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DOES FINANCIAL INCLUSION IMPROVE INCOME EQUALITY? THE CASE OF TÜRKİYE

Kemal Erkişi[1], Semra Boğa[2]

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[1] Antalya Bilim University, Antalya, Türkiye

[2] Final International University, Kyrenia, Cyprus

Corresponding Author:
Semra Boğa

Email: semra.boga@final.edu.tr

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ABSTRACT

Income inequality has become an important economic and humanitarian problem for both advanced and emerging economies, especially with the increase in financialization trends. The equitable distribution of income has garnered attention in both developed and developing nations, given the rise in global trade and production. However, limited research has explored the impact of financial inclusion on income inequality. To address this gap, this study investigated the effect of financial inclusion on income inequality in Türkiye, contributing to the very limited literature. In this study, the financial inclusion variable is measured using a six-dimensional index encompassing financial institutions and financial markets with depth, access, and availability sub-dimensions. During the time frame spanning from 1980 to 2021, estimations of parameters are conducted employing cointegration regression techniques, including FMOLS, DOLS, and CCR. The analysis revealed that inflation, per capita income, urbanization, and financial inclusion have a negative impact on income equality, whereas education has a positive impact. A 1% rise in financial inclusion is associated with a proportional rise in income inequality of approximately 0.012%. Contrary to the findings of previous empirical studies in general, the increase in financial inclusion in Türkiye has a distorting effect on income equality. The findings of this study offer important implications for Türkiye. While the relationship between the financial inclusion indicator and income inequality is not negative, increasing the income of low-income groups across all financial sectors is likely to improve income equality.

Keywords: *Financial Inclusion, Income Inequality, FMOLS, DOLS, CCR*

1. INTRODUCTION

Türkiye has emerged as the 19th largest economy globally, owing to its strategic location in the Middle East and a youthful demographic. The country has maintained an average growth rate of 5% since the post-1980 period (The World Bank, 2023). Although the main drivers of economic growth have been consumption expenditures and exports, it is undeniable that the financial liberalization reforms made in 1980 transformed the Turkish economy and contributed to economic growth. With the January 24, 1980 decisions, Türkiye removed the legal obstacles for capital flows and opened its economy to the outside world, integrating itself to the global capital. While an important feature of the post-1980 period in which neo-liberal policies were adopted is the increase in trade volumes, the increase in financialization trends was another important reflection of the paradigm shift. The free entry of foreign capital flows into the country has undoubtedly been the main trigger of financialization in Türkiye. The establishment of the Capital Market in 1981, making the Turkish Lira convertible against other currencies with the decision taken in 1989, selling government securities to foreigners and allowing banks to borrow freely from abroad were important practices that accelerated financialization (Heydarova, 2020). Especially in the post-1990 period, with the acceleration of globalization movements all over the world, the increase in credit volumes in Türkiye accelerated, and the share of the finance sector in

GDP gained a significant momentum.

Although it is clear that the financial liberalization and deepening efforts in Türkiye after 1980 contributed significantly to economic growth, it is difficult to state that this growth is shared fairly among different classes of society since the share of the richest 10% received 54.47% of total income in 2022, causing Türkiye rank as the 34th most unequal country in the world (World Inequality Database, 2022). According to Dabla-Norris et al. (2016), income inequality is an important factor affecting growth and sustainability. The proposition put forth by the researchers posits that augmenting the proportion of income allocated to the lowest 20% of the population will result in a corresponding upsurge in economic growth. They also state that high level of income inequality leads to macroeconomic instability and prepares the ground for financial crises. Considering that in addition to economic problems, income inequality causes an increase in crime rates, corruption, misuse of resources and nepotism, it is clear that this problem will have social and political outcomes. Therefore, establishing income equality in the economy as well as stable economic growth is vital for a healthy society.

A substantial body of literature exists on the determinants of income equality, encompassing economic, financial, and social factors. When the current literature is examined, it is seen that the variable whose relationship with income equality is most investigated is economic growth, especially in developing countries (Digdowiseiso, 2009; Shahbaz, 2010; Baloch et al., 2018). However, as in the case of Türkiye, it is obvious that economic growth does not guarantee income equality. Conversely, empirical investigations that consider the era following 1990, during which financialization surged worldwide, primarily assess the effect of financialization on income inequality within advanced economies (Lin & Tomaskovic-Devey, 2013; Van Arnum & Naples, 2013; Alvarez, 2015). In these studies, it has been concluded that financialization reduces the income share of the working class and increases the share of the financial sector in the economy which eventually worsens income inequality. With the understanding that financialization alone cannot contribute to all segments of society, the concept of “financial inclusion”, which means a kind of “financial equality”, entered the literature. Studies investigating the link between income inequality, which has become a global problem, and financial inclusion began to enter the literature in the late 2000s. The first empirical findings that income equality can be achieved with “financial inclusion”, which means the spread of financialization to all segments of society, is based on the study of Kim (2016). The researcher presented findings that financial inclusion can improve income equality, especially in low-income and high-vulnerability countries. The relationship between the two variables has recently started to be investigated for Türkiye as well. In the study conducted by Çalış & Gökçeli (2022), it was revealed that financial inclusion reduces income inequality. In another study on Türkiye, no clear relationship was found between the two variables (Takmaz et al., 2022). The limited number of studies is not sufficient to reveal the effect of financial inclusion on income inequality in Türkiye. This study, which empirically investigates this effect, aims to fill an important gap in the literature. Revealing the relationship between income inequality and financial inclusion, which is one of the most important problems of Türkiye, will offer important implications for both policy makers and decision makers of the financial sector.

