2001	ΔlnINF	− 5.159***	2000
lnG	− 4.144	2007	ΔlnG	− 6.466***	2008
lnFD	− 3.568	2001	ΔlnFD	− 6.684***	2002
lnBAD	− 3.802	2007	ΔlnBAD	− 6.534***	2011
lnSMD	− 4.340*	2004	ΔlnSMD	− 5.370***	2001
lnBND	− 3.512	2001	ΔlnBND	− 6.482***	2001

* and *** indicates statistical significance at 10 and 1% level, respectively

non-stationary at the level form of variables. However, the null of the unit root process is rejected at the first-differenced form and all series have become stationary.

In the next step, we used the time break (2001) that was obtained from the unit root test of Lee and Strazicich (2004) as a dummy variable and the ARDL cointegration test is employed. As shown in Table 3, the F-statistic exceeds critical bound for all

Table 3 Results of the ARDL cointegration and diagnostic tests

Estimated model	*INE = f(Y,INF,G,FD)*	*INE = f(Y,INF,G,BAD)*	*INE = f(Y,INF,G,SMD)*	*INE = f(Y,INF,G,BND)*
Lag order	(2,1,2,2,1)	(2,0,2,1,2)	(1,2,2,1,2)	(1,0,2,0,0)
F-statistics	6.568**	7.706***	8.741***	7.475**
Structural breaks	2001	2001	2001	2001
Critical values	1%	5%	10%	
Lower bound	5.856	4.154	3.430	
Upper bound	7.578	5.540	4.624	
Diagnostic tests				
$\chi^2 NORMAL$	0.463 [0.793]	0.590 [0.744]	0.913 [0.633]	0.438 [0.803]
$\chi^2 SERIAL$	1.476 [0.263]	1.541 [0.244]	2.959 [0.155]	1.587 [0.236]
$\chi^2 ARCH$	1.777 [0.196]	0.036 [0.851]	0.259 [0.615]	0.298 [0.590]
$\chi^2 WHITE$	0.836 [0.636]	1.665 [0.149]	1.113 [0.448]	1.707 [0.173]
$\chi^2 RAMSEY$	0.842 [0.389]	0.918 [0.371]	1.536 [0.162]	0.041 [0.967]
CUSUM	Stable	Stable	Stable	Stable
CUSUMQ	Stable	Stable	Stable	Stable

** and *** indicates statistical significance at 5 and 1% level, respectively. The numbers in brackets are *p* values. The critical values are obtained from Narayan (2005)

models. Therefore, we conclude that there is a long-run relationship between variables for all models. In addition, diagnostic tests are shown in Table 3. From these results, it seems ARCH test results support that residuals are homoscedastic, Ramsey–Reset test confirms the correct functional form. To examine the normality behavior of estimated residuals, we utilized the Jarque–Berra statistic and the result confirms the normality behavior. The result of the Breusch–Godfrey LM test rejects the presence of serial correlation of the residuals. Moreover, as shown in Fig. 2, the stability properties are examined with CUSUM and CUSUMQ tests.

Next, we examined the short- and long-run effects of the real income per capita, inflation, and the government expenditures with different financial development indicators on income inequality, as shown in Table 4. In the short run, the results from the first model show that increasing real income per capita and government expenditures reduce the income inequality. However, inflation has an insignificantly positive effect on income inequality. In the case of financial development, it seems the coefficient of overall financial development index (the square of the financial development index) which is positive (negative). Since there is evidence of an inverted U-shaped relationship between financial development and income inequality, we confirm the validity of the GJ hypothesis for overall financial development index. It can be observed from Model II that the effect of real income per capita and government expenditures are negative, and the effect of inflation on income inequality is positive. In addition, the coefficient of the banking development index is positive and the coefficient of the square of the banking sector development is negative. Therefore, we validate the inverted U-shaped relationship between banking sector development and income inequality. For Model III and Model IV, similar to the first and second model, the effects of the real income per capita and government expenditures are negative, and the effect of inflation is negative. In the

