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ORIGINAL ARTICLE

The handicap for enhanced solidarity across advanced economies: The greater the economic openness higher the unequal distribution of income

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Abstract

Perspectives on degree of economic openness have been discussed by different paradigms resulting in concluding remarks which differ widely from each other. However, the existing literature has passed over the importance of the relationship between economic openness and income distribution in relation to the bargaining framework and thus its effect on social solidarity. In this paper, the rationale is based on the argument that an increase in the degree of economic openness has a significant impact on the distribution of income through the channels of trade and financial account. The empirical results show that a more open economic environment leads to an unequal distribution of income, both reducing labor's share and increasing income inequality, through the collapse of bargaining power of labor in favor of capital across advanced economies over the 1996–2014 period. However, the results also suggest that the trade channel is a more effective influence on income distribution than the finance channel. Additionally, other globalization measures, including economic, social and political factors, also have negative effects on income distribution

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by way of reducing the bargaining power of labor, which results in the collapse of social solidarity and cohesion over time.

KEYWORDS
bargaining framework, economic openness, globalization, income distribution, social solidarity

JEL classification:
D33, F4, F60

1 | INTRODUCTION

Recent decades have seen the declining bargaining power of labor, institutional breakdown, rising social conflicts and increasing income gap between poor and rich. These facts are linked to each other through different channels but they have as a whole eroded the social cohesion by reducing the solidarity between *homo economicus* and *homo socius*. This paper focuses on the relationship of two major phenomena, namely economic openness and income inequality, which have undermined social cohesion in many different ways and thus inhibit enhanced solidarity within advanced countries where the capitalist ingredients are well-organized and well-developed in contrast to economies from developing and emerging country groups. Investigation of this relationship may give important clues to the changing dynamics of social solidarity in advanced economies and may also provide important policy components for future generations to avoid such factors which affect social cohesion. Thus, the major focus of this study is on the two basic issues ofeconomic openness, proxied by several macroeconomic indicators, and income inequality, proxied by the labor share of income and the Gini coefficient.[1]

Economic openness can be assumed as one of the leading components of globalization. Therefore, analysis of the changing dynamics of social solidarity within the context of the openness–income distribution nexus, the globalization factor can also be added into the theoretical considerations. In particular, if it is assumed that economic freedom has a significant impact on the degree of social solidarity through increasing the level of income inequality, the sub-components of the economic channel of the globalization phenomenon should be examined with caution. For instance, Young and Tackett (2018: 18) argue that the overall perception of globalization is based on the idea that it favors firms/corporations and their profits at the expense of workers and their wages. The reasons for this fact can be either institutional failures or economic instabilities rooted in the unfettered globalization of capital. According to Thomas Piketty, “if we want to have a regulated globalization that benefits the majority of people, then we need closer political and fiscal cooperation.”[2] Therefore, whilst investigation of the social components of globalization needs much attention to understand the economic changes occurring in the societal environment, the

[1] The major reason why the study uses two variables in terms of income distribution is because the labor share of income provides a way to make class-based analysis whereas the Gini coefficient leads us to investigate the household-based differences in the distribution of total national income.

[2] Quoted in the European Magazine (12.12.2014), available at https://www.theeuropean-magazine.com/thomas-piketty-2/9351-thomas-piketty-on-globalizations-ills.

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economic factors in themselves are important to clarify the changes that emerge in social solidarity, especially in the case of the bargaining positions of capital and labor.

What drives the unequal distribution of income between capital and labor and has caused income inequality to increase in recent years across advanced countries when there is an unfettered openness in economic indicators? This paper attempts to answer this question by way of concentrating on the role of changing bargaining positions and thereby social relations in terms of solidarity among individuals. While the bulk of studies focus on household-based analysis in terms of the relationship between economic openness and income inequality, there are also other studies focusing on the importance of class ingredients of this nexus in which they are examined through the way of looking at differing bargaining power over time. Asteriou, Dimelis, and Moudatsou (2014) empirically investigate the reasons behind income inequality in the presence of an increasing degree of globalization, measured with both trade and financial variables to determine the effects of each specific factor which as a whole may indirectly be influential on social cohesion. As pointed out by the authors (p. 598), "…trade openness exerts an equalizing effects, while financial globalization through FDI, capital account openness and stock market capitalization has been the driving force of inequality…". Additionally, Jayadev (2007) empirically presents a class-level analysis for the relationship between financial globalization, measured with capital account openness, and the labor share of income, and finds that a higher level of financial flows decreases the labor share of income through increasing the threat option of capital at the economy-wide level since getting returns from abroad becomes easier than before.

In this context, today, the distributional conflicts are more complex than in the past in the presence of increasing globalization of capital, which in turn leads to much more complicated economic relations among different income-owners. Therefore, a significant issue within these complicated economic and social relations arising from an increase in the level of globalization concerns differing bargaining positions of capital and labor and thereby indirectly the solidarity between social groups. In globalized economic relations, growing inequalities, and increasing competitive pressures, the distribution of national income is significantly changed as well. However, in such an economic environment, labor's share is firmly decreased due to the fact that increased capital mobility provides more opportunities for capital to get higher returns from abroad, thus boosted the bargaining position of capital to the detriment of labor. Figure 1 shows that the labor share has been shrinking in the advanced countries; on the other hand, Figure 2 shows that income inequality has increased over the same 1996–2014 period. In addition, Figure 3 shows that the economic openness measures, proxied by openness to trade and financial flows, skyrocketed in these years.

Although the reasons are numerous, the downward trend of labor's share is obvious for different countries. IMF (2017) highlighted that the labor share of national income was has been falling at the global level since the early 1990s, as well as its heterogeneity across countries, sectors, and skill groups. Some previous studies lead us to understand that the change in labor returns (i.e., in particular, the decline of labor share and the growing income inequality) coincides with a higher level of globalization movement following the 1980s across different countries (Chordokrak & Chintrakarn, 2011; Goldberg & Pavcnik, 2007; Wan, Lu, & Chen, 2007; Xue, Luo, & Li, 2014).

Figure 4 shows the medium-run causal linkages between the selected indicator for the bargaining power of labor which is the unemployment rate and labor's share.[3] Figure 5 depicts the changes in the unemployment rate and the Gini coefficient in case of medium-run causal

[3] The left-hand side of the graph shows the trade openness index and the right-hand side shows the financial openness index.

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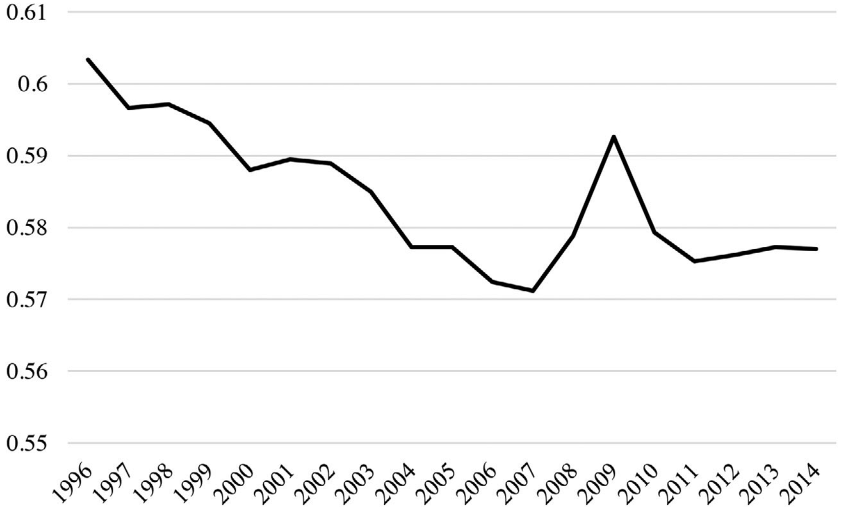

FIGURE 1 Trends in labor's share
Source: Penn World Tables 9.0

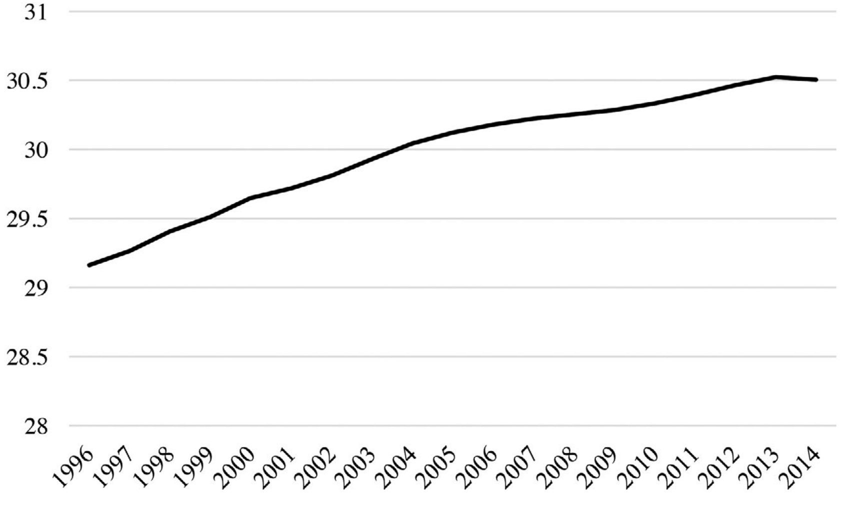

FIGURE 2 Trends in income inequality *Source*: Solt (2016)

linkage for the period between 1996 and 2014. The conceptual background of this paper is guided by the hypothesis that, given the other macroeconomic indicators, the change in the distribution of income is to a large extent affected by changes in the bargaining power of labor. In that sense, worker's bargaining power and thereby labor's share accruing in national income can be changed by the following three factors: the economy's overall health (e.g., inflation rate, current account balance, GDP per capita, exchange rates, technological change, human capital, etc.), financial motives (e.g., capital market liberalization, financial development, etc.), and the political context (e.g., political risks, democratic context of the country, etc.).

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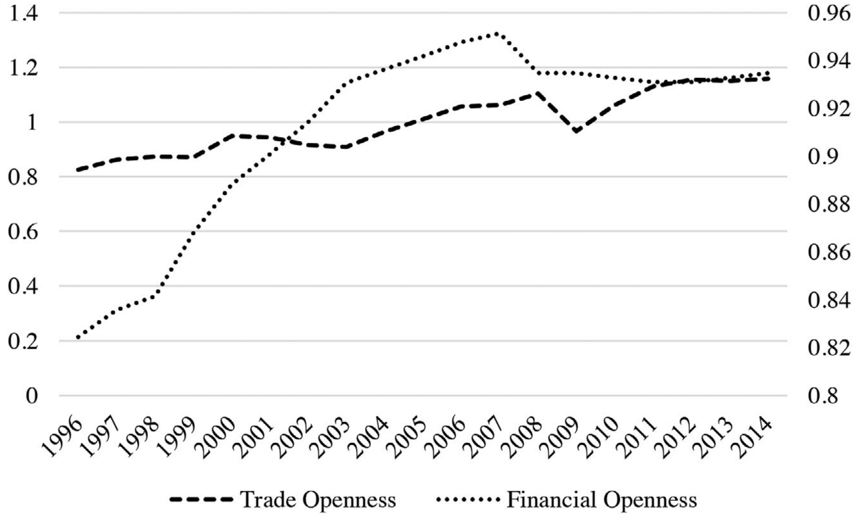

FIGURE 3 Trends in trade openness and financial openness
Source: World Bank, World Development Indicators Database; Aizenman, Chinn, and Ito (2008)

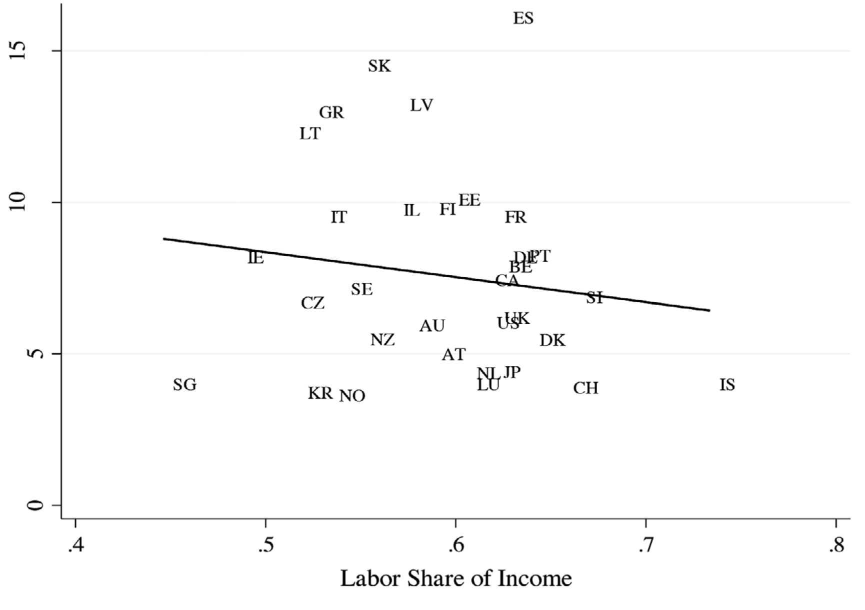

FIGURE 4 Unemployment rate and labor's share, 1996–2014
Source: World Bank, World Development Indicators Database; Penn World Tables 9.0

All else being constant, an increase in the overall health of the economy and institutional progress increases labor's share, whereas an increase in financial motives decreases labor's share by way of reducing employment opportunities. Detecting the main causes of unequal distribution of income pursuant to the changes in bargaining power is fundamental to constitute sound policy regimes that can provide enhanced social solidarity for a more equalized socio-economic and political environment.