2. THEORETICAL AND EMPIRICAL LITERATURE

According to Leeladhar (2005), who was one of the first to define the concept, financial inclusion is the provision of banking services to all segments of society, including disadvantaged and low-income groups, at reasonable costs. In other words, it is to officially include the unbanked in the financial system. On the other hand, income inequality is the situation where the income created in the economy and the resources owned are not distributed equally among the citizens of the country (Atkinson, 1997). According to Brune et al. (2011), by increasing their saving ability individuals can get stronger against financial shocks, continue their consumption, and invest in their health and education. Therefore, access to financial resources has the potential to improve income distribution by eliminating poverty. In the theoretical model built by Kling et al. (2022) to investigate the relationship between financial inclusion and income inequality, it was tested whether individuals’ increased access to credit leads to education investments, thus more human capital and higher income. According to the result of the theoretical estimation, only informal loans can create an income increase for low-income individuals. Therefore, it can be argued that certain financial products may not be efficacious in mitigating the issue of income inequality.

According to intensive and extensive margin theories, the relationship between financial participation and income inequality may vary depending on whether economic agents have been involved in the financial system before. The intensive margin theory predicts that wealthy individuals and powerful companies that already have access to financial instruments can increase their financial resources more and affect income inequality (Chipote et al., 2014). On the other hand, the extensive margin theory argues that income inequality can be eliminated by incorporating those who have not had access to financial resources before into the financial system (Chiwira et al., 2016).

Greenwood & Jovanovic (1990) posit that the impact of financial inclusion on income inequality is non-linear, which contradicts the margin theories. Researchers have proposed the presence of a Kuznets-type correlation between financial inclusion and economic development. They argue that financial inclusion can solely benefit the entire society during the advanced stages of economic development, as individuals with elevated income levels are the only ones capable of accessing financial services during the initial stages of economic development. However, this view can be easily refuted when the United States of America, which has a high income inequality since the 1970s, is taken into account (Piketty & Saez, 2003). According to Menyelim et al. (2021), in the OECD report of 2008, income inequality deepened in developed countries, which are well above the Kuznet curve, causing the margin theory to be questioned.

Upon analyzing empirical studies that assess the link between financial inclusion and income inequality, it becomes apparent that diverse financial inclusion variables are employed for different countries and groups of countries. Agyemang-Badu et al. (2018) conducted an analysis of 48 African countries and developed a financial inclusion index that incorporates specific characteristics of each country. The fixed-effect panel regression estimator was utilized by researchers to gauge the interactions between financial inclusion, poverty, and income inequality. The research revealed that financial inclusion is inversely related to both poverty and income inequality in African countries. Khan et al. (2022) conducted a study utilizing the multiple regression method to examine the relationship between financial inclusion and poverty, income inequality, and financial stability in 54 African nations. The research indicates that financial inclusion is a vital mechanism for alleviating poverty and income inequality in these states.

The study conducted by Kim (2016) aims to evaluate the potential impact of financial inclusion, specifically in terms of financial accessibility, on the reduction of income inequality for 40 countries consisting of OECD and European Union (EU) members. The first outcome of the analysis is that the income inequality exerts a detrimental impact on the growth. There exists a robust inverse link between income inequality and income growth, particularly in nations with lower income levels. Furthermore, the effect of income inequality on diminishing economic growth is more pronounced in nations with high levels of fragility. Secondly, progressivity does not appear to be a significant determinant in mitigating income inequality in nations with low-income or those with high levels of fragility. Moreover, Kim (2016) argues that financial inclusion may strengthen the relationship between economic growth and income equality in general. Thus, it transforms the previously negative relationship between income inequality and economic growth into a positive relationship. Thus, the formerly adverse association between income inequality and income growth is now transformed into a positive as a direct result of this. The above-mentioned trend is more prevalent in countries with high levels of vulnerability as opposed to countries with low levels of vulnerability. Le et al. (2019), in their research on 22 transition countries, revealed that financial inclusion reduces income inequality.

Menyelim et al. (2021) examined the potential for financial inclusion in the context of Kuznets Curve to mitigate the negative effects of income inequality on the economic growth of 48 Sub-Saharan African countries. The results of the study show that promoting financial inclusion can have a short-term effect in reducing income inequality. The study revealed that inclusive financial access has a positive net effect in reducing the impact of income inequality on economic growth.