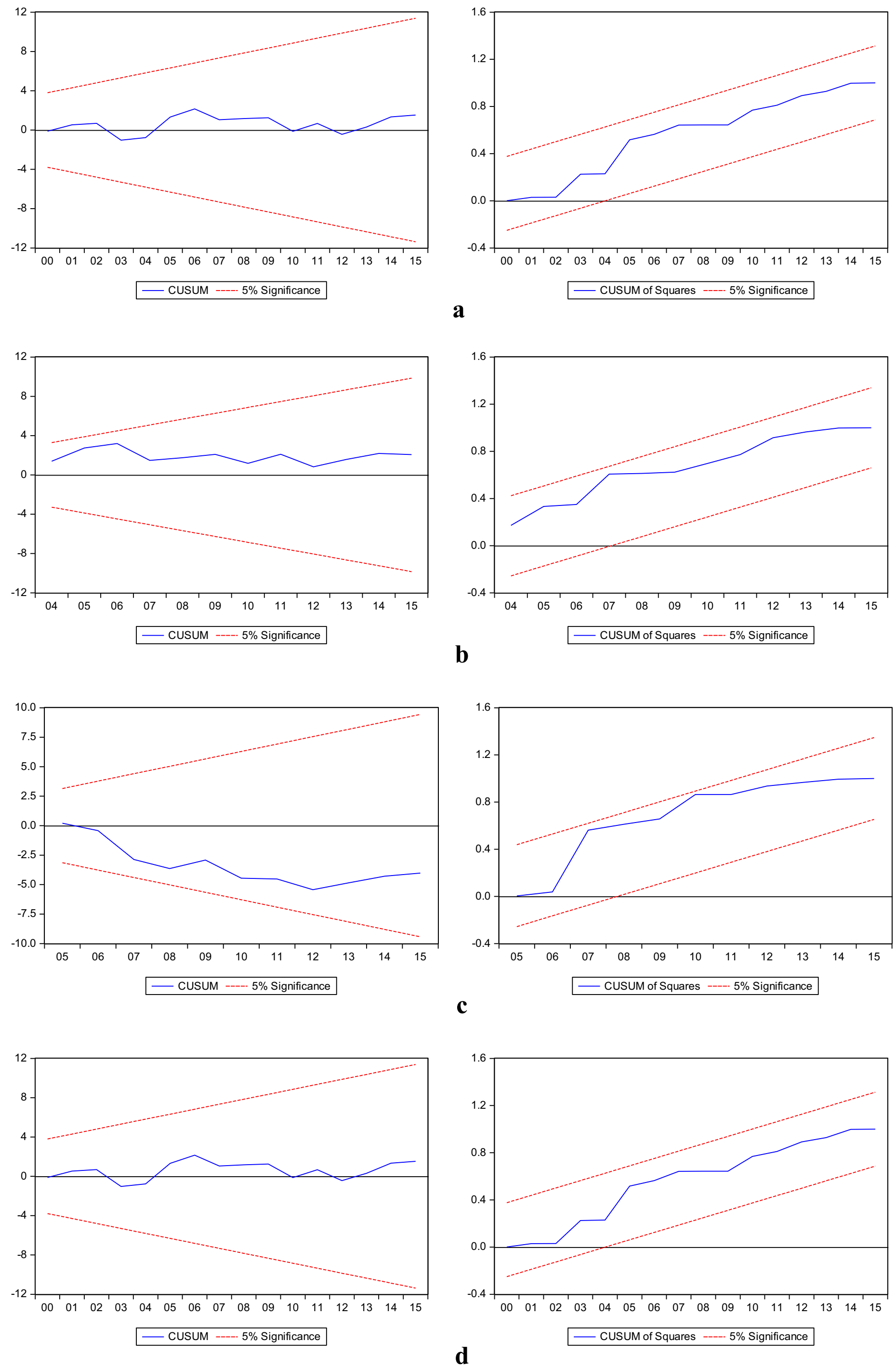

Fig. 2 The plots of the cumulative sum and the cumulative sum of the squares of recursive residuals. **a** Overall financial development index**. b** Banking sector development index. **c** Stock market development index. **d** Bond market development index

Table 4 The results of the short run and long run

Variables	Model I	Model II	Model III	Model IV
Short-run results				
lnY	− 0.145** (− 2.721)	− 0.051 (− 0.856)	− 0.163*** (− 4.714)	− 0.209*** (− 3.342)
lnINF	0.020 (0.512)	0.056*** (5.209)	0.178*** (6.646)	0.087* (2.036)
lnG	− 0.103** (− 2.910)	− 0.205*** (− 4.982)	− 0.068** (− 2.475)	− 0.226*** (− 4.253)
lnFD	0.044*** (4.465)	–	–	–
$lnFD^2$	− 0.019*** (− 4.198)	–	–	–
lnBAD	–	0.035** (2.597)	–	–
$lnBAD^2$	–	− 0.020*** (− 3.152)	–	–
lnSMD	–	–	− 0.020*** (− 5.754)	–
$lnSMD^2$	–	–	− 0.005*** (− 4.206)	–
lnBND	–	–	–	− 0.006 (− 0.543)
$lnBND^2$	–	–	–	0.004 (1.043)
ECT (− 1)	− 0.090*** (− 4.970)	− 0.218*** (− 4.882)	− 0.911*** (− 6.277)	− 0.582*** (− 6.894)
Long-run results				
lnY	− 0.051* (− 1.997)	− 0.023** (− 2.525)	− 0.124*** (− 4.807)	− 0.155*** (− 3.313)
lnINF	0.030*** (4.803)	0.025*** (4.557)	0.009** (2.220)	0.012** (2.920)
lnG	− 0.108 (− 1.728)	− 0.174*** (− 3.590)	− 0.227** (− 3.205)	− 0.287** (− 3.021)
lnFD	0.051*** (3.359)	–	–	–
$lnFD^2$	− 0.015** (− 2.416)	–	–	–
lnBAD	–	0.037*** (3.023)	–	–
$lnBAD^2$	–	− 0.021** (− 2.835)	–	–
lnSMD	–	–	− 0.009* (− 1.967)	–
$lnSMD^2$	–	–	− 0.006* (− 1.998)	–
lnBND	–	–	–	0.012 (0.747)
$lnBND^2$	–	–	–	0.002 (0.443)