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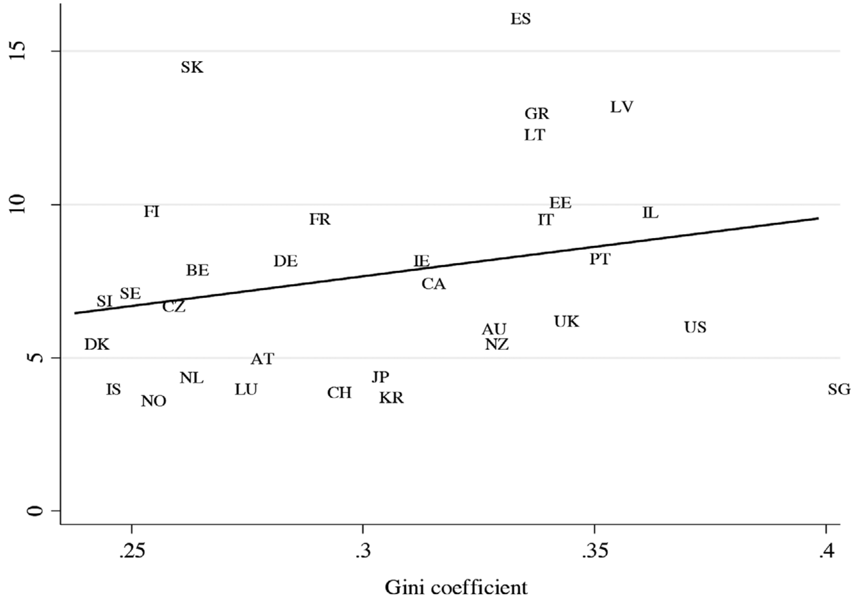

FIGURE 5 Unemployment rate and Gini coefficient, 1996–2014
Source: World Bank, World Development Indicators Database; Solt (2016)

In general, this article differs from previous studies in three aspects: First, it is focused on specific components of globalization, namely trade openness and financial openness, for the investigation of total income distribution through linking the causal pathways of bargaining power. Second, it concentrates on advanced economies, namely high-income countries where capitalist ingredients are well developed. Third, beyond empirical considerations of the effects of globalization measures discussed in relation to income inequality on the basis of household level, this study extends the current paradigm by incorporating the class-based dynamics through measuring the changes in income shares of labor and capital.

Additionally, this study seeks to contribute to the existing literature in different ways. First, it examines the importance of the role of bargaining power in case of income distribution and thus indirectly the differing societal formation. Differences in labor market policies, as well as in the macroeconomic policies affecting labor institutions across advanced countries can lead to the emergence of negative effects on bargaining positions. Second, it separately measures the effects of openness indicators on income distribution to understand their differential impacts in the case of class-based analysis. Third, it allows for the varieties for the methodological framework by way of considering both a fixed-effects model of country-level income distribution and lagged values of variables to account for the endogeneity issue.

In particular, a related question regards the major causes of a decline in labor's share since the late 1970s in advanced economies.[4] According to Dabla-Norris, Kochhar, Suphaphiphat, Ricka, and Tsounta (2015: 18–22), the drivers of income inequality are several and can potentially affect social cohesion: global trends, technological change, trade globalization, financial globalization, financial deepening, changes in labor market institutions, redistributive policies, and the level of education. However, none of these factors are linked to bargaining power. For instance, Blanchard

[4] This topic is further developed and discussed in the following section.

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and Giavazzi (2003) argue that looking at the changes in marginal productivities provides limited and misguided information on declining labor's share and thus labor movements (e.g., unionization) should be considered in the case of income distribution. Kristal (2010: 737) also extends this case by way of regarding the historical sequence of labor movements in advanced countries in three aspects: "global context of workers' bargaining power", "labor-affiliated political parties", and "expansion of the welfare state". Additionally, Jaumotte and Tytell (2007) deal with the effects of immigration flows and income shares accruing to labor. In this paper, all these potential factors are considered in lieu of the relationship with the globalization phenomenon and their effects on the income distribution through the bargaining channel, which negatively dampens the social cohesion.

In view of these factors provided in the existing literature, this article provides a rationale by analyzing the role of economic components of globalization in explaining the change in income distribution on the basis of differing bargaining positions and thus social class segregation, using a panel of 32 advanced economies over the 1996–2014 period. This article is organized as follows. Section 2 reviews the literature about the globalization–income distribution nexus, especially considering the bargaining power in relation to the decline in labor share of income and income inequality. Sections 3 and 4 elaborate on the dataset and the empirical methodology that is used in the analysis, respectively. Section 5 presents the results of the benchmark and sensitivity analyses. Section 6 concludes.

2 | CHANNELS OF INFLUENCE

The existing literature has thoroughly investigated the relationship between trade and financial globalization and income distribution (e.g., Asteriou et al., 2014; Bergh & Nilsson, 2010; Doan & Wan, 2017; Young & Tackett, 2018). However, only a few of them point to the importance of the role of bargaining power and thereby the social cohesion in searching for the links between these two indicators. In this paper, the globalization process is only addressed in economic terms. It is assumed that economic openness may affect class- and household-based distributional contexts on the basis of two channels: (a) openness in trade and (b) financial flows. The effects of these two channels on distributional/allocational issues are specified through various control variables that have been used by previous theoretical and empirical studies.

Dreher (2006: 3) notes that globalization is "a process that erodes national boundaries, integrates national economies, cultures, technologies, and governance and produces complex relations of mutual interdependence". Whereas this conceptual definition is to a large extent analogous with the previous explanations, the empirical considerations based on the same topic are erratic; and therefore an appropriate framework to present a congruent background for the relationship between globalization and income distribution has not been established yet. On the one hand, the proponents of more globalized economic relations argue that the increase in total national income over the post-1980 period substantially stemmed from an increase in openness to trade and growing financial flows, resulting in the reduction of global income inequality.[5] On the other hand, the opponents point out that the fruits of the globalization process are not equally distributed among groups of people, despite the fact that aggregate income increases both within and between countries over time. While the second point is most applicable for the case of

[5] However, Milanovic (2016) argues that within-country income equality is still very problematic across nations even though the global inequality gap has narrowed over time.

developing and emerging economies, the inequality gap has also widened through several channels in advanced economies. However, the key point is to determine the major factors driving income inequality in relation to the globalization phenomenon.

Following the work of Asteriou et al. (2014): 593–594), the factors that spur change in income inequality should be discussed through three different components: (a) the links to trade openness, (b) the links to financial globalization, and (c) other links with regard to macroeconomic determinants, capital-based technological change, governmental problems, educational attainment, and so forth. Some of these topics are discussed in what follows in detail.

2.1 | The logical background of openness–distribution nexus

Standard trade theory predicts that international trade can affect the within-country distribution of income through increased income for the owners of the abundant factors in production. According to this theory, there are several ways through which the trade channel of globalization can reduce global inequality on the basis of the Heckscher–Ohlin (H-O) theorem (factor endowment theory) and Stolper–Samuelson (S-S) theorem. The predictions of these models suggest that a higher degree of openness in the trade regime increases demand for unskilled labor in developing countries where the labor–capital ratio is relatively higher than that of developed countries. On the other hand, trade openness benefits skilled labor more in developed countries and thus leads to production of more capital-intensive goods where the capital–labor ratio is relatively higher than that of developing countries. For this reason, a wage differential would be narrowed in favor of unskilled labor in developing countries and hence the liberal arguments supporting more open trade regimes in developing countries would be at the forefront. All in all, whereas developing countries would tend to export unskilled-labor abundant goods, developed countries would tend to export skilled-labor abundant goods. This implies that the abolishment of trade barriers (or reductions in tariff and trade taxes) and a higher degree of trade integration would provide higher returns under liberalized trade regimes to the workers with the abundant factor. Therefore, it is supposed that international trade would reduce the labor share of income in advanced economies and would increase it for other countries (Giovannoni, 2014; Stockhammer, 2013).

However, the theoretical arguments and empirical observations were separated from each other due to restrictive and naïve assumptions of the neoclassical trade model in which the decline of labor share has been captured in developing countries over the post-1980 period. While the theoretical reasons are several, the general view on this decline has raised concerns. Until recently, traditional wisdom has argued that international trade tended to increase economic growth which was the key dimension of globalization. In that sense, the pro-active policies in favor of trade regimes have neglected the increasing gap in the distribution of national income both within and between countries.

The growth-based trade regime can increase the income gap by way of two factors as follows (Dünhaupt, 2017: 289): (i) the global competition and (ii) the prices of traded goods. First, increased competition in the global era leads to a higher degree of price competition, which canalizes firms to use labor-saving technologies in order to compensate for the price decline and for the unit labor costs through the demand for reducing unskilled labor which creates a negative impact on labor's share. Second, the changes in prices of raw materials and semi-finished goods can negatively affect the income distribution over the medium- and long-run. If these prices are higher

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relative to wage costs in the case of free movements of capital covering both money and production, labor's share decreases.[6]

It is clear that the negative relationship between openness to trade and income distribution is explicitly related to the growth channel. However, it conceals the change in skill composition of workers, which is crucial to understanding the change in "skill premium" over time.[7] Despite the limitations of the traditional theorem for further analyses on widening income inequality in developing economies, Yay, Taştan, and Oktayer (2016): 583) use several approaches to clarify the global rise in the skill premium such as purchasing-power differences, outsourcing, differences in goods prices and factor returns, high technology transfers, and foreign direct investment. First, trade specialization may not be beneficial for poor countries because their purchasing-power differs significantly from rich countries in that the demand for the goods is different between those countries. In other words, the trade of goods is based on the income-level of the countries, whereby poor countries demand unskilled-labor-intensive goods from poor countries, whereas rich countries demand skilled-labor-intensive goods from rich countries.

The second critical dimension of trade openness can be deduced from the participation in "global value chains" (Dao, Das, Koczan, & Lian, 2017: 18–21). In advanced countries, the major reflection of this case is an increasing level of "capital-outsourcing" and "offshoring of production" including both intermediate inputs and services (Amiti & Wei, 2009). However, among developing countries, this is externalized through the re-exportation process in global value chains (Hummels, Jorgensen, Munch, & Xiang, 2014; Koopman, Wang, & Wei, 2014). According to Feenstra and Hanson (1996), the shift of production of intermediate goods and the division of the production process into different stages have resulted in skill /upgrading in developing countries. Therefore, the demand for skilled labor has significantly increased in global integration, which, in turn, has downwardly pressured the wages of unskilled labor by transferring them into the unqualified and low-return projects. In other words, the production of the final good is divided into different stages whereby most are produced in the most cost-efficient places (Grossman & Rossi-Hansberg, 2008). Related to this, Aghion and Williamson (1998) argue that the prices of complementary and substitute factors used in the production process would be altered to the detriment of the latter by way of reducing the demand for substitute goods.

Third, unlike the traditional perspective, there may be some conflicting results in relation to goods prices and factor returns due mostly to the problems in the reallocation of labor across different sectors, which, in turn, causes increasing wage dispersion among workers. The neoclassical arguments are based on the fact that freedom in trade would readjust the sectoral allocation of workers in which they would shift from contracting sectors to expanding sectors. However, this may not be the case when the labor markets are subjected to rigidities among domestic markets (Goldberg & Pavcnik, 2007). Additionally, besides the current problems in labor markets, there may be other factors which restrict to reach an optimal point in the economy and lead to prices and wages to change over time (Yay et al., 2016: 583).

Fourth, the competitive pressures in international markets may motivate firms/corporations to invest in high-tech goods, which then demand a higher level of productivity, educational progress, and quality enhancement, thereby rising skill premium and wage inequality. This is the result of promoting new information technology which leads to a higher level of productivity and

[6] Depending on these factors, several studies find a negative relationship between trade openness and income distribution. See Harrison (2005), Guscina (2006), EC (2007).

[7] Unlike the mainstream arguments, globalization tended to increase the wage gap between skilled and unskilled workers in favor of the former and thus has widened the income inequality through increased demand for skilled workers.

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endowment and thus boosts the overall skill premium, resulting in increased labor income inequality (Dabla-Norris et al., 2015: 18–19). Since there is a change in the technological production method, the demand for unskilled and skilled labor would be disproportionately allocated between sectors by dispelling the current job opportunities in line with the robotization process of the production system[8].

Finally, an increase in the degree of financial capital flows in the form of foreign direct investment (FDI) may push the wage level downward because it may necessitate a higher level of skilled labor in line with an ongoing high-technology transfer to the host country.

The traditional perspectives in favor of a higher degree of openness in the trade regime also suffer from several drawbacks, which are rooted in the assumptions of the neoclassical theorem. The post-1980s revealed that the decline in labor's share is obvious in developing economies and developed economies as well. Therefore, this leads us to examine the general framework of trade openness within the context of traditional assumptions. First, the factor immobility assumption is contradicted with an increasing amount of capital flows and labor migration across borders (EC, 2007). Second, the workers are heterogeneous in the sense of their skills and qualities and thus the skilled labor and unskilled labor may gain in developing and developed countries, respectively (Stockhammer, 2009). Third, even if the neoclassical trade theory is not supposed to be a well-functioning in terms of its predictions, it needs the government intervention in case of a growing gap between the rich and the poor to compensate the disintegrating majority[9] (Roine, Vlachos, & Waldenström, 2009: 977). Finally, Kratou and Goaied (2016: 136) argue that the potential effects of openness to trade on income inequality can be widened by increasing volatility and economic shocks. In particular, low-income households may be affected more than the others because of worsening income inequality if the country has a significant dependence on private capital flows and thus be open to the detrimental effects of short-run fluctuations.