Neaime and Gaysset (2018), who tested the effect of financial inclusion on income inequality, poverty and financial stability for MENA countries, revealed that financial inclusion reduces income inequality, while population size and inflation increase income inequality. Omar & Inaba (2020) measured the impact of financial inclusion on poverty and income inequality in 116 developing countries by devising a unique financial inclusion index. The results of the empirical study demonstrated conclusively that financial inclusion lessens poverty and income inequality in emerging states. Similarly, Ouechtati (2020), who analyzed for developing countries, demonstrated that income inequality can be reduced if a high bank penetration rate and inclusive loans are made available to low-income individuals. Park & Mercado (2018) conducted a study utilizing a sample of 176 countries, which included 37 emerging Asian economies. The researchers developed a financial inclusion indicator and determined that financial inclusion has a significant impact on reducing poverty and income inequality across the entire sample. However, no significant relationship was found between these variables in developing Asian countries. Çalış & Gökçeli (2022), who analyzed Türkiye with the VAR approach, found an inverse relationship between financial inclusion and income inequality. Researchers who questioned the direction of causality between the variables revealed that financial participation is the granger cause of lower income inequality.

The present literature review reveals a notable void in empirical research pertaining to the correlation between financial inclusion and income inequality in nations with varying degrees of development. The existing body of literature on Türkiye appears to be limited.

3. EMPIRICAL ANALYSIS

This research explores the relationship between financial inclusion and income inequality from 1980 to 2021 in the case of Türkiye. The empirical part of the research commences with Model Specification. This section, titled "Empirical Analysis," provides an account of the theoretical underpinnings of the empirical methods employed as well as the results obtained through these methods. The study commences with examining the stationarities of the series through the application of the Augmented Dickey-Fuller (1979), Phillips-Perron (1988), and Zivot-Andrews (1992) techniques to ascertain the appropriate model specification. After determining the optimal lag-length, The Johansen cointegration approach is used in order to calculate the number of cointegrating relationships. This serves the purpose of confirming the existence of a long-term connection between the variables. The long-term coefficients are subsequently computed through the employment of FMOLS, DOLS, and CCR estimation techniques.

3.1. ECONOMETRIC METHODOLOGY

3.1.1. UNIT ROOT TESTS

Time series data mostly exhibit trending behavior, which is known as a scholastic trend; to put it another way, there may be an issue with the lack of stationarity. In order to acquire unbiased findings, it is vital to eliminate this kind of trending behavior. Therefore, determining the stationary level of a series is a prerequisite for time series analysis. The results of this unit root test are critical in choosing appropriate parameter estimation methods. The literature on econometrics has a variety of different unit root tests. This research makes use of the Augmented Dicky- Fuller (ADF) unit root test and the Philips Perron (PP) unit root test, both of which are used often in the applied economics.

The expression for ADF is presented in Equation (1) with a constant term, and in Equation (2) with both a

$$\Delta Y_t = \gamma_0 + \gamma_1 Y_{t-1} + \sum_{i=1}^{n} \alpha_i \Delta Y_{t-i} + \mu_t \tag{1}$$

$$\Delta Y_t = \gamma_0 + \gamma_1 t + \gamma_2 Y_{t-1} + \sum_{i=1}^{n} \alpha_i \Delta Y_{t-i} + \mu_t \tag{2}$$

$$H_0: \gamma_1 = 0 \; (unit\ root) \quad H_1: |\gamma_1| < 0 \quad ; \quad H_0: \gamma_2 = 0 \; (unit\ root) \quad H_1: |\gamma_2| < 0$$

constant term and a trend component.

Using ADF regression, which is represented by equations (1) and (2), the existence of a unit root of is investigated. The terms "t" and "n" are the number of time and the number of lags respectively. ΔY_{t-i} stands for the first difference of the series with "n" lag. As it is expressed in Eq. (2) "the null hypothesis of series has a unit root" is examined.

PP unit root test equation can be expressed as $\Delta Y_t = \vartheta_0 + \gamma Y_{t-1} + \alpha_t$ and hypothesis tested as H_0: $\gamma = 0$ (*unit root*); H_0:$\gamma < 1$. By directly altering the test statistics $t_{\gamma=0}$, the PP tests ensure that any potential serial correlation as well as heteroscedasticity in the errors at of the test regression are eliminated.

3.1.2. COINTEGRATION

If the two series in the cointegration system exhibit similar behaviors or trends, there is a non-random link, or a common stochastic trend. Long-term relationships exist between series exhibiting similar stochastic trend behavior. In this direction, the concept of a long-term relationship, which implies the existence of equilibrium between variables, can be tested by the presence of a common stochastic trend. If it is determined that the two integrated series share a common stochastic trend, we can say that they are cointegrated. In this research the existence of cointegration tested by Johansen tests for cointegration

$$Y_t = \delta D_t + \phi_1 Y_{t-1} + \cdots + \phi_p Y_{t-p} + \varepsilon_t \tag{3}$$

The basis of Johansen's methodology is the vector autoregression (VAR) of order p given by

y_t is a (nx1) vector of variables, which is typically represented as I(1) that stands for "integration of order one ε_t and is a (nx1) vector represent residual called innovations vector."