*, **, and *** indicates statistical significance at 10, 5, and 1% level, respectively. Numbers in parentheses are *t*-statistics. The coefficients of lagged variables are not shown for short-run results

case of stock market development, both the coefficient of the stock market development index and the square of the stock market development index are significantly negative. This confirms a monotonically decreasing relationship between stock market development and income inequality. The empirical results further reveal that the relationship between the bond market development index and income inequality is statistically insignificant. When we observe the error-correction terms (ECT), it appears the error-correction terms for all models are significantly negative, confirming the speed of correcting previous disturbances into an equilibrium state.

In the long run, increasing real income per capita and government expenditures negatively affect the Gini coefficient for all models similar to the short run. In addition, the effect of the inflation on income inequality is positive for all models, similar to the finding from the short-run. Moreover, in the long run, it seems there is an inverted U-shaped relationship between overall financial development index and income inequality, and monotonically decreasing relationship between stock market development index and income inequality. The validity of an inverted U-shaped relationship between banking sector development index and income inequality is confirmed. However, there is no statistically significant relationship between bond market development and income inequality.

In totality, economic growth and increasing government size reduce income inequality for both in the short run and the long run. However, the effect of inflation on income inequality is positive both for the short run and long run. In the case of financial development, the GJ hypothesis is confirmed, which argues that there is an inverted U-shaped relationship between financial development and income inequality for overall financial development index in Turkey. Similarly, we confirm an inverted U-shaped relationship for the banking sector development. Our results confirm a monotonically decreasing relationship between stock market development and income inequality. However, the findings show that there is no statistically significant relationship between bond market development and income inequality in Turkey.

5 Conclusion

This paper examined the validity of Greenwood–Jovanovic hypothesis for different financial development indicators for the period 1990 to 2015 in Turkey. In doing so, we first utilized the principal component analysis to construct four financial development indicators (i.e., overall financial development index, banking sector development index, stock market development index, and bond market development index). After this procedure, we investigated the effect of the real income per capita, inflation, government expenditures, and financial development indicators on the Gini coefficient. We utilized the unit root test with structural breaks and examined the relationship between the variables with the ARDL bounds testing procedure including the time break which was endogenously obtained from the unit root test.

The empirical findings showed that increasing real income and government expenditures reduce income inequality in Turkey for both short- and long-run relationship. The first finding revealed that low-income segments benefit more than high-income segments from increasing prosperity as a result of economic growth. The second finding showed that increasing government expenditures reduces inequality, implying that transfer spending has been successfully implemented in terms of reducing income inequality in Turkey. In addition, our result showed that the positive impact of inflation on income inequality is valid in both the short run and the long run. In the case of financial development, we found evidence of the validity of the GJ hypothesis which confirms an inverted U-shaped relationship between financial development and income inequality for both overall financial development index and banking sector development index. These results mean that income distribution is adversely affected by financial development in the initial stages of the development of the banking sector. However, at a later stage, minimized systemic risk in the banking sector facilitates the accessibility of credit by the low-income segment, thereby reducing income inequality. We also found a monotonically decreasing relationship between stock market development and income inequality. It appears that the stock market liquidity directly benefits the population with low-income levels in Turkey. On the other hand, our results revealed that the effect of bond market development on income inequality is positive but statistically insignificant. The failure of the bond market to reduce inequality is largely due to the structure of the domestic public debt. Because the public sector debt securities with high-interest rates are purchased mainly by banks in recent decades. When the average distribution of consolidated budget expenditures of Turkey is observed for recent decades, it can be observed that the share of education and health expenditures are 13.2% and 3.4%, while the share of domestic debt interest payments is 20.5% (Bedir and Karabulut 2011).

As a policy implication, the implementation of policies that ensure equal opportunities for low-income sectors to access financial support and financial instruments for creating their own businesses should be enhanced. Investment in education and health sectors that are considered to reduce income inequality should also be supported by the financial sector.

Authors' contributions

MAD conceptualized and analyzed the data; AS helped in the write-up of the manuscript; and SAS assisted in writing and proofreading. All authors read and approved the final manuscript.

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Availability of data and materials

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Competing interests

The authors declare that they have no competing interests.

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