All these above-mentioned factors imply that there is no definite theoretical conclusion about the trade openness–income inequality nexus; thus the topic is highly debatable. While the theoretical model is at the core of the debate in terms of its predictions and assumptions, the empirical studies also provide inconclusive results. For instance, Kraay (2006) and Goldberg and Pavcnik (2007) find that there is a positive link between trade openness and wage inequality. Wood (1995) states that the wages of unskilled labor are negatively affected by a more open trade regime and Sebastian (1997) finds that this negative effect is visible only for developed economies. Harrison (2005) and Guscina (2006) remark that the increase in trade share is negatively associated with labor's share. In the case of FDI, Faustino and Vali (2013) and Decreuse and Maarek (2015) find a direct link with decreases in labor's share, whereas Alderson and Nielsen (2002) argue that outward FDI can increase wage dispersion through the de-industrialization of an aggregate economy in advanced economies. Additionally, inward FDI may increase wage inequality by way of undermining solidarity among workers (Brady & Wallace, 2000). Bigsten and Munshi (2014) also find that the freedom to trade internationally can increase inter-occupational wage inequality in poorer OECD countries. Meschi and Vivarelli (2009) compare the effects of trade volume between developed and developing economies on income inequality and find that a higher volume of trade with high-income economies exacerbates the wage inequality in developing economies. This is also parallel to the empirical results found by Bergh and Nilsson (2010), IMF, 2007a, 2007b),

[8] For more information about the theoretical background and empirical outcomes for the relationship between technological advances and the labor allocation and thereby the widening wage inequality, see also Acemoglu (1998); Card and DiNardo (2002); OECD (2011); Ford (2016).

[9] See also Bardhan, Bowles, and Wallerstein (2006); Scheve and Slaughter (2007).

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Milanovic (2005), and Pica and Rodríguez (2011). Moreover, Rodrik (1997) empirically ponders the role of factor allocation on inequality and, in that sense, finds that the mobile factor benefits more from the trade openness. On the other hand, Gourdon, Maystre, and de Melo (2008) examine the impact of tariff reduction on income inequality which is subjected to relative factor endowment and conclude that lower rate of tariffs may increase income inequality in poor countries where the educational background for workers is at low levels. Finally, IMF (2017) supports some empirical evidence that there is a negative effect of the relative price of capital on the wage share along with technological upgrading in non-tradeable sectors, but not in tradeable sectors.

Neoclassical trade theory is criticized within the context of widening wage dispersion for various reasons across both developed and developing countries; it neglects the market structure in economies where the imperfect competition in goods and labor markets is highly effective on functional income distribution, which alters the shift parameters such as bargaining power and mark-up. In particular, this study assumes that income distribution is conditional on the mark-up pricing of firms, which, in turn, depends on change in the bargaining framework. In other words, wage inequality is indirectly determined by the conditions in the bargaining power of capital over the labor pursuant to the assumption of imperfectly competitive markets. According to Guschanski and Onaran (2017: 8), bargaining power measures are divided into two categories: direct and indirect factors.[10] While the former increase labor's power in negotiations, the latter emphasize their role in the fallback option when a problem occurs in negotiations. Indeed, each factor has a critical effect on the wage share based on the countries' conditions. Several studies have found that the wage share is subject to changes in direct factors such as union density, minimum wages, unemployment rate, and strike intensity (Dünhaupt, 2017; ILO, 2011; Jayadev, 2007; Kristal, 2010). For instance, the unemployment rate and union density are employed in numerous regressions to show their significant and differential effects on the wage share covering both advanced and emerging country-level analyses. However, there are also other indicators (e.g., collective bargaining power) which extend and correct the estimation outcomes in the empirical studies (Visser, 2006).

Besides the negotiation-based labor power, government measures may also lead to an increase in wage inequality. Welfare state retrenchment (Starke, 2006) has become a popular endeavor, especially in the post-crisis period of the global environment, with its impact on socio-economic problems, which, in turn, has stimulated a change in the distribution of national income. In particular, a fall in wage share can be deduced as a part of this stylized fact over time and across countries, covering both developed and developing economies. The only way in which labor can compensate their basic needs where the unemployment rate is at high levels is if the fallback option can support the condition of workers against capital (Jayadev, 2007; Onaran, 2009; Stockhammer, 2016). However, even if the more regulated labor markets have potential effects on wage dispersion, this is empirically not robust for the case of advanced economies (EC, 2007; IMF, 2007a, 2007b).

Moreover, Guschanski and Onaran (2017) find a negative correlation between functional income distribution and a higher degree of inequality in personal income distribution.[11] For example, unlike wage inequality, Daudey and Garcia-Penalosa (2007) find that the effect of a change in the wage share in distributing personal income is significant. On the other hand, it creates

[10] By following these categorization, these are assumed as the major point of this study in which the social solidarity is altered due to the change in bargaining framework.

[11] According to Guschanski and Onaran (2017: 8), widening personal income inequality strengthens the power of only one class over available resources.

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negative pressure on power relations between capital and labor against the latter in terms of both economic and political contexts, and, in turn, leads to increasing control over the means of production by privileged classes and elites. This is similar to what led Stiglitz (2008) to stress the importance of regulatory reforms and agenda to consolidate the confidence in financial markets. The closing of "revolving doors" and thereby the destruction of lobbyist behaviors of regulators narrow their opportunities for more control over the redistribution of resources and the rules in the socio-economic and political structure.

Another way that the bargaining framework can change over time relates to the technological progress, which has been exacerbated through Information and Communication Technology (ICT) developments in the post-1980 period, that disintegrated the labor markets into different categories by increasing the tasks of skilled labor and by canalizing unskilled labor to routine duties emerging mostly in the service sector. However, most importantly, it led to two distinguishing results that occurred in income distribution. On the one hand, developments in ICT displaced workers instead of machines for existing tasks in production (IMF, 2017). On the other hand, it contributed to an increase in the capital–output ratio (Karabarbounis & Neiman, 2014). Both of these two factors strengthened the negative pressure on wages and thus widened the wage dispersion across regions, countries, and sectors as well. The basic link for this negative correlation can be derived from the change in the elasticity of substitution between capital and labor (Guschanski & Onaran, 2017). If, for example, the elasticity is greater than one, it is expected that a higher capital usage leads to a lower wage share where the capital is a complement for skilled labor, but also a substitute for unskilled labor. Therefore, in countries where the unskilled–skilled labor ratio is high, any change in this ratio would create downward pressure on the wage share. In other words, the impact of a higher capital–output ratio on the wage share would be negative. According to Lavoie (2014), a firm target for the case of a higher profit rate would be in compliance with a higher mark-up and a higher capital–output ratio.

However, the inclusion of technological progress into this framework can allow for a much more flexible framework against the substitution effects between capital and labor in order to understand the fluctuations that occurred in wage–profit share. This case can be easily deduced from the gap between a wage rise and productivity level. If a rise in the overall wage level is placed under the increase in productivity level, the gap between profits and wages widens in favor of the former (Bhaduri, 2006). Additionally, if this widening income gap is stimulated by technological change and that stimulates the replacement of workers by automated techniques, the possible result is an increase in firing threat and thereby the reduction of the bargaining power of labor (Guschanski & Onaran, 2017). Hence, the social effect of technological progress (Marglin, 1974) should be considered as an important determinant of the income distribution. It means that the social effect of technology on the production process and thereby the distribution structure of national income is an indivisible part of the bargaining framework, which has a different theoretical concept than that of the elasticity of substitution[12] and productivity.

3 THE DATA

This paper investigates the effects of economic openness on income distribution employing a dataset derived mostly from the World Bank (World Development Indicators Database) and Penn

[12] For more information on empirical considerations about the effect of elasticity of substitution on wage share, see Harrison (2005), Karabarbounis and Neiman (2014), Doan and Wan (2017), IMF (2017).

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World Tables (PWT) version 9.0. However, the other variables are taken from different data sources.[13] The study has information for a total of 32 advanced economies in our sample, classified on the basis of income-level as proposed by the International Monetary Fund and these are listed in Table A2, for the period 1996 to 2014.[14]

The dependent variable is constructed on two alternative indicators. In other words, the paper estimates panel regression models by including both household-based and class-based measures, which are the Gini coefficient derived from Solt's (2016) database and the labor share of income comes from the PWT version 9.0, respectively. Therefore, the empirical results provide a piece of information on the changes in income distribution from both personal and functional perspectives.

On the one hand, our first measure of income distribution is the Gini coefficient (GINI), which is calculated by Solt (2016) and is produced in the Standardized World Income Inequality Database (SWIID). The unique feature of this database is that the other sources do not have balanced panel data for the Gini coefficient, but SWIID fills this gap by providing an entire time-series for many countries. Additionally, SWIID classifies data for the Gini coefficient into two categories as Gini Gross (pre-tax, pre-transfer) and Gini Net (post-tax, post-transfer). Furthermore, the general context of Solt's (2016) database has a further importance for examining the effects of changes in Gini coefficient on an aggregate economy in the context of low-income economies. In this study, the Gini coefficient is employed as a Gini Net measure to account for the effects of economic openness on a household-based measure of income distribution.

Our second measure of income distribution is the labor share of national income. Traditionally, labor's share is calculated as compensation of employees divided by GDP or value-added and thus the remainder is assumed as capital's share. Recent studies (e.g., Gollin, 2002; Krueger, 1999) assert that labor's share may produce a biased result if the compensation of employees does not include the earnings of self-employed people. Therefore, to account for this problem, labor's share is adjusted from this indicator taken from PWT version 9.0 and is named as "adjusted labor share of income" (i.e., LABSH) (Dünhaupt, 2013). In this database, labor's share contains both compensation of employees and self-employed and GDP by ignoring taxes and subsidies[15]. All in all, there is a negative correlation between labor's share and capital's share which means that if labor's share increases accruing in national income, the rest of the share of income goes to the capital.

Our main determinants of interest—namely the Gini coefficient and labor's share—are relied upon both personal and functional perspectives on income distribution. However, these measures are separated from each other on a theoretical basis. Therefore, the comparison between the two indicators cannot be easily accomplished across countries. But a virtue of the Gini coefficient over labor's share can provide a way to compare income inequalities among countries which have different income levels. However, it has no power to analyze the sectoral income distribution. In that sense, the functional income distribution analyses sum up the class inequality on the basis of the capital–labor antagonism for the share of national income. However, one disadvantage is the nature of the informal sector, for which data is lacking all over the world. As Jayadev (2007: 426) states that "while, in theory, the informal sector is to be included in the data, in practice, by their very nature, enterprises from this sector may not be". Hence, the legal authorities may report labor's share much higher or lower than the true level in an aggregate economy which

[13] For more information about the list of variables, please see Table A1.

[14] The selection of this period basically depends on the aim for getting a balanced panel structure and thereby the availability of the data sources.

[15] For more information about different types of calculation methods for the labor share of income, see Guerriero (2012).

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is very common in advanced and developing economies even if the weights of importance are different in each country.

Table 1 presents selected descriptive statistics for the 18 high-income economies in our sample. As the statistical outcomes show, the trends in the Gini coefficient and labor's share are volatile and thus may not provide a significant correlation with each other. For instance, across the sample, LABSH has on average 60% in the total national income. In countries like Israel, Italy, Norway, and Sweden, the distribution of income on the basis of aggregate national share is lower than average for the first and the last year of the sample period, which in fact depends on different factors from political, social and sociological spheres. On the other hand, in Belgium, Canada, Denmark, and Switzerland, class inequality is much narrower than for the rest of the sample.

However, the same situation is not valid for the Gini coefficient. In other words, while the countries may have labor's share above the average level, they may have a higher Gini coefficient in which it indicates that the household-based income differences are at high levels. Therefore, the descriptive statistics lead us to argue that even countries where the labor's share is above average may have a lower Gini coefficient. This result is important as there is no systematic relationship between these two measures. Hence, it means that the intended of personal income distribution and functional income distribution are different from each other.

Additionally, in Table 1, Belgium has the highest openness level in trade (i.e., TRADEOPEN) and many of the countries are completely open to the financial transactions (i.e., FINOPEN). A variable with considerable effect on income distribution—the economic globalization, including both trade openness, financial openness and FDI flows—is on average equal to 66.7 in 1996 and 76.0 in 2014. There are, however, significant differences across the selected countries with Belgium having the highest degree of economic globalization (83.8 in 1996 and 88.7 in 2014) and Japan the lowest (44.3 in 1996 and 61.1 in 2014). In Table 1, as expected, the government share (i.e., GOVSHARE) follows the negative trend with globalization measures since the globalization movement isolates the role of government in the functioning of the economic system: with an average government share of total GDP of 14.5% in 1996 and 18.5% in 2014, Israel shows the highest level (22.1% in 1996 and 23.3% in 2014) and Switzerland the lowest (7.3% in 1996 and 7.59% in 2014), on average.

The study also includes industry value added (annual % growth) as an additional variable to investigate the effects of change in capitalist dynamics over the post-1980s period of globalization on income distribution. Whereas structural change in the capitalist system is measured by different indicators, the industrial sector share (i.e., INDUSTRY) is one of the most crucial variables to account for the negative relationship between financialization and industrialization over the neoliberal period of the 1980s, which is obtained from the World Bank's World Development Indicators Database. Across the selected sample in Table 1, the average growth ratio of industrial sector is approximately 2.27% in 1996 and 1.77% in 2014 with Korea (7.83% in 1996 and 3.15% in 2014) and Israel (8.69% in 1996 and 1.49% in 2014) at the largest record and Italy (0.26% in 1996 and -1.39% in 2014) and Switzerland (-1.83% in 1996 and 0.37% in 2014) at the lowest, on average.

Finally, one of the critical measures for bargaining power of labor, which is the unemployment rate (% of total labor force) (i.e., UNEMP), is included in the analysis to examine the antagonism between capital and labor over time and thereby the class inequality. In Table 1, the selected economies with the highest unemployment rate are Spain (22.1% in 1996 and 24.4% in 2014), France (12.4% in 1996 and 10.3% in 2014), and Israel (8.46% in 1996 and 5.89% in 2014); those countries with the lowest percentage of unemployment level are Korea (2.05% in 1996 and 3.53% in 2014), Japan (3.35% in 1996 and 3.58% in 2014), and Switzerland (3.71% in 1996 and 4.83% in 2014).