If $det(\mathrm{I}_n - \phi_1 Z - \cdots - \phi_p Z^p) = 0$ all of the model's roots are located outside the unit circle, then the VAR(p) model is stable. Considering the VECM we can rewrite the Johansen's cointegration equation is spec-

$$\Delta Y_t = \Gamma_1 D_t + \Pi y_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j} + \varepsilon_t \tag{4}$$

ified in Eq.(4)

where,

- D_t is a vector comprising of deterministic variables, which may include constants, trends, and/or seasonal dummy variables.
- $\Gamma_j = -I + \sum_{j=1}^{p-1} \phi_j$ are mxm matrices;
- The long-run impact matrix is represented by the equation Π= γA, where A and γ are matrices of dimensions (m×k), and ε_t represents the errors.
- All roots of $det(\mathrm{I}_n - \sum_{j=1}^{p-1} \Gamma_j B^j)$ are located outside the unit circle.

In this framework, cointegration happens when Π has reduced rank. This is the basis of the test: By checking

$$J_{trace} = -T \sum_{i=r+1}^{n} \ln(1 - \widehat{\lambda_i}) \tag{5}$$

$$J_{max} = -T \ln(ln(1 - \widehat{\lambda_{r+1}})) \tag{6}$$

the rank of Π, we can determine if the system is cointegrated. Equations (5) and (6) depict the trace and maximum eigenvalue test, as proposed by Johansen (Hjalmarsson and Österholm, 2007).

3.1.3. ESTIMATION METHODS - FMOLS, DOLS AND CCR

FMOLS, DOLS and CCR techniques are all capable of producing valid findings even with very small sample sizes. Phillips & Hansen (1990) introduced the FMOLS, which is an instrumental variable estimate with non-stationary regressors. FMOLS endeavors to address the endogeneity issue by utilizing kernel estimators of the parameter in question. Furthermore, the FMOLS method utilizes the error term's covariance matrix to tackle issues arising from the correlations between cointegration equations and stochastic processes in the long-term.

Thus, the FMOLS initially modifies the variables and estimates to eradicate the extant nuisance parameters. The FMOLS estimators rectify the OLS estimation method's inadequacy so that it is possible to determine the optimum value for cointegrating regressions. The FMOLS estimator uses an assumption that is asymptotically biased while yet being strictly exogenous (Chen ve Huang, 2013). The FMOLS estimator can be expressed in the following way:

$$y_t^* = y_{t-1} \overline{w}_{12} \Omega_{22}^{-1} u_{2t} \tag{7}$$

the biased correlation term can be express as $\gamma_{12} = \gamma_{12} - \overline{w}_{12} \Omega_{22}^{-1} \xi_{2t}$ (8)

The long-run covariance coefficients Ω and ξ are computed utilizing the residuals $u_t = (u_{1t} u_{2t}')'$.

The equation

(9) can be utilized to conduct FMOLS estimation in this instance.

$$\widehat{\theta_{FMOLS}} = \begin{bmatrix}\alpha\\ \beta\end{bmatrix} = \left(\sum_{i=1}^{T} s_t\ s_t'\right)^{-1} \left(\sum_{i=1}^{T} s_t\ y_t^* - T\begin{bmatrix}\gamma_{12}\\ 0\end{bmatrix}\right) \tag{9}$$

here $(s_t = x_t' d_t')'$

$$\widehat{\theta_{FMOLS}} = \left(\left(\sum_{i=1}^{T} x_t'\ d_t'\right)'\left(\sum_{i=1}^{T} x_t'\ d_t'\right)^{-1}\ y_t^* - T\begin{bmatrix}\gamma_{12}\\ 0\end{bmatrix}\right) \tag{10}$$

In Equation (10), the terms y_t* and γ_{12} correspond to the correction terms for endogeneity and serial correlation, respectively. This asymptotically unbiased estimator has a completely efficient mix-normal asymptotic distribution, allowing for the use of asymptotic chi-square inference in traditional Wald test (Adom et al., 2015).

Another approach, referred to as CCR, was presented by Park (1992). The statistical theory behind this method is basically the same as that which underpins the Phillips & Hansen (1990). The main difference between the two models is that CCR is only concerned with the transformation of data, whereas FMOLS is concerned with the transformation of both data and parameters. CCR is a single equation regression that is capable of applying multivariate regression without requiring any modifications or sacrificing its effectiveness (Park, 1992). In accordance with the findings of Adom et al. (2015), the CCR estimator is shown as in Equation (11)

The third estimation method employed in this research is DOLS which is developed by Stock and Watson (1993). They suggested that it would be advantageous to include the lag and lead values of the independent

$$\widehat{\theta_{CCR}} = \left(\sum_{t=1}^{T} Z_t^* Z_t^{*1}\right)^{-1} \sum_{t=1}^{T} Z_t^* Z_t^* \tag{11}$$

variables in addition to the level values of the independent variable in order to rectify the deviation and endogeneity problem of OLS method. This DOLS (Dynamic Ordinary Least Squares) method is an effective estimator that that are resistant to autocorrelation and heteroscedasticity. In this regard, it is deemed superior to other estimators (Esteve & Requena, 2006).