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TABLE 1 Descriptive statistics

Country	LABSH		GINI		TRADEOPEN		FINOPEN		ECONGLOB		GDPGRW		INDUSTRY		GOVSHARE		UNEMP		HUMCAP	
	1996	2014	1996	2014	1996	2014	1996	2014	1996	2014	1996	2014	1996	2014	1996	2014	1996	2014	1996	2014
Australia	59.7	56.8	31.3	33.2	29.0	60.8	93.9	87.9	61.2	66.5	3.87	2.56	3.18	3.32	14.3	16.0	8.51	6.07	3.49	3.48
Austria	62.5	58.8	27.5	27.8	70.9	117.6	100	100	75.9	83.3	2.35	0.83	1.66	0.56	14.2	17.9	5.28	5.62	3.05	3.33
Belgium	62.6	62.9	25.2	25.8	121.9	200.5	93.9	100	83.8	88.7	1.59	1.29	2.25	3.31	16.7	22.7	9.48	8.52	2.91	3.12
Canada	62.7	61.9	29.5	31.2	47.1	75.1	100	100	64.7	70.1	1.61	2.86	N/A	4.32	14.7	17.2	9.62	6.91	3.42	3.68
Denmark	65.1	64.0	22.2	25.2	83.1	143.3	100	100	77.9	83.7	2.90	1.62	0.16	0.48	17.8	24.5	6.84	6.59	3.22	3.51
France	62.3	63.0	28.3	29.9	46.4	67.8	100	100	67.6	77.3	1.39	0.95	−0.35	−0.28	18.1	22.8	12.4	10.3	2.86	3.13
Germany	66.3	62.2	26.6	28.9	47.3	87.7	100	100	68.1	78.7	0.82	1.93	−2.90	4.99	13.4	17.7	8.82	4.98	3.51	3.66
Israel	60.1	53.6	33.1	36.9	48.3	77.6	16.6	100	54.7	70.7	5.31	3.41	8.69	1.49	22.1	23.3	8.46	5.89	3.24	3.68
Italy	54.1	53.9	34.1	33.2	34.7	57.2	100	100	63.7	68.1	1.28	0.11	0.26	−1.39	12.8	17.8	11.9	12.7	2.68	3.07
Japan	65.9	60.2	28.7	29.9	21.3	38.6	93.9	100	44.3	61.1	3.09	0.37	4.16	3.00	13.2	20.9	3.35	3.58	3.28	3.54
Korea	57.2	51.9	28.5	30.7	35.5	77.6	16.6	71.7	36.3	66.5	7.59	3.34	7.83	3.15	9.3	15.1	2.05	3.53	3.05	2.59
Netherlands	62.1	59.5	25.3	26.5	105.6	169.7	100	100	82.7	89.0	3.57	1.42	2.19	-0.17	16.5	22.4	6.42	7.42	2.08	3.32
Norway	56.6	53.3	24.1	24.8	70.8	87.8	81.9	100	72.3	73.4	5.03	1.97	5.80	1.43	14.3	14.1	5.04	3.48	3.29	3.61
Spain	64.4	57.9	33.7	34.3	35.7	57.9	87.9	100	65.1	75.0	2.67	1.38	1.50	0.85	14.2	18.9	22.1	24.4	2.57	2.88
Sweden	53.7	56.5	22.9	25.7	76.9	119.8	93.9	100	74.1	83.8	1.52	2.60	2.26	0.58	19.7	23.1	9.5	7.92	3.16	3.38
Switzerland	66.4	65.4	29.6	29.3	97.8	162.1	100	100	74.6	85.4	0.54	2.45	-1.83	0.37	7.03	7.59	3.71	4.83	3.50	3.66
United Kingdom	58.7	61.2	34.1	32.9	41.8	70.9	100	100	74.4	79.8	2.54	3.05	1.49	3.65	13.7	19.4	8.19	6.11	3.38	3.73
United States	61.2	60.3	35.8	37.8	16.8	31.7	100	100	59.5	66.8	3.79	2.57	N/A	2.21	9.10	11.3	5.4	6.17	3.53	3.72
Average	*61.2*	*59.1*	*28.9*	*30.2*	*57.3*	*94.7*	*87.7*	*97.8*	*66.7*	*76.0*	*2.86*	*1.93*	*2.27*	*1.77*	*14.5*	*18.5*	*8.17*	*7.5*	*3.18*	*3.45*

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The average rate of unemployment of the sample in Table 1 is approximately 8.71% in 1996 and 7.5% in 2014.

Apart from the multi-functional role of indicators among themselves, Table 2 shows the correlation of these variables in order to assess whether the link between them is strong enough or not. Both globalization variables, including TRADEOPEN and FINOPEN, are not perfectly correlated with the data for LABSH and GINI. In addition, the covariates of these two measures are not so high and differ in sign, especially for the data about the distributional issues. The same case is also valid for the bargaining power measure (e.g., UNEMP) which is both negatively but not perfectly correlated with the globalization measures: TRADEOPEN (-0.222) and FINOPEN (-0.040). However, INDUSTRY, which is an essential parameter for the real investment and thereby the economic prosperity in advanced economies, correlates strongly and positively with GDP per capita growth (0.872) (i.e., GDPGRW). Finally, Table A1 presents the details for the list of variables used in our analysis.

4 | HYPOTHESIZED PROPOSITIONS AND ECONOMETRIC MODEL

With the above-mentioned variables in mind, the major channel through which the globalization measures, including both trade openness and financial openness, can affect the income distribution and thereby the social solidarity is the relative bargaining position between labor and capital. The paper theoretically implies that the rise in the degree of openness in trade regime and financial account possibly affects the relative bargaining power of labor by way of increasing the threat option of capital; and therefore, the share of total national income is biased towards capital. In this context, the change in labor's share is affected through different channels by which globalization interacts with various channels of influence in an aggregate economy. This prompts us to propose the following hypotheses which are empirically tested in the next part.

H1. Trade openness decreases the bargaining power of labor relative to that of capital, decreasing labor's share of national income, increasing income inequality, and unsettling social cohesion.

H2. Financial openness decreases the bargaining power of labor relative to that of capital, decreasing labor's share of national income, increasing income inequality, and unsettling social cohesion.

H3. Globalization decreases the bargaining power of labor relative to that of capital, decreasing labor's share of national income, increasing income inequality, and unsettling social cohesion.

All three of these hypotheses mutually interact with each other and thus any change in one indicator may in principle affect the whole hypothesized structure based on alleviating or intensifying the capital–labor antagonism over time and across countries. However, it is worth noting that the effect of this change in variables has not the same weight on labor's share and income inequality. Therefore, the robustness of these measures may not be significant in all models even though the expected results are provided in terms of their correlation with the other variables.

To briefly summarize our empirical results on the basis of these hypotheses, both openness measures covering trade and finance tend to be negatively related to labor's share and positively related to the Gini coefficient (consistent with H1 and H2) in most of the specifications. Similar to these empirical outcomes, the globalization variable which includes sub-factors in terms of

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TABLE 2 Correlation matrix

Variables	LABSH	GINI	TRADEOPEN	FINOPEN	ECONGLOB	GDPGRW	INDUSTRY	GOVSHARE	UNEMP	HUMCAP
LABSH	1.000									
GINI	−0.339	1.000								
TRADEOPEN	−0.159	−0.061	1.000							
FINOPEN	−0.059	0.185	0.141	1.000						
ECONGLOB	−0.121	−0.075	0.595	0.502	1.000					
GDPGRW	−0.216	0.095	0.050	−0.104	0.072	1.000				
INDSHARE	−0.195	0.031	−0.002	−0.128	0.034	0.872	1.000			
GOVSHARE	−0.041	−0.031	−0.158	−0.158	−0.037	−0.022	0.025	1.000		
UNEMP	−0.132	0.199	−0.222	−0.040	−0.114	−0.124	−0.043	0.486	1.000	
HUMCAP	0.032	−0.227	−0.052	−0.044	−0.079	−0.096	−0.035	−0.062	−0.240	1.000

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trade and financial openness measures tend to be negatively related to labor's share and positively related to the Gini coefficient (consistent with H3). Therefore, the empirical analysis expects that both hypotheses are in compliance with each other.

Our aim is to test these hypotheses on the basis of the relationship between economic openness and income distribution over the 1996–2014 period for the sample which includes 32 advanced economies. To estimate this relationship, the study considers the dynamic panel data model in Equation (1) as follows:

$$DIST_{it} = \alpha DIST_{it-k} + \gamma OPEN_{it-k} + \beta X'_{it-k} + \theta_t + \mu_i + \varepsilon_{it}, \quad (1)$$

where $DIST_{it}$ is the income distribution variable of country i ($i = 1, \ldots, N$) in period t ($t = 1, \ldots, T$), $DIST_{it-k}$ refers to the lagged value of income distribution, $OPEN_{it-k}$ is the lagged openness measures for trade and finance, X'_{it-k} is a vector of lagged control variables, θ_t captures the unobserved time effects and μ_i captures the unobserved country-based fixed heterogeneity. ε_{it} is the white-noise error term. Since there might be a heterogeneity problem in the estimated models, standard errors are clustered at the country level. Additionally, the paper checks cross-sectional dependency through the dynamic fixed-effect models because there might be some correlation on the basis of other factors, irrespective of the economic side.

The major reason why the empirical methodology is primarily constituted on the estimation of a dynamic structure for the regression model can be deduced from the formation of dataset. In particular, the data that the empirical analysis is used might be characterized by complex error structures which indicate that the disturbances are likely to be heteroskedastic and contemporaneously correlated across panels. Therefore, especially for the fixed-effects linear regression model, this method provides a way to solve diagnostic problems. Moreover, the use of the GMM method allows us to get rid of the specific type of endogeneity issue of explanatory variables, simultaneity and unobserved heterogeneity of the sample countries. Since the time span is large enough, the use of dynamic structure in the fixed-effects model indirectly asserts that the standard nonparametric time-series covariance matrix estimators can be used for all general forms of spatial and periodic correlations as robust. Therefore, in some cases, instead of dynamic models, the static models are almost always wrongly specified since the within-group error terms are serially autocorrelated. This is somehow problematic for the estimation results due to the fact that statistical inference and point estimates become statistically invalid. In contrast, the use of dynamic models also allows us to make a distinction between short- and long-run effects of explanatory variables on income distribution, which is crucial for the policy-making process in sample countries.

In Equation (2), the study also uses the system-GMM estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998) in order to solve the correlation issue emerging between μ_i and $DIST_{it-k}$. Unlike Equation (1), Equation (2) presents the first-difference transformation in terms of reducing the potential country-fixed effects μ_i:

$$\Delta DIST_{it} = \alpha \Delta DIST_{it-k} + \gamma \Delta OPEN_{it-k} + \beta \Delta X'_{it-k} + \Delta\theta_t + \Delta\varepsilon_i t, \quad (2)$$

where Δ is the first-difference estimator denoted as $\Delta DIST_{it-k} = DIST_{it} - DIST_{it-l}$. Also, orthogonality conditions are determined in the case that $E(DIST_{it-k}\Delta\varepsilon_{it}) = 0$. $DIST_{it-k}$ is the optimal lag length of income distribution indicators. The second and further lags of $DIST_{it}$ are also included in the model as an instrument for the residuals (Heid, Langer, & Larch, 2012). According to Che, Lu, Tao, and Wang (2013: 161), the OLS estimates in Equation (1) produce biased results for α and γ because $DIST_{it-k}$ is a function of ε_{it} and thus $Cov(\Delta DIST_{it-k}, \Delta\varepsilon_{it})$ is different from zero.

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The consistent estimation results need no second-order serial correlation in $\Delta\varepsilon_{it}$. Therefore, the paper benefits from the use of the AR(2) test suggested by Arellano and Bond (1991) to control for the second-order serial correlation of $\Delta\varepsilon_{it}$. Additionally, it utilizes the J-test of Hansen (1982) to confirm whether the orthogonality conditions are still relevant.

However, the persistence of the dependent variable in time and the small sample period lead to the emergence of the weak instrument problem which then affect the consistency of the standard errors. The major way to solve this problem is to take averages of the panels in terms of the specified years. Though this is accepted as relevant in a theoretical context, an alternative method may be needed in some cases if the problem persists in the analysis since the number of observations is reduced together with causing unreliable point estimates and hypothesis tests. Therefore, Arellano and Bover (1995) and Blundell and Bond (1998) suggest the transformation of the Equation (1) type of level estimations by augmenting the difference-GMM. This leads to the use of lagged first differences of the dependent variable as an instrument for lagged dependent variables. In that vein, the new estimator changes the moment conditions to E $(\Delta d_{it-k}(\delta_i + u_{it})) = 0$ where $t = 3,\ldots,T$.

To foreshadow the significance of this estimator, Hansen (1982) provides a difference J-test. The extension case of this new econometric procedure is referred to as the system-GMM. In particular, the system-GMM estimator increases the asymptotic efficiency gains along with the moment conditions produced by Arellano and Bover (1995) and Blundell and Bond (1998). However, it also negatively affects the orthogonality conditions for both J-test and difference J-test as the number of instruments increases in time, where $\mathrm{E}(d_{it-k}\ \Delta u_{it}) = 0$ and $\mathrm{E}(\Delta d_{it-k}(\delta_i + u_{it})) = 0$, respectively. This is called the "size distortion" in which there is a positive link between the time dimension and the number of instruments.[16] According to Roodman (2009a), the proliferation of the number of instruments leads to a finite sample bias by virtue of the endogeneity problem and to an inaccurate specification test results. Additionally, Roodman (2009a) classifies the reasons for instrument proliferation as follows: (i) overfitting endogenous variables, (ii) tentative estimates of the GMM optimal weighting matrix, and (iii) bias in two-step standard errors. The best way to solve this problem depends on the fact that the collapse of the instruments may reduce the possibility for finite sample bias and hence may possibly validate the orthogonality conditions.

5 | EMPIRICAL RESULTS

The empirical results are sequentially presented in three steps. The empirical analysis starts by showing our baseline findings based on Equation (1) for dynamic pooled OLS and dynamic fixed-effects methods and Equation (2) for bootstrap-corrected fixed effects and two-step system-GMM procedures by employing annual data. In this framework, the estimated models ignore the control variables in order to reveal the causal relationship between the variables related to the income distribution (i.e., labor share of income) and the globalization indicators (i.e., trade openness and financial openness).

In the following step, the study performs some robustness/sensitivity analysis regarding the methodology employed. First, it implements the two-step system-GMM with different and varied collapsed instruments. The empirical part also provides some evidence using an alternative methodology. In particular, it shows results obtained by fractional response fixed-effects

[16] For more information, see Andersen and Sørensen (1996), Bowsher (2002), and Roodman (2009a).

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models (Ramalho, Ramalho, & Coelho, 2018) in order to to make robustness check in comparison to dynamic model results.

In the last step, the study performs some extensions by including some more control variables to explore some complementary channels to baseline analysis. All in all, the standard errors are fully adjusted from the diagnostic issues and they are robust for the full sample. In addition, time- and country-fixed effects are statistically significant in all regressions, including both baseline and sensitivity tests.

5.1 | Baseline estimation results

Table 3 summarizes the estimation results from the baseline analysis for the advanced economies by employing annual data over the 1996–2014 period. In that context, Table 3 benefits from the labor's share data and Gini coefficients in order to understand each effect of globalization measures on class- and household-based income distribution, respectively. Furthermore, in Table A3, the sample period becomes larger for the period between 1985 and 2014.[17] The major reason for using a long time span is the consideration that openness may possibly take time to affect income distribution. Therefore, this extended time period allows us to check whether the effect of openness on income distribution changes over time. However, it should be noted that the former socialist countries are excluded from the sample since there is no data for most of the variables which leads to failure in using a balanced sample. All in all, the results for each case do not change significantly, which means that the estimation outputs for different time periods complement each other.

The study divides the data into two periods 1996–2005 and 2006–2014 to investigate different economic dynamics over time in Tables A4 and A5. According to the extended empirical results, the empirical findings also hold for time-based analyses as shown in Table 3. Therefore, the long-run effects of openness in trade regime and financial accounts on the income distribution are validated in the statistical framework even if the periods become narrower or longer. These outcomes are crucial to show the theoretical difference from the mainstream approach which vigorously argues that implementing a higher openness in both trade and financial accounts would result in a decline in an unequal distribution of income, especially in developed economies. However, the results of this paper both validate our hypotheses and strengthen those hypotheses in the analytical structure.

In particular, the baseline estimations are grounded on four different estimation methods to test the significance of the above-mentioned hypotheses in consideration of possible econometric problems that may emerge in those techniques and each of which is statistically powerless to address these technical issues. First, the paper uses a fixed-effects panel data method to correct the potential heterogeneity problem that may occur in the estimated models and this allows us to cluster the standard errors at the country level. Regarding the use of this method, the estimation process can test the effects of cross-sectional dependency through the models since there may be potential dynamics among the variables towards adding control and proxy variables. While these advantages are given for the fixed-effects models for panel data, the limitations of these models are as follows Hill, Davis, Roos, and French (2019): (i) a culture of omission, (ii) low

[17] Note that some of the countries are excluded from the sample since there is no data for this countries in case of labor share of income, Gini coefficient, trade openness, and financial openness. These excluded countries can be ranged as follows: Estonia, Iceland, Latvia, Lithuania, Luxembourg, Slovakia, and Slovenia.

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TABLE 3 Baseline estimation results

	Annual Data, 1996–2014									
	Pooled	**FE**	**BCFE**	**Sys-2**	**FE**	**Pooled**	**FE**	**BCFE**	**Sys-2**	**FE**
	OLS	**OLS**	**Dynamic**	**GMM**	**OLS**	**OLS**	**OLS**	**Dynamic**	**GMM**	**OLS**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10)**
	Dependent variable is labor's share					**Dependent variable is Gini coefficient**				
$LABSH_{t\text{-}5}$	0.842***	−0.079	−1.041***	0.862***						
	(0.069)	(0.079)	(0.063)	(0.078)						
$GINI_{t\text{-}5}$						0.947***	0.136*	1.535***	0.961***	
						(0.021)	(0.078)	(0.047)	(0.025)	
$TRADEOPEN_{t\text{-}5}$	−0.007**	−0.052	−0.018**	−0.007*	−0.043***	−0.000	0.380	0.023	0.046	1.090***
	(0.003)	(0.033)	(0.007)	(0.004)	(0.007)	(0.098)	(0.480)	(0.096)	(0.078)	(0.213)
$FINOPEN_{t\text{-}5}$	0.001	0.006	0.002	0.004	0.004	−0.168	0.196	0.069	0.906	0.595**
	(0.019)	(0.020)	(0.008)	(0.019)	(0.009)	(0.403)	(0.697)	(0.213)	(1.292)	(0.269)
Hansen *J*-test				[0.25]					[0.25]	
AR(2) test				[0.21]					[0.20]	
Wald χ^2 (prob.)				0.000					0.000	
No. of inst.				30					14	

(Continues)

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TABLE 3 (Continued)

	Annual Data, 1996–2014									
	Pooled OLS (1)	FE OLS (2)	BCFE Dynamic (3)	Sys-2 GMM (4)	FE OLS (5)	Pooled OLS (6)	FE OLS (7)	BCFE Dynamic (8)	Sys-2 GMM (9)	FE OLS (10)
	Dependent variable is labor's share					Dependent variable is Gini coefficient				
No. of obs.	448	448	448	448	448	448	448	448	192	448
No. of countries	32	32	32	32	32	32	32	32	32	32
R-squared	0.776	0.909			0.899	0.955	0.984			0.982

Notes: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Pooled OLS regression in columns (1) and (6), with robust standard errors clustered by country are in parentheses. Fixed-effects (FE) OLS regressions in columns (2), (5), (7), and (10), with country dummies and robust standard errors clustered by country are in parentheses. Columns (3) and (8) use the bootstrap-corrected fixed effects (BCFE) method of De Vos et al. (2015) with robust standard errors, and columns (4) and (9) use the two-step system-GMM (Sys-2 GMM) of Arellano and Bover (1995) and of Blundell and Bond (1998) extended by Roodman (2006), with robust standard errors and by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. Year dummies are included in all pooled OLS, fixed-effects, and system-GMM regressions. In columns (4) and (9), the regression results are also provided Hansen *J*-test statistics to specify whether the overidentifying restrictions are valid or not. The collapsing method is used in system-GMM models to get rid of from too-many-instruments problem and to provide the orthogonality conditions. The joint null hypothesis implies that the instruments are valid, which means that they are uncorrelated with the error term and that the excluded instruments are correctly ignored from the estimated equation. The results of AR(2) are also provided to check for autocorrelation in case of proper maximum lag distance. For the case of the BCFE method representing in models (3) and (8), the selection of the error resampling pattern is based on cross-sectional heteroskedasticity error terms and the analytical heterogeneous initialization is used in the regression. In addition, the convergence option is selected in all models. Standard errors are estimated from the bootstrap distribution of the BCFE estimator. Confidence bounds for the *t*-distribution are calculated on the basis of bootstrapped standard errors.

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statistical power, (iii) limited external validity, (iv) restricted time periods, (v) measurement error, (vi) time invariance, (vii) undefined variables, (viii) unobserved heterogeneity, (ix) erroneous causal inferences, (x) imprecise interpretations of coefficients, (xi) imprudent comparisons with cross-sectional models, and (xii) questionable contributions vis-à-vis previous work.

In Table 3, columns (1)–(4) show that there is a high persistency for labor's share in which the coefficients are statistically significant at the 1% level. However, the results are not the same for trade openness and financial openness. Though the effect of the globalization movement in the trade channel on labor's share is significant and negative in many of the models, the coefficient for financial openness is not significant in each specification (the only exception is model 10). In that vein, the empirical outcomes indicate that openness in the financial account does not matter for the distribution of income in the class-based framework across advanced economies. This equivalently means that the trade channel is more important than the financial transactions and relations for understanding the changing antagonism between capital and labor over time and thereby the decline in the significance of social solidarity and collective conscience. Moreover, column (5) shows a simpler specification in which lagged labor share of income is dropped. The major reason for eliminating this lagged dependent variable from the regressions is to show that the hypothesized correlation is still relevant and statistically significant between the dependent variable and regressors. With the trade openness, there is still evidence of a significant effect of income distribution on a measure of openness, and in this case, the corresponding fixed-effects estimates are accepted inside the standard error bands.

Second, the remaining columns (6) to (10) show that the same arguments cannot be made for the case of using the Gini coefficient to estimate the effects of openness variables on income inequality. The coefficient estimates on trade openness and financial openness are to a large extent statistically insignificant even though they are positive in almost all columns. The only exception to this generalization of the estimation results is seen in column (10). On the one hand, the empirical result for fixed-effects model in column (10) indicates that financial openness is positively correlated with the level of income inequality, which means that a higher degree of openness in financial accounts increases the income inequality to the detriment of the bottom segments of households. On the other hand, the empirical result for fixed-effects model in column (10) also shows that openness in trade regimes increases the inequality in the distribution of an aggregate income on the household-based levels.[18]

In addition to the fixed-effects linear models, the study also uses bootstrap-corrected fixed effects (BCFE) estimations and inference in dynamic panel data models, measured in columns (3) and (8). The major aim of using the BCFE method is to estimate the given models with the fixed-effects estimator produced by De Vos, Everaert, and Ruyssen (2015) and thus to correct its

[18] In consideration of the results in Table A3, the study has to explain one crucial point which is related to the coefficient of lag LABSH (i.e., $LABSH_{t-5}$) in model 3 and the coefficient of lag GINI (i.e., $GINI_{t-5}$) in model 8 since they are higher than one. While this seems to be problematic in a statistical context, one possible reason can be deduced from the measuring procedures of the labor share of income and the Gini coefficient. In other words, whether the slightly high labor share of income and Gini coefficients in models are statistically given, they basically depend on how the dependent variables are measured. The study assumes that since the labor share of income and Gini coefficient are measured by using annual data rather than monthly data, the coefficients may in some cases exceed the given gap. Additionally, another reason may be the lack of instrumental variables, which is substantially related to the theoretical background of the BCFE method. Therefore, the inclusion of these variables or the implementation of higher lag values of given variables may reduce the number of the coefficient of the labor share of income and the Gini variable. Finally, lagged labor share of labor and lagged Gini coefficients are highly significant and indicate that there is a considerable degree of persistence in income distribution.

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small T bias.[19] The algorithm estimates the bias of fixed-effects to get rid of using the analytical correction formulas, which is therefore proper for higher-order models with a potential nonstandard error structure. Since there are critical diagnostic problems in the models, all regressions take into account heteroskedasticity and cross-sectional dependence patterns in line with the choice of the resampling scheme. The estimation results are substantially crossed with the linear fixed-effects estimation results. However, since the estimates can still suffer from the endogeneity problem, the paper also considers the system-GMM method in further estimations.

The paper also benefits from other methods to control these issues. For example, the system-GMM is used to get an estimator to solve endogeneity, heteroskedasticity, and serial correlation problems in static panel data models. In that sense, there are several advantages to consider the system-GMM estimator since it controls for time-invariant country-fixed effects; solves the endogeneity problem of lagged dependent variable; leads to a certain degree of endogeneity in the other regressors, and optimally integrates information on cross-country differences in levels along with within-country variation in changes (Fukase, 2010: 205). However, Roodman (2009b: 87) states that the disadvantage of system-GMM is that it is complicated and so can easily generate invalid estimates where the estimator's purpose, design, and limitations unwittingly misuse the estimators. The dynamic panel data regressions were corrected by using two-step system-GMM estimators by using Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. The major reason to use the two-step robust option is that standard errors are possibly downwardly biased even though the standard covariance matrix is already robust in theory in a two-step estimation. To correct this problem, the Arellano–Bover and the Blundell–Bond system-GMM estimations were regressed by the variance correction proposed by Windmeijer (2005) in columns (4) and (9) of Table 3. The same procedure was also followed in other system-GMM estimations. The major reason to estimate the system-GMM with the lagged dependent variables, in contrast to fixed-effects model without the lagged labour share or lagged Gini coefficient, was to account for the persistence of the lagged dependent variables, which may affect the statistical validity of the regressors and the chosen instruments as well.

In particular, the empirical results presented in Table 3 complement each other since each method has its own dynamics where the technical deficiency in one method is controlled by the other one. For instance, the system-GMM accounts for the country-fixed effects (Roodman, 2009b). The fixed-effects are part of the disturbance term of the untransformed level model. Therefore, the instrumental variables are chosen to adjust from this issue such that they do not correlate with these unobserved time-fixed effects. Regarding the system-GMM, the instruments are typically the one-time changes of the given variables. The statistical validity is conditioned to the case that all variables in the regression model should be jointly mean stationary. However, the jointly mean stationary condition of the variables is not directly tested in the model. Therefore, the paper deals with the Hansen overidentification test results whether it validates the conditions for the instruments. In that vein, system-GMM accounts for differences among the sample countries that are stable in time if the assumption of jointly mean stationary is provided in the regression models. In other words, the instruments should be statistically significant where there is no correlation with the disturbance term. Mean stationary, in that sense, can provide the validity of these outcomes. So, the system-GMM estimator accounts for the fixed-effects since all these results are provided in Table 3. Therefore, each method complements their empirical results to a large extent.

[19] For more information on the theoretical background, see Nickell (1981).

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All in all, the investigation of the causal relationship between the globalization indicators and income distribution needs further attention to be tested by additional specifications, which include different types of control variables and alternative measures for economic openness. However, note that the estimations representing in all these above-given models provide almost significant results for the case of economic openness-income distribution nexus across advanced economies in control of the inclusion of fixed-effects, which control for time-invariant country-specific characteristics. Hence, our empirical results, mostly on the basis of trade openness indicator, show that the traditional wisdom on this issue is not prevailing for the advanced countries at least.