The DOLS technique offers an estimator that is asymptotically efficient and eradicates the feedback effects present in the cointegration equation. The DOLS approach can be mathematically represented by the

$$y_t = x_t'\beta + d_{1t}\Psi_1 \sum_{j=q}^{r} \Delta\ x_{t+j}'\ \delta + u_{1t} \tag{12}$$

Equation (12).

Here, “q” and “r” allow the difference of the explanatory variables, which allows to eradicate the long-run correlation between the error terms. The DOLS methodology’s estimation procedure yields parameter estimates that possess an asymptotic distribution, similar to the FMOLS and CCR approaches. The DOLS approach incorporates the first difference of the independent variables, thereby enabling the inclusion of time lags in the estimation process, while the CCR approach enables the asymptotic execution of the $\chi 2$ test.

3.1.4. MODEL SPECIFICATION AND FINDINGS

In the model, income inequality is defined as the dependent variable, while financial inclusion, education, income, inflation and urban population are defined as independent variables. The variables that are used in the model, along with their abbreviations and proxies for those variables, are listed in Table 1.

Table 1. Variables

Variables	Abbr.	Proxy	Source
Income inequality	GINI	GINI coefficient	World Bank
Financial inclusion	FDI	Financial development index	IMF
Education	ED	School enrollment, tertiary (% gross)	World Bank
Income	GDP	GDP per capita (constant LCU)	World Bank
Inflation	CPI	Inflation, consumer prices (annual %)	World Bank
Urban population	POP	Urban population (% of total population)	World Bank

Source: World Bank (2023); IMF (2023)

A significant corpus of scholarly literature exists that provides estimations of the impact of financial inclusion on economic activity and inequality. In academic literature, the most commonly used indicators for measuring financial development are the private credit to GDP ratio and the stock market capitalization to GDP ratio. These indicators are often utilized as proxies for financial development in empirical research. Nevertheless, the proxies utilized fail to consider the intricate and diverse nature of financial inclusion.

As per the Staff Discussion Paper (Sahay et al., 2015), the "financial development index" offers a comprehensive evaluation of the development of financial institutions and markets based on their depth, access and efficiency. Depth is assessed by measuring size and liquidity, access by evaluating the accessibility of financial services to individuals and businesses, efficiency by examining the capability of institutions to deliver financial services at a sustainable cost, and level of activity by analyzing the degree of participation in financial markets.

The following Eq. 13 depicts the functional formulation of the model that will be employed in empirical analysis.

$$Income\ Inequality = f(Financial\ Inclusion, Education, Income, Inflation, Polulation) \quad (13)$$

$$GINI = f(FIN, \quad ED, \quad GDP, \quad CPI, \quad POP)$$

Equation 1 is the functional representation of the model. We can express our model in the form of cointegration regression as in Eq. 14

$$lnGINI_t = a + \beta_1 lnFIN_t + \beta_2 lnED_t + \beta_3 lnGDP_t + \beta_4 lnCPI_t + \beta_5 lnPOP_t + \varepsilon_t - \lambda ECT_{t-1} \quad (14)$$

Where, is the error term, λ is the speed of adjustment parameter, and is the lagged error correction term. The cointegrating vectors are incorporated into the model as the coefficients on the independent variables, β_1, ..., β_5. These coefficients represent the long-run relationship between the variables. The error correction term, λECT_{t-1}, captures the short-run dynamics that drive the adjustment process towards the long-run equilibrium. The speed of adjustment parameter, λ, represents the extent to which the dependent variable adjusts to deviations from the long-run equilibrium in the previous period.

Table 2 displays the results of ADF and PP, which test the null hypothesis of "series contains a unit root". The output of the report indicates that all series contain unit roots in levels, but exhibit stationary in the first difference. Thus, GINI, FIN, GDP, CPI, ED, and POP are integrated order of one $Yt \sim I(1)$.

Table 2. ADF and PP unit root test

H0: series contain unit root	ADF				PP			
	I(0)		I(1)		I(0)		I(1)	
Variable	Stat.	Prob.	Stat.	Prob.	Stat.	Prob.	Stat.	Prob.
GINI	-1.168	0.687	-5.004***	0.000	-1.226	0.662	-5.067***	0.000
FIN	-0.990	0.757	-7.373***	0.000	-0.977	0.7617	-7.674***	0.000
GDP	2.166	0.999	-5.072***	0.000	2.769	1.0000	-5.011***	0.000
CPI	-2.061	0.261	-7.479***	0.000	-2.158	0.2219	-7.717***	0.000
ED	0.165	0.970	-6.558***	0.000	0.222	0.9735	-6.559***	0.000
POP	-1.908	0.328	-3.995**	0.0014	-1.701	0.4306	-5.233***	0.000

Source: Authors' calculation. *** p<.01, ** p<.05, * p<.1

Unit root tests that do not account for structural breaks may produce biased results. By considering the numerous internal and external disruptions experienced by the Turkish economy during the review period, it would be appropriate to employ the unit root tests allowing for structural break. For this purpose, Zivot-Andrews (1992) unit-root test which is considered structural break is employed to robust the output of ADF and PP Test. The outcomes displayed in Table 3 confirms the ADF and PP findings that all variable are $Y_t \sim I(1)$.