5.2 | Analysis

5.2.1 | More results using two-step system-GMM

Table 4 presents the two-step system-GMM estimation results for both labor share of income and Gini coefficient. The main rationale for using this econometric procedure is to solve the weak instrument problem and the endogeneity issue, which cannot be done by fixed-effects even though the lagged values of all variables are used in the models. Therefore, the fixed-effects estimator is supposed to provide biased and inconsistent results for small T and large N panels. Additionally, the inclusion of lagged values of dependent variables into the regression may lead to spurious results since it may create an endogeneity problem in itself. Therefore, in this section, we employ a two-step system-GMM estimator to dispel those kinds of problems.

Following Arellano and Bover (1995), Blundell and Bond (1998), and Roodman (2009b), the study compares the initial results provided for Equation (1) with the estimates of system-GMM through using collapsing instruments along with the orthogonality conditions and two-step method presented in Table 4. The number of observations may decrease due to lagged structures thus it follows Arellano and Bover (1995) method by using forward orthogonal deviations instead of first-differencing of the variables. The AR(2) and Hansen (1982) J-test results are also reported to check the validity of instruments and orthogonality conditions, respectively. Furthermore, lag years of GMM-style instruments are used in system-GMM estimations to reduce the impacts of instrument-type problems on coefficients. Finally, it furthers practice GMM-style instruments by way of including additional explanatory variables treated as endogenous.

The empirical results in Table 4 show that the lagged value of income distribution, including both labor share of income and Gini coefficient, is statistically significant in most of the regressions. Theis means that the persistent characteristics of within-country income distribution for two dependent variables are substantially significant. As can be seen in Table 4, the estimation results are similar to what the empirical analysis found in baseline analysis. While the degree of openness to trade is negatively correlated with labor share and statistically significant in almost all regressions, the reverse is valid for financial openness. Therefore, even in the presence of solving for weak instrument problem and endogeneity issue, the trade channel is more effective in contrast to financial accounts on the distribution of income. However, the same is not true in the case of income inequality measured by the Gini coefficient in which none of the openness indicators are statistically significant in the models although the sign of coefficients is almost positive. In addition to these estimation results, the same models are used in Table A6 to check the statistical validity of openness–income distribution nexus for a longer time period. Although the endogeneity problem is considered in the models, the estimation results do not change

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TABLE 4 System-GMM estimation results

	Annual Data, 1996–2014							
	Sys-2 GMM-IV (TL, FL) (1)	**Sys-2 GMM-IV (TL, FL) (2)**	**Sys-2 GMM-IV (TL) (3)**	**Sys-2 GMM-IV (FL) (4)**	**Sys-2 GMM-IV (TL, FL) (5)**	**Sys-2 GMM-IV (TL, FL) (6)**	**Sys-2 GMM-IV (TL) (7)**	**Sys-2 GMM-IV (FL) (8)**
	Dependent variable is labor's share				**Dependent variable is Gini coefficient**			
$LABSH_{t-5}$	−0.140	−0.181**	-0.183	0.940***				
	(0.125)	(0.088)	(0.450)	(0.167)				
$GINI_{t-5}$					−0.115	−0.159	0.789***	0.689
					(0.438)	(0.286)	(0.207)	(0.451)
$TRADEOPEN_{t-5}$	−0.035**	−0.039***	−0.037**	−0.006	0.315	0.363	0.125	0.095
	(0.014)	(0.011)	(0.015)	(0.010)	(0.391)	(0.435)	(0.111)	(0.350)
$FINOPEN_{t-5}$	−0.020	0.221	−0.056	0.000	5.527	8.282	1.622	1.269
	(0.122)	(0.198)	(0.098)	(0.009)	(4.955)	(12.21)	(2.576)	(2.369)
Hansen *J*-test	[0.41]	[0.57]	[0.27]	[0.37]	[0.84]	[0.75]	[0.32]	[0.62]
AR(2) test	[0.11]	[0.61]	[0.10]	[0.20]	[0.12]	[0.36]	[0.11]	[0.12]
Wald χ^2 (prob.)	0.097	0.000	0.006	0.000	0.409	0.000	0.000	0.000
No. of inst.	15	14	15	15	11	10	9	9
No. of obs.	416	416	416	416	288	288	224	224
No. of countries	32	32	32	32	32	32	32	32

Notes: ***, **, and * denote the significance at 1%, 5%, and 10% levels, respectively. First, Sys-2 GMM-IV (TL, FL) shows the regression results for the variables using as instruments in estimations including both trade openness and financial openness. While in model (1), the standard instruments include these given variables instead of GMM-style instruments, in model (2), only the trade openness is used as a standard instrument. Second, Sys-2 GMM-IV (TL) represents the case that the regression results for given variables using as instruments in GMM-style for including only trade openness. Third, Sys-2 GMM-IV (FL) represents the case that the regression results for given variables using as instruments in GMM-style for including only financial openness. All models perform the two-step system-GMM developed by Arellano and Bover (1995) and Blundell and Bond (1998) and extended by Roodman (2006), with robust standard errors and by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. Hansen *J*-test statistics are provided to specify whether the overidentifying restrictions are valid or not. The collapsing method is used in all GMM models to get rid of from too-many-instruments problem and to provide the orthogonality conditions. The joint null hypothesis implies that the instruments are valid instruments, which means that they are uncorrelated with the error term and that the excluded instruments are correctly excluded from the estimated equation. The results of AR(2) are also provided to check for autocorrelation in case of proper maximum lag distance.

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significantly in the longer time period compared to the baseline results, which is approached by the fixed-effects method. Therefore, even if the endogeneity problem is considered to be controlled in the models, the estimation results in each model with different estimation procedures theoretically complement each other.

5.2.2 | Exponential regression models

Given the nature of the dependent variables, which are normalized between [0,1], Table 5 introduces the exponential regression results of fractional-response fixed-effects models proposed by Ramalho et al. (2018), which are based on the seminal paper of Papke and Wooldridge (2008). Besides the original fixed-effects estimators, the new fixed-effects estimators on the basis of these models are designed for linear-fractional logit and probit regression models. Different from the empirical outputs produced in Table 3, the standard specifications of these models become a form of exponential regression along with multiplicative individual-effects and time-variant heterogeneity.

One of the most technical advantages of these newly founded fixed-effects estimators are robust to both time-variant and time-invariant heterogeneity and thus can reconcile fractional responses with observations at the boundary value of zero (Ramalho et al., 2018). Moreover, the dynamic panel data models are statistically proper for new fixed-effects estimators where the endogenous explanatory variables are easily included in the regression models by way of ignoring the specification of the reduced-form model. Therefore, the empirical results obtained in Table 5 provide an informative background to the validity of regressions for dynamic models and thereby statistically confirm the initial findings.

Furthermore, the same empirical strategy is used for the longer time series in Table A7 by excluding the former socialist countries from the sample in order to statistically show that exponential regression results of fractional-response fixed-effects models are still significant, even though the sample is restricted to only a few countries.

The empirical results produced in Table 5 shows that the negative correlation between measures of openness and income distribution still prevails as in Table 3. However, the crucial difference in the results obtained by logit and probit models from the results of Table 3 is that none of the financial openness coefficients is statistically significant. In contrast, the main channel of influence in changing conditions of income distribution depends on the change in the degree of openness in the trade regime. Also, the empirical findings show that the persistence of dependent variables makes the coefficient of measures in openness insignificant. In that sense, the exclusion of lagged dependent variables from the regressions in both logit and probit models leads the coefficients of independent variables to be statistically significant. This is also supported by the fact that the number of observations remains constant as in Table 3.

Furthermore, the empirical findings imply that the channel of openness measures are more powerful on the level of income inequality measured by the Gini coefficient. Therefore, besides the class-based inequality changes, the household-based inequality changes have a significant meaning for the correlation with the openness measures, especially for the openness in the trade regime. However, the extension of time in terms of the periods change the pattern of effect of openness on income distribution as represented in Table A7. In this new framework, the channel of influence for income distribution is based on openness in a financial account. Moreover, the effect of financial openness is vigorous in the labor share of income rather than the Gini coefficient. Therefore, this extended version of the fractional-response fixed-effects method based on

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TABLE 5 Exponential regression results of fractional-response fixed-effects models

	Annual Data, 1996–2014							
	Logit models				Probit models			
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
	Dependent variable is labor's share		Dependent variable is Gini coefficient		Dependent variable is labor's share		Dependent variable is Gini coefficient	
$LABSH_{t-5}$	−0.311				−0.196			
	(0.323)				(0.200)			
$GINI_{t-5}$			0.633*				0.384*	
			(0.344)				(0.209)	
$TRADEOPEN_{t-5}$	−0.219	−0.183*	0.020	0.054**	−0.135	−0.113*	0.012	0.032**
	(0.138)	(0.102)	(0.022)	(0.021)	(0.085)	(0.063)	(0.013)	(0.012)
$FINOPEN_{t-5}$	0.030	0.020	0.008	0.028	0.018	0.012	0.005	0.017
	(0.083)	(0.075)	(0.034)	(0.034)	(0.051)	(0.046)	(0.020)	(0.020)
No. of obs.	448	448	448	448	448	448	448	448
No. of countries	32	32	32	32	32	32	32	32
R-squared	0.166	0.081	0.203	0.096	0.167	0.081	0.204	0.095

Notes: ***, **, and * denote significance at the 1%, 5% and 10% levels, respectively. Robust standard errors are in parentheses.

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logit and probit models imply that the financial account openness may have a significant effect on income distribution by way of changing the labor's share in a long time period.

All in all, dynamic system-GMM estimation results indicate that income distribution in the case of class- and household-based analyses have a mild significant and slowly changing measure. In this context, although dynamic fixed-effects results suffer from the weak instrument problem and endogeneity issue and thus the estimation results may possibly be biased and insufficient, the empirical outcomes produced from the system-GMM procedure are not very different from these initial results. Therefore, one needs further estimations in which the control variables are included in the same models and the proxy variable for openness in trade and finance is alternatively used in the same models. Section 5.3 presents a robustness check using additional variables.

5.3 | Extensions using more control variables

We assess the robustness of the empirical results to inputs from the baseline analysis with robust standard error method produced by Driscoll and Kraay (1998) in control of several variables.[20] Tables 6 and 7 present the results of the fixed-effects estimations over advanced economies. First, in the context of labor share of income for Table 6, trade openness (i.e., TRADEOPEN) is statistically significant in all regressions, which is similar to baseline and system-GMM results. However, the same is not the case for financial openness (i.e., FINOPEN) in which none of the empirical outcomes show that the coefficients are significant (excluding model 1) even though they are almost negative. Consistent with the theoretical background of the models, the bargaining power measures, including both unemployment rate (i.e., UNEMP) and human capital index (i.e., HUMCAP), are mostly significantly significant and negative, which indicates that the negative pressure on the bargaining framework of workers reduces the labor share through dispelling the social cohesion across different sectors and markets.

Additionally, the crisis dummy (i.e., CRISIS) is negative and statistically significant in all regressions showing that the overall effect of the global crisis of 2007/2008 on the labor share accruing in the national income is negative. Regarding the impact of governance indicators, which are included in political stability (POL_STAB) and rule of law (RULE_LAW), give the expected empirical results. However, though the RULE_LAW is positive and statistically significant in many of the models, the POL_STAB is negative and only statistically significant in model 8. This indicates that the rule of law is much more effective than political stability in the allocation of resources.

The study also uses other control variables for robustness tests by using foreign direct investment flows, including both inflows (FDI_INW) and outflows (FDI_OUT), and sectoral shares

[20] The major reason is to focus on an extended fixed-effects model with additional variables is based on the fact that the correlation between the openness measures and income distribution is also conditioned to other socio-economic variables. However, there are some technical problems to include all these potential variables into the system-GMM, BCFE, or fractional-response fixed-effects method in which the inclusion of these variables into the regression is limited by the selection process of endogenous and instrument variables. Therefore, the previous sections of robustness checks based on the fixed-effects method have a partial understanding of openness–distribution nexus. However, the author is aware of the limitations of using the only fixed-effects method in the presence of other control variables. Although the problems of the fixed-effects method are relevant, the empirical results still make sense in certain conditions that the estimations control the country-fixed effects to a large extent. Therefore, they are still statistically and theoretically meaningful to understand the openness–distribution nexus for advanced economies. Also, note that Hypothesis 3 is based on an extended analysis in case of inclusion of globalization variables such as the overall globalization index and economic globalization index to analyze their effects on income distribution. In that sense, the empirical results obtained in Tables 7 and 8 statistically confirm the hypothesized proposition for the case of the effects of globalization indices on income distribution.