Table 3. ZA Test Results

H0: series contain unit root	I(0)	I(1)	Decision
Variable	Stat.	Stat.	
GINI	-3.111	-6.239***	I(1)
FIN	-4.597	-7.581***	I(1)
GDP	-2.153	-6.088***	I(1)
CPI	-4.550	-8.770***	I(1)
ED	-4.458	-9.152***	I(1)
POP	-4.012	-6.979***	I(1)

Source: Authors' calculation. *** p<.01: (-5.34), ** p<.05: (-4.80), * p<.1: (-4.58)

The selection of the lag length is a fundamental issue in applied time series analysis. In time series analysis, lag refers to the number of time periods between the current observation and a past observation. Rarely is there an immediate response or interaction between dependent variable, Y, and independent variable, X. Most of the time, it takes a considerable period of time before Y responds to X. In general, having an excessive number of lags causes the standard errors of coefficient estimates to be inflated, which in turn implies an increase in the amount of error associated with the prediction. The model may become overly complex, leading to overfitting and poor generalization to new data. If the lag length is too short, the model may not capture all the relevant information from the past, leading to poor predictions. Therefore, selecting an appropriate lag length is crucial for building accurate and reliable time series models.

Table 4 summarizes the findings of the variety of information criterions that determined the optimal lag length. BIC indicates that the optimal lag length is 1, while LL, FPE, AIC, and HQIC all determined as 4. As a result, it has been concluded that lag length of four is the optimal choice.

Table 4. Optimal lag-length selection

lag	LL	LR	df	p	FPE	AIC	HQIC	SBIC
0	233.527	0.000			2.5e-13	-11.975	-11.883	-11.717
1	555.807	644.560	36	0.000	7.4e-20	-27.043	-26.398	-25.2325*
2	615.811	120.010	36	0.000	2.4e-20	-28.306	-27.110	-24.9445
3	652.395	73.169	36	0.000	3.6e-20	-28.337	-26.589	-23.4238
4	719.142	133.49*	36	0.000	1.9e-20*	-29.9549*	-27.655*	-23.4907

Source: Authors' calculation

The Johansen test is a statistical procedure utilized to ascertain the existence and quantity of cointegrating associations between the series. The trace statistic is one of two test statistics used in the Johansen test, the other being the maximum eigenvalue statistic. The choice between the two statistics depends on the number of variables and the order of integration of the variables under consideration. The trace statistic is more appropriate when the variables are integrated of the same order, while the maximum eigenvalue statistic is more appropriate when the variables are integrated of different orders. Therefore, trace statistics are employed to ascertain the existence and quantity of cointegrating relationships among the series. The outcomes of the test reported in Table 5.

Table 5. Johansen tests for cointegration

No. of CE(s)	parms	LL	Eigenvalue	Trace Statistic	%5 Critical Value
0	114	613.37185	.	211.5408	94.15
1	125	662.27554	0.92376	113.7334	68.52
2	134	685.14911	0.69997	67.9863	47.21
3	141	704.66283	0.64193	28.9588*	29.68
4	146	711.73162	0.31067	14.8212	15.41
5	149	716.50462	0.22214	5.2752	3.76

Source: Authors' calculation

The Johansen test posits a null hypothesis "there is no cointegration among the variables" against the alternative hypothesis of "there is at least one cointegrating vector". As the calculated trace statistic exceeds the critical values from the tables provided by Johansen, then the null hypothesis is rejected indicating the variables are cointegrated. As it is "3 cointegrating equations at the 0.05 level" indicated we can estimate the long-run parameters by employing FMOLS, CCR and DOLS cointegration regression techniques. The FMOLS, CCR and DOLS estimator yielded the following results, as shown in Table 6.

Table 6. Cointegration Regression Estimations

LnGINI	**FMOLS**	**CCR**	**DOLS**
LnFIN	0.0115*** (3.87)	0.0116*** (5.21)	0.0660*** (6.76)
LnED	-0.114*** (-35.24)	-0.0984*** (-34.01)	-0.114*** (-3.83)
LnGDP	0.116*** (16.84)	0.115*** (16.19)	0.163*** (4.67)
LnCPI	0.0205*** (40.21)	0.0210*** (43.65)	0.0297*** (4.13)
LnPOP	0.461*** (41.70)	0.342*** (49.12)	0.313*** (7.48)
_cons	1.034*** (13.61)	1.483*** (21.27)	1.106** (3.17)
R2	.799	.950	.982
N	41	41	41

Source: Authors' calculation. * $p<0.05$, ** $p<0.01$, *** $p<0.001$ - t statistics in parentheses

The outcomes of FMOLS, CCR, and DOLS estimators, respectively, are as follows

- The coefficient for financial inclusion exhibits a positive sign and is statistically significant at the 0.01 level. An increase of 1% in financial inclusion results in increase in income inequality of approximately 0.012%, 0.012%, and 0.07%, respectively.
- The coefficient for tertiary education exhibits a negative sign and is statistically significant at the 0.01 level. A 1% increase in tertiary education is associated with a decrease in income inequality of -0.114%, -0.098%, and -0.114%, respectively.
- The coefficient for GDP per capita exhibits a positive sign and is statistically significant at the 0.01 level. An increase of 1% in GDP per capita increases income inequality by 0.116%, 0.115%, and 0.163%, respectively.
- The coefficient for inflation, specifically consumer price inflation, exhibits a positive sign and is statistically

significant at the 0.01 level. An increase of 1% in inflation results in a distortion of income inequality by 0.0205%, 0.0210%, and 0.0297%, respectively.

• The coefficient for the urban population displays a positive sign and exhibits statistical significance at the 0.01 level. Moreover, a 1% rise in urban population leads to distortions in income inequality of 0.406%, 0.342%, and 0.313%, respectively.

The CCR and DOLS estimators are utilized to enhance the robustness of the FMOLS estimation outcomes. Based on the results obtained from the three estimation techniques, it has been concluded that the parameter estimates exhibit a high degree of proximity and that the direction of the relationship between the dependent and independent variables is consistent. Table 7 presents a summary of the results obtained from three estimations.

Table 7. Summary of estimations

Income inequality –Gini-	**FMOLS**		**CCR**		**DOLS**		**Impact on Gini**
	Sign	Signific.	Sign	Signific.	Sign	Signific.	
Financial inclusion	+	√	+	√	+	√	Distort
Education-tertiary	-	√	-	√	-	√	Improve
Income-GDP per capita	+	√	+	√	+	√	Distort
Inflation-CPI	+	√	+	√	+	√	Distort
Urban Population	+	√	+	√	+	√	Distort

Source: Authors' evaluation

3.1.5. MODEL VALIDATION

The LM test, also referred to as the Lagrange-multiplier test, is a diagnostic tool employed to ascertain the presence of autocorrelation in the residuals of a regression model. Autocorrelation is a statistical phenomenon that occurs when the residuals of a model display a correlation with one another, indicating the existence of an identifiable pattern in the model's errors that has not been appropriately addressed. The test's null hypothesis posits the absence of autocorrelation at the specified lag order. The initial step in conducting the LM test involves estimating the regression model and acquiring the residuals. The subsequent step involves conducting a regression analysis of the squared residuals against the lagged values of the squared residuals, with consideration of a predetermined lag order. The test statistic denoted as LM is derived by multiplying the R-squared value obtained from the regression analysis with the sample size. The null hypothesis assumes the absence of autocorrelation and the distribution of the test statistic follows chi-squared distribution.

Table 8. LM Test

lag	$\chi2$	df	Prob > $\chi2$
1	36.5457	36	0.44332
2	27.4017	36	0.84783
3	46.6648	36	0.10982
4	30.5180	36	0.72655

Source: Authors' calculation

Upon scrutiny of Table 8, it is evident that the p-values surpass 0.05 for every lag duration. The null hypothesis stating the absence of autocorrelation cannot be rejected.

Normality tests are utilized as diagnostic instruments to evaluate whether a given dataset conforms to a normal distribution, which is a crucial assumption of our statistical model. Non-normal distribution of data can result in biased parameter estimations, imprecise p-values, and erroneous conclusions. Consequently, it is crucial to verify the normality assumption prior to utilizing statistical techniques that are dependent on it.

The Jarque-Bera (JB) test is a statistical method utilized to ascertain whether a given dataset conforms to a normal distribution. The JB test relies on the sample skewness and kurtosis as statistical measures of the extent to which the sample distribution deviates from a normal distribution. Skewness quantifies the degree of asymmetry of the distribution, while kurtosis measures the heaviness of the tails of the distribution. In a Gaussian distribution, the measures of skewness and kurtosis are both equal to zero.

Table 9. Normality

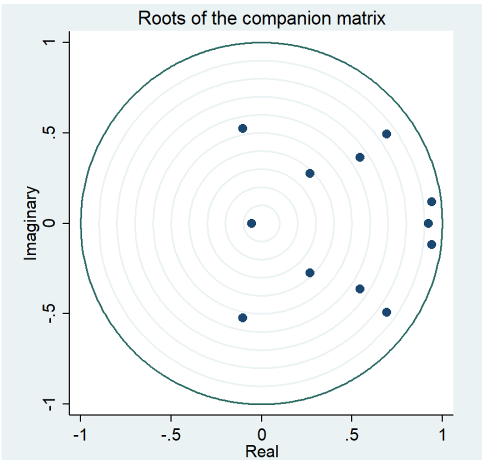
	Jarque-Bera test		Skewness test		Kurtosis test	
Eqs.	χ2	Prob > χ2	χ2	Prob > χ2	χ2	Prob > χ2
LnGINI	2.024	0.36350	1.952	0.16234	0.072	0.78889
LnFIN	2.848	0.24073	2.842	0.09186	0.007	0.93497
LnED	1.773	0.41209	1.246	0.26424	0.527	0.46803
LnGDP	6.565	0.03753	0.681	0.40920	5.884	0.01528
LnCPI	1.014	0.60221	0.082	0.77493	0.933	0.33420
LnPOP	3.188	0.20315	2.023	0.15491	1.164	0.28054
ALL	17.412	0.13473	8.826	0.18359	8.586	0.19822

Source: Authors' calculation

Table 9 presents the results of the normality test, which consisted of JB-Skewness and Kurtosis. Given that the p-value corresponding to the test statistic is greater than 0.05, we are unable to reject the null hypothesis. Thus, we can infer that there is insufficient evidence to support the claim that the sample data does not conform to a normal distribution.