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TABLE 6 Extensions to base specifications (dependent variable is lag of labor's share)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Dependent variable is the lag of labor's share										
TRADEOPEN	−0.052***	−0.042***	−0.031***	−0.037***	−0.036***	−0.032***					
	(0.007)	(0.006)	(0.010)	(0.009)	(0.008)	(0.008)					
FINOPEN	−0.016*	−0.016	−0.008	−0.006	0.001	0.004					
	(0.009)	(0.010)	(0.008)	(0.007)	(0.007)	(0.006)					
FDI_INW		0.000	0.000	0.000	0.000	0.000					
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)					
FDI_OUT		−0.000***	−0.000***	−0.000***	−0.000***	−0.000**					
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)					
ECONGLOB							−0.000	−0.000			
							(0.000)	(0.000)			
OVRGLOB									−0.002***	−0.002***	−0.002***
									(0.001)	(0.000)	(0.000)
UNEMP		0.000	0.000	−0.000	−0.001*	−0.000	−0.000	-0.001	−0.001	−0.001**	−0.001
		(0.001)	(0.001)	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
HUMCAP			−0.043***	−0.052***	−0.059***	−0.056***	−0.068***	-0.083***	−0.038***	−0.044***	−0.044***
			(0.010)	(0.011)	(0.013)	(0.011)	(0.009)	(0.011)	(0.009)	(0.011)	(0.010)
CRISIS			−0.007***	−0.007***	−0.007***	−0.010***	−0.012***	-0.013***	−0.011***	−0.011***	−0.011***
			(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
GOVSHARE				0.177**	0.155*	0.134*		0.097	0.057	0.030	0.018
				(0.076)	(0.077)	(0.074)		(0.070)	(0.077)	(0.076)	(0.080)

(Continues)

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TABLE 6 (Continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Dependent variable is the lag of labor's share										
INDUSTRY				0.000				-0.000	−0.000		
				(0.000)				(0.000)	(0.000)		
MANUFACTURE					0.001***					0.001*	
					(0.000)					(0.001)	
SERVICE						−0.001*					−0.002***
						(0.000)					(0.000)
RULE_LAW			0.028	0.044***	0.042***	0.038***	0.023	0.035**			
			(0.016)	(0.014)	(0.011)	(0.013)	(0.015)	(0.014)			
POL_STAB				−0.011	−0.009	−0.009		−0.016**			
				(0.009)	(0.008)	(0.008)		(0.006)			
GDPGRW					−0.002***					−0.003**	
					(0.001)					(0.001)	
INFLATION						0.001					0.001
						(0.001)					(0.001)
R-squared	0.147	0.139	0.191	0.226	0.270	0.237	0.149	0.178	0.169	0.213	0.196
No. of obs.	576	565	565	552	552	552	576	563	563	563	563
No. of countries	32	32	32	32	32	32	32	32	32	32	32

Notes: ***significant at 1%; **significant at 5%; *significant at 10%. Robust standard errors are in parentheses.

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TABLE 7 Extensions to base specifications (dependent variable is lag of Gini coefficient)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Dependent variable is lag Gini coefficient										
TRADEOPEN	1.933***	1.852***	0.823***	0.761***	0.744***	0.824***					
	(0.259)	(0.312)	(0.217)	(0.198)	(0.207)	(0.211)					
FINOPEN	0.037	0.111	−0.312	−0.223	−0.251	−0.205					
	(0.394)	(0.376)	(0.397)	(0.450)	(0.441)	(0.403)					
FDI_INW		0.002	0.001	0.002	0.001	0.001					
		(0.002)	(0.002)	(0.002)	(0.002)	(0.002)					
FDI_OUT		−0.001	0.000	0.000	0.000	0.001					
		(0.004)	(0.003)	(0.003)	(0.003)	(0.003)					
ECONGLOB							0.052***	0.050***			
							(0.009)	(0.011)			
OVRGLOB									0.099***	0.100***	0.110***
									(0.011)	(0.011)	(0.011)
UNEMP		−0.002	0.002	0.004	0.006	0.001	0.016	0.016	0.009	0.012	0.019
		(0.012)	(0.015)	(0.017)	(0.019)	(0.023)	(0.017)	(0.017)	(0.021)	(0.023)	(0.028)
HUMCAP			2.718***	2.658***	2.723***	2.561***	2.577***	2.527***	1.275***	1.337***	1.234***
			(0.471)	(0.478)	(0.519)	(0.477)	(0.297)	(0.336)	(0.334)	(0.357)	(0.306)
CRISIS			0.113**	0.156**	0.157**	0.146	0.104*	0.158**	0.065	0.065	−0.014
			(0.049)	(0.067)	(0.064)	(0.094)	(0.057)	(0.064)	(0.062)	(0.062)	(0.058)
GOVSHARE				2.974	3.208*	2.539		4.007*	3.761*	4.138**	3.577*
				(1.922)	(1.765)	(2.049)		(1.999)	(1.886)	(1.856)	(2.038)

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TABLE 7 (Continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Dependent variable is lag Gini coefficient										
INDUSTRY				0.005				0.004	0.009		
				(0.007)				(0.006)	(0.006)		
MANUFACTURE					−0.003					−0.002	
					(0.007)					(0.007)	
SERVICE						−0.005					0.016
						(0.020)					(0.019)
RULE_LAW			0.263	0.321	0.348	0.263	0.213	0.293			
			(0.411)	(0.335)	(0.313)	(0.326)	(0.398)	(0.331)			
POL_STAB				0.212	0.193	0.236		0.282			
				(0.129)	(0.134)	(0.143)		(0.174)			
GDPGRW					0.018					0.025	
					(0.022)					(0.017)	
INFLATION						−0.000					0.040
						(0.036)					(0.028)
R-squared	0.141	0.118	0.211	0.208	0.209	0.207	0.242	0.244	0.259	0.262	0.263
No. of obs.	576	565	565	552	552	552	576	563	563	563	563
No. of countries	32	32	32	32	32	32	32	32	32	32	32

Notes: ***significant at 1%; **significant at 5%; *significant at 10%. Robust standard errors are in parentheses.

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such as industry (i.e., INDUSTRY), manufacturing (i.e., MANUFACTURE), and services (i.e., SERVICE) shares as part of GDP, which are all added into the empirical models. On the one hand, as expected in the theoretical sense, FDI_OUT is negative and statistically significant in all regressions. However, FDI_INW is insignificant although the coefficient is positive. On the other hand, the measures on sectoral shares give controversial results. While the coefficients of the industrial and service sectors are negatively correlated with labor's share, the same is not true for the coefficient of manufacturing sector. In each regression results, the coefficients of these variables are almost statistically significant. Although the understanding behind the service sector for why it has a negative impact on labor's share is clear since most of the low-skilled workers are employed in this sector and thus the wages are close to the reservation level. Additionally, the level of employment is so high in the service sector across advanced countries, which thus creates negative pressure on income distribution in detriment to the labor share. However, the controversial case depends on the sign of the coefficients of industrial and the manufacturing sectors. One of the major reasons why the industrial sector negatively affects the labor share of income is that if the production process in the industry is automated along with use of machines, the general employment strategy needs high-skilled workers. Therefore, if there is an imbalance in the number of workers in terms of their skills, the result is a decline of labor's share. This is the possible reason behind the positive coefficient of MANUFACTURE in which the production method may have not automated in comparison with the industrial sector.

The study implements further robustness tests in terms of the Gini coefficient to investigate the effect of economic openness variables on income inequality and the results are summarized in Table 7. In particular, it uses the same models and the econometric procedure as presented in Table 6. The most striking result comes from the TRADEOPEN measure in which the coefficient of this variable is highly statistically significant and positive in all models. Thus, a higher degree of openness in trade channel increases the level of inequality in the context of household-based analysis across advanced economies. This result also indicates that trade openness is the most effective channel for the change in income distribution over time. However, the study cannot make the same conclusion for the degree of openness in financial accounts in searching for its effect on distributional issues between capital and labor, and also among different income-levels of households. As presented in Table 6, the coefficient of FINOPEN is statistically insignificant in Table 7. Therefore, the handicap for enhanced solidarity comes from the side of trade regime effects on income distribution in control of several variables. Additionally, the same insignificant conditions are prevailing for FDI flows, including both inflows and outflows. The bargaining framework is also changed in the context of using the Gini coefficient. In Table 7, however, UNEMP becomes insignificant in explaining the change in allocational problems. However, HUMCAP is still statistically significant in all models and positively correlated with the Gini coefficient. Furthermore, the significant characteristics of sectoral variables turn into unaffected positions in the case of their impact on income distribution in Table 7. Finally, both economic globalization (i.e., ECONGLOB) and overall globalization (i.e., OVRGLOB) coefficients are statistically significant and positive in all regressions, which indicates that the negative effect of globalization channel on income inequality is completely straightforward. All in all, the estimation results lead us to argue that the class-based determinants of income distribution are very different from the household-based components, even though the impact of trade openness is effective in a mutual framework. Therefore, it is clear that both sub-components and the major components of economic openness are one of the major determinants in the case of negatively affected income distribution across advanced economies over the 1996–2014 period.

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6 | CONCLUDING REMARKS

The principal theme of social solidarity basically depends on the sources of different disciplines; and therefore, the social order in society can differ due to multiple factors. While some of the studies often encountered in Western countries investigate the historical movement of solidarity in terms of capitalist development, others focus on the situations for under-capitalist countries. However, the allocational or more preciously distributional consequences of the changes in the social system over the globalization process for the post-1980s period have been somewhat neglected in the literature, especially in the economic discipline. One of the reasons is the lack of foundations in explaining the relationship between the changes in distributional mechanisms and social solidarity. Therefore, this paper is based on the hypothesis that the income distribution as a whole is mainly determined through changes in the bargaining framework in the globalized world, which thus alter the focal dynamics of social solidarity and social cohesion. In other words, the main rationale of this study is founded on the fact that if the bargaining positions of the social classes change over time, this then leads to direct changes in the distribution of income and indirect changes in the social system. In that vein, the paper has argued that in contrast to traditional wisdom, globalization of the major economic channels, such as the trade regime and financial accounts, impedes equal distribution of income since they too much favor capital over labor through decreasing the bargaining power of labor and thus erode social solidarity, especially in the working environment.

In particular, the income distribution is analyzed by way of two variables, which are the labor share of income and the Gini coefficient in order to consider both class-based and household-based analyses, respectively. First, it can be specified that the relationship between economic openness and income distribution is straightforward over time. However, one thing that should be mentioned is that the effect of financial channel on the measures of income distribution is almost insignificant in contrast to the same effect of the trade channel. In other words, the degree of trade openness is significantly effective on the distribution of income across advanced economies whereas the level of financial transactions and relations do not have a considerable effect in most of the specifications. In that sense, more open trade regimes are both effective in the reduction of labor's share accruing in national income in favor of the capital and in an increase of income inequality in different quintiles of income groups based on households. Although the effects of openness are differentiated in two types of analyses, it is not possible to reject the causality for the economic openness–income distribution nexus in control of several econometric procedures. Therefore, the estimation results show that the dynamic effects of openness on income distribution may provide different outputs in terms of measuring the income distribution in which the nexus among these two variables need much attention for further studies through benefiting from the other disciplines. Each variable has its own characteristics which are also effective on social order and thus have its own power to affect the allocation of resources and income distribution separately.

Although this paper has shown that there is a causal relationship between economic openness and income distribution and that it changes on the basis of methodological differences of each measurement of income distribution, the same correlations are also significant and robust in control of several variables, which reveals the inner dynamics of the potential relationship among these two indicators through pointing out the importance of the bargaining framework. In that sense, the striking fact of the empirical results is the negative effects of bargaining power—proxied by unemployment rate and human capital index—on income distribution. Actually, this was an

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expected result since a decrease in the bargaining power of labor is the main stimulus behind the breakdown of social solidarity and social cohesion. Although the increase in unemployment rate reduces the power to bargain on higher wages, the increase in the level of human capital favors the wages of high-skilled workers in detriment of low-skilled labor and thus widens the wage dispersion across different sectors and negatively affects the income distribution.

Furthermore, besides the negative influence of measures on the bargaining framework, the combined globalization indices, such as economic, political and social indicators, are negatively correlated with income inequality. On the one hand, the economic globalization index has negative effect on income distribution variables. First, the index has a downward pressure on labor's share by way of reducing the bargaining power of workers, especially the low-skilled labor. Therefore, it widens the gap between capital and labor share and distorts the accord among these two classes through eroding the social cohesion across the national boundaries. Second, income inequality on the basis of household income analysis is also increased due to the implementation of more openness in the economic sphere. On the other hand, further robustness checks for the effects of globalization movement on income distribution are also tested by the overall globalization index, which comprises both economic, political and social factors. The results are the same as represented in the models measured by economic globalization index. The coefficients of overall globalization index are highly statistically significant in all models, which mean that there is a strong relationship between the allocation of resources/distribution of income and the globalization indicators.

All in all, the context of this paper initiates a multi-dimensional investigation for the economic openness–income distribution nexus. In that vein, in addition to the indicators for economic openness, the historical movements and changes in the labor market on the basis of bargaining framework included into the current analysis and thus the differentials in social solidarity and cohesion was examined through the changes in bargaining power of labor, which contributed to our understanding of the inner dynamics of social segmentation on the basis of income levels across advanced economies. However, the paper also points out that the dimensions of the relationship between economic openness and income distribution and thereby the social solidarity need further examination by way of looking at other variables from the other disciplines. Therefore, there remains a need for further analyses for further understanding of extended generalizations about the economic openness–income distribution nexus.