Table 10. Eigenvalue stability

Eigenvalue		Modulus
.9408056	+ .118923i	.948292
.9408056	- .118923i	.948292
.9231298		.92313
.6934538	+ .4937039i	.851247
.6934538	- .4937039i	.851247
.5438024	+ .3648882i	.654877
.5438024	- .3648882i	.654877
.1045889	+ .5235808i	.533925
.1045889	- .5235808i	.533925
.268261	+ .2745281i	.383835
.268261	- .2745281i	.383835
.053905		.053905



Source: Authors' calculation

If a given matrix A is characterized by eigenvalue stability, then its solutions exhibit stability even when subjected to minor perturbations. The concept of eigenvalue holds significant importance in the domain of matrices, as it plays a crucial role in ensuring the precision and stability of numerical computations. As demonstrated in Table 10, all of the roots of the companion matrix are situated within the unit circle. The relationship between the roots of a polynomial lying inside the unit circle and the absolute value of the eigenvalues of the companion matrix being less than one is a result of the principles of linear systems and the spectral theorem. The stability and convergence of specific numerical methods and systems, such as difference equations and dynamical systems, are significantly impacted by this phenomenon. The diagnostic tests conducted on the model's results suggest that the model is valid.

4. DISCUSSION AND CONCLUSION

The issue of income inequality has gained significant attention in both developed and developing nations, primarily due to the emergence of financialization trends resulting from globalization. The issue of income distribution remains a crucial topic on the political and economic agenda, given the rise in global trade volume and production output. Despite being limited in scope, academic research has demonstrated a keen interest in investigating the effects of financial inclusion on the equitable distribution of income. In this regards, this study investigated the impact of financial inclusion on income inequality in Türkiye, contributing to the very limited literature. Financial inclusion indicators in the literature mainly cover the number of ATMs per capita, the number of bank branches and the share of credits in GDP. Unlike the literature, in this study, indicators including the depth, access and effectiveness of financial markets and institutions were incorporated in the analysis by using the IMF's financial development index. In addition to financial inclusion, variables such as education, income, inflation and urban population, which are thought to affect income inequality, were also added to the model as control variables.

According to the results of the analysis, the increase in financial inclusion creates a distorting effect on income equality in Türkiye. This result, which seems to contradict the findings of previous studies, may be partly due to measurement differences. While the representative variable of financial inclusion was narrowly defined in previous studies, the financial development index employed as a proxy for financial inclusion in this study is a much broader indicator that includes depth, access and efficiency of stock exchange and other debt markets. Considering the studies in which stock exchange markets deteriorate income inequality or do not contribute to income equality, especially in developing countries (Das & Mohapatra, 2003; Seven & Coskun, 2016; Blau, 2018), the finding obtained from this study is not surprising. This result can also be explained by the economic structure of Turkey. Considering the growth composition, it can be easily said that Türkiye is a consumption-based economy. As orthodox economists emphasize, the use of credit facilities by individuals for consumption rather than investment opportunities can also contribute to income inequality.

An important variable that has been agreed upon in influencing income distribution is education. Consistent with previous studies, this study confirms that education can help to cope with income inequality in Türkiye. Education enables individuals to acquire the knowledge and skills necessary to find better paying jobs. Therefore, the contribution of education to income equality is an expected result for this study. Another variable whose effect on income inequality was measured in the study is GDP per capita. There is no consensus in the literature on how GDP per capita affects income distribution. In this study, findings reveal that increase in GDP per capita distorts income inequality. It has been determined that inflation is another important macroeconomic variable that distorts income equality in Türkiye. Similarly, the increase in the urban population aggravates income inequality in Türkiye.

To conclude, it is not possible to solve income inequality in any economy with just one variable. Income inequality is a complex problem with economic, social and institutional dimensions. Therefore, the approach to the solution of this problem should include a solution beyond just considering financial factors. The findings of the study provide important implications to solve income inequality in Türkiye. First of all, although there is no negative relationship between the financial inclusion indicator and income inequality, inclusion of low-income groups in all areas of the financial world by increasing their income will likely improve income inequality. Another macroeconomic indicator that will contribute to the solution of income inequality is inflation. Although inflation causes chronic problems in many areas of the economy, its greatest impact is reducing the purchasing power of individuals. Considering that the low-income group is the most affected by inflation, it can be predicted that ensuring price stability in the economy will improve income equality. Based on the findings of the study, policies aimed at raising the level of education and bringing more individuals into education will also ensure a more equitable distribution of income in Türkiye.

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