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APPENDIX

TABLE A1 Variables, definitions and sources*

Variable	Definition	Source
Labor's share	Compensation of employees / employees, persons	Penn World Tables 9.0
Gini coefficient	Gini coefficient (0 = fully equal; 100 = fully unequal)	Solt (2016)
Trade openness	(Exports + Imports)/GDP]	World Development Indicators, Penn World Tables 9.0, Author's Calculation
Financial openness	Standardized principal component of the variables that indicate the presence of multiple exchange rates, restrictions on current account transactions, on capital account transactions, and the requirement of the surrender of export proceeds.	The Trilemma Indexes, Aizenman (2018)
Foreign direct investment, inwards	Foreign direct investment: inwards flows (% of GDP)	World Development Indicators
Foreign direct investment, outflows	Foreign direct investment: outwards flows (% of GDP)	World Development Indicators
Economic globalization	The index of economic globalization which covers both trade globalization and financial globalization indicators	The KOF Globalisation Index Database
Overall globalization	The index of overall globalization which covers both economic, political and social globalization indices	
Unemployment rate	Unemployment rate (% of total labor force)	World Development Indicators
Human capital	Human capital index, based on years of schooling and returns to education	Penn World Tables 9.0
Crisis	Crisis (Dummy)	Author's Calculation
Government share	Share of government consumption at current PPPs	Penn World Tables 9.0

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TABLE A1 Variables, definitions and sources*

Variable	Definition	Source
Industry growth rate	Annual growth rate for industrial value added based on constant local currency. Aggregates are based on constant 2010 U.S. dollars. Industry corresponds to ISIC divisions 10–45 and includes manufacturing (ISIC divisions 15–37). It comprises value added in mining, manufacturing (also reported as a separate subgroup), construction, electricity, water, and gas. Value added is the net output of a sector after adding up all outputs and subtracting intermediate inputs. It is calculated without making deductions for depreciation of fabricated assets or depletion and degradation of natural resources. The origin of value added is determined by the International Standard Industrial Classification (ISIC), revision 3.	World Development Indicators
Manufacturing growth rate	Annual growth rate for manufacturing value added based on constant local currency. Aggregates are based on constant 2010 U.S. dollars. Manufacturing refers to industries belonging to ISIC divisions 15–37. Value added is the net output of a sector after adding up all outputs and subtracting intermediate inputs. It is calculated without making deductions for depreciation of fabricated assets or depletion and degradation of natural resources. The origin of value added is determined by the International Standard Industrial Classification (ISIC), revision 3.	World Development Indicators
Services growth rate	Annual growth rate for value added in services based on constant local currency. Aggregates are based on constant 2010 U.S. dollars. Services correspond to ISIC divisions 50–99. They include value added in wholesale and retail trade (including hotels and restaurants), transport, and government, financial, professional, and personal services such as education, health care, and real estate services. Also included are imputed bank service charges, import duties, and any statistical discrepancies noted by national compilers as well as discrepancies arising from rescaling. Value added is the net output of a sector after adding up all outputs and subtracting intermediate inputs. It is calculated without making deductions for depreciation of fabricated assets or depletion and degradation of natural resources. The industrial origin of value added is determined by the International Standard Industrial Classification (ISIC), revision 3 or 4.	World Development Indicators
Rule of law	Rule of Law captures perceptions of the extent to which agents have confidence in and abide by the rules of society, and in particular the quality of contract enforcement, property rights, the police, and the courts, as well as the likelihood of crime and violence. Estimate gives the country's score on the aggregate indicator, in units of a standard normal distribution, i.e. ranging from approximately -2.5 to 2.5.	Worldwide Governance Indicators

Continues

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TABLE A1 Continued

Variable	Definition	Source
Political Stability and Absence of Violence/Terrorism	Political Stability and Absence of Violence/Terrorism measures perceptions of the likelihood of political instability and/or politically-motivated violence, including terrorism. Estimate gives the country's score on the aggregate indicator, in units of a standard normal distribution, i.e. ranging from approximately -2.5 to 2.5.	Worldwide Governance Indicators
GDP per capita growth	Annual percentage growth rate of GDP per capita based on constant local currency. Aggregates are based on constant 2010 U.S. dollars. GDP per capita is gross domestic product divided by midyear population. GDP at purchaser's prices is the sum of gross value added by all resident producers in the economy plus any product taxes and minus any subsidies not included in the value of the products. It is calculated without making deductions for depreciation of fabricated assets or for depletion and degradation of natural resources.	World Development Indicators
Inflation, consumer prices	Inflation as measured by the consumer price index reflects the annual percentage change in the cost to the average consumer of acquiring a basket of goods and services that may be fixed or changed at specified intervals, such as yearly. The Laspeyres formula is generally used.	World Development Indicators

*** All of these definitions for variables are quoted from their related database.

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TABLE A2 List of countries and country codes

Country	Codes
Australia	AU
Austria	AT
Belgium	BE
Canada	CA
Czech Republic	CZ
Denmark	DK
Estonia	EE
Finland	FI
France	FR
Germany	DE
Greece	GR
Iceland	IS
Ireland	IE
Israel	IL
Italy	IT
Japan	JP
Korea Republic	KR
Latvia	LV
Lithuania	LT
Luxembourg	LU
Netherlands	NL
New Zealand	NZ
Norway	NO
Portugal	PT
Singapore	SG
Slovakia	SK
Slovenia	SI
Spain	ES
Sweden	SE
Switzerland	CH
United Kingdom	UK
United States	US

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TABLE A3 Baseline estimation results (extensions to time span, 1985–2014 period)

	Annual data, 1985–2014									
	Pooled OLS (1)	**FE OLS (2)**	**BCFE Dynamic (3)**	**Sys-2 GMM (4)**	**FE OLS (5)**	**Pooled OLS (6)**	**FE OLS (7)**	**BCFE Dynamic (8)**	**Sys-2 GMM (9)**	**FE OLS (10)**
	Dependent variable is labor's share					**Dependent variable is Gini coefficient**				
$LABSH_{t-5}$	0.863***	0.277***	1.038***	0.726***						
	(0.032)	(0.040)	(0.052)	(0.101)						
$GINI_{t-5}$						0.946***	0.445***	1.342***	0.996***	
						(0.018)	(0.090)	(0.047)	(0.030)	
$TRADEOPEN_{t-5}$	−0.004	0.016	−0.007*	−0.011*	−0.023***	−0.033	−0.619	−0.067	−0.242	1.971***
	(0.002)	(0.010)	(0.004)	(0.006)	(0.006)	(0.161)	(0.910)	(0.117)	(0.155)	(0.344)
$FINOPEN_{t-5}$	0.008	−0.007	−0.001	0.044*	−0.033***	−0.048	−0.405	−0.115	0.270	0.919
	(0.007)	(0.010)	(0.003)	(0.023)	(0.004)	(0.457)	(0.549)	(0.118)	(0.526)	(0.241)
Hansen *J*-test				[0.11]					[0.37]	
AR(2) test				[0.10]					[0.29]	
No. of inst.				20					12	
No. of obs.	600	600	600	216	600	600	600	600	120	600
No. of countries	24	24	24	24	24	24	24	24	24	24
R-squared	0.861	0.916			0.880	0.947	0.971			0.946

Notes: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Pooled OLS regression in columns (1) and (6), with robust standard errors clustered by country are in parentheses. Fixed-effects (FE) OLS regressions in columns (2), (5), (7), and (10), with country dummies and robust standard errors clustered by country are in parentheses. Columns (3) and (8) use the bootstrap-corrected fixed effects (BCFE) method of De Vos et al. (2015) with robust standard errors, and columns (4) and (9) use the two-step system-GMM (Sys-2 GMM) of Arellano and Bover (1995) and of Blundell and Bond (1998) extended by Roodman (2006), with robust standard errors and by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. Year dummies are included in all pooled OLS, fixed-effects, and system-GMM regressions. In columns (4) and (9), the regression results are also provided Hansen *J*-test statistics to specify whether the overidentifying restrictions are valid or not. The collapsing method is used in system-GMM models to get rid of from too-many-instruments problem and to provide the orthogonality conditions. The joint null hypothesis implies that the instruments are valid, which means that they are uncorrelated with the error term and that the excluded instruments are correctly ignored from the estimated equation. The results of AR(2) are also provided to check for autocorrelation in case of proper maximum lag distance. For the case of the BCFE method representing in models (3) and (8), the selection of the error resampling pattern is based on cross-sectional heteroskedasticity error terms and the analytical heterogeneous initialization is used in the regression. In addition, the convergence option is selected in all models. Standard errors are estimated from the bootstrap distribution of the BCFE estimator. Confidence bounds for the *t*-distribution are calculated on the basis of bootstrapped standard errors.

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TABLE A4 Baseline estimation results (1996–2005 period)

	Annual data, 1996–2005									
	Pooled OLS (1)	FE OLS (2)	BCFE Dynamic (3)	Sys-2 GMM (4)	FE OLS (5)	Pooled OLS (6)	FE OLS (7)	BCFE Dynamic (8)	Sys-2 GMM (9)	FE OLS (10)
	Dependent variable is labor's share					Dependent variable is Gini coefficient				
$LABSH_{t-5}$	0.964***	0.494***	0.746***	0.971***						
	(0.016)	(0.064)	(0.159)	(0.016)						
$GINI_{t-5}$						0.989***	0.926***	1.291***	0.986***	
						(0.006)	(0.042)	(0.121)	(0.012)	
$TRADEOPEN_{t-5}$	−0.001*	−0.005	−0.056***	−0.002*	−0.015	0.016	0.224	−0.030	0.016	0.846**
	(0.000)	(0.013)	(0.021)	(0.001)	(0.011)	(0.027)	(0.317)	(0.211)	(0.038)	(0.396)
$FINOPEN_{t-5}$	0.002	0.003	0.077**	0.001	−0.027**	−0.204*	−0.020	−0.612*	−0.117	1.365***
	(0.003)	(0.006)	(0.037)	(0.005)	(0.010)	(0.100)	(0.148)	(0.368)	(0.099)	(0.387)
Hansen *J*-test				[0.17]					[0.16]	
AR(2) test				[0.49]					[0.28]	
No. of inst.				12					12	
No. of obs.	288	288	160	160	288	288	288	160	160	288
No. of countries	32	32	32	32	32	32	32	32	32	32
R-squared	0.945	0.965			0.939	0.997	0.998			0.984

Notes: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Pooled OLS regression in columns (1) and (6), with robust standard errors clustered by country are in parentheses. Fixed-effects (FE) OLS regressions in columns (2), (5), (7), and (10), with country dummies and robust standard errors clustered by country are in parentheses. Columns (3) and (8) use the bootstrap-corrected fixed effects (BCFE) method of De Vos et al. (2015) with robust standard errors, and columns (4) and (9) use the two-step system-GMM (Sys-2 GMM) of Arellano and Bover (1995) and of Blundell and Bond (1998) extended by Roodman (2006), with robust standard errors and by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. Year dummies are included in all pooled OLS, fixed-effects, and system-GMM regressions. In columns (4) and (9), the regression results are also provided Hansen *J*-test statistics to specify whether the overidentifying restrictions are valid or not. The collapsing method is used in system-GMM models to get rid of from too-many-instruments problem and to provide the orthogonality conditions. The joint null hypothesis implies that the instruments are valid, which means that they are uncorrelated with the error term and that the excluded instruments are correctly ignored from the estimated equation. The results of AR(2) are also provided to check for autocorrelation in case of proper maximum lag distance. For the case of the BCFE method representing in models (3) and (8), the selection of the error resampling pattern is based on cross-sectional heteroskedasticity error terms and the analytical heterogeneous initialization is used in the regression. In addition, the convergence option is selected in all models. Standard errors are estimated from the bootstrap distribution of the BCFE estimator. Confidence bounds for the *t*-distribution are calculated on the basis of bootstrapped standard errors.

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TABLE A5 Baseline estimation results (2006–2014 period)

	Annual data, 2006–2014									
	Pooled OLS (1)	FE OLS (2)	BCFE Dynamic (3)	Sys-2 GMM (4)	FE OLS (5)	Pooled OLS (6)	FE OLS (7)	BCFE Dynamic (8)	Sys-2 GMM (9)	FE OLS (10)
	Dependent variable is labor's share					Dependent variable is Gini coefficient				
$LABSH_{t-5}$	0.988***	0.630***	0.817***	0.983***						
	(0.014)	(0.057)	(0.131)	(0.018)						
$GINI_{t-5}$						0.996***	0.847***	1.351***	0.997***	
						(0.004)	(0.031)	(0.168)	(0.004)	
$TRADEOPEN_{t-5}$	0.001	−0.002	−0.016	−0.001	−0.022**	0.005	0.232	0.045	0.003	0.563**
	(0.000)	(0.010)	(0.013)	(0.001)	(0.009)	(0.018)	(0.251)	(0.257)	(0.032)	(0.246)
$FINOPEN_{t-5}$	0.001	0.014	0.032	−0.007*	0.018	0.250*	0.522	0.579	0.146*	0.849*
	(0.003)	(0.013)	(0.069)	(0.004)	(0.017)	(0.148)	(0.450)	(1.608)	(0.077)	(0.475)
Hansen *J*-test				[0.82]					[0.29]	
AR(2) test				[0.11]					[0.10]	
No. of inst.				12					10	
No. of obs.	256	256	128	160	256	256	256	128	128	256
No. of countries	32	32	32	32	32	32	32	32	32	32
R-squared	0.963	0.973			0.943	0.997	0.998			0.991

Notes: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Pooled OLS regression in columns (1) and (6), with robust standard errors clustered by country are in parentheses. Fixed-effects (FE) OLS regressions in columns (2), (5), (7), and (10), with country dummies and robust standard errors clustered by country are in parentheses. Columns (3) and (8) use the bootstrap-corrected fixed effects (BCFE) method of De Vos et al. (2015) with robust standard errors, and columns (4) and (9) use the two-step system-GMM (Sys-2 GMM) of Arellano and Bover (1995) and of Blundell and Bond (1998) extended by Roodman (2006), with robust standard errors and by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. Year dummies are included in all pooled OLS, fixed-effects, and system-GMM regressions. In columns (4) and (9), the regression results are also provided Hansen *J*-test statistics to specify whether the overidentifying restrictions are valid or not. The collapsing method is used in system-GMM models to get rid of from too-many-instruments problem and to provide the orthogonality conditions. The joint null hypothesis implies that the instruments are valid, which means that they are uncorrelated with the error term and that the excluded instruments are correctly ignored from the estimated equation. The results of AR(2) are also provided to check for autocorrelation in case of proper maximum lag distance. For the case of the BCFE method representing in models (3) and (8), the selection of the error resampling pattern is based on cross-sectional heteroskedasticity error terms and the analytical heterogeneous initialization is used in the regression. In addition, the convergence option is selected in all models. Standard errors are estimated from the bootstrap distribution of the BCFE estimator. Confidence bounds for the *t*-distribution are calculated on the basis of bootstrapped standard errors.

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TABLE A6 System-GMM estimation results (extensions to time span, 1985–2014 period)

	Annual data, 1985–2014							
	Sys-2 GMM-IV (TL, FL) (1)	Sys-2 GMM-IV (TL, FL) (2)	Sys-2 GMM-IV (TL) (3)	Sys-2 GMM-IV (FL) (4)	Sys-2 GMM-IV (TL, FL) (5)	Sys-2 GMM-IV (TL, FL) (6)	Sys-2 GMM-IV (TL) (7)	Sys-2 GMM-IV (FL) (8)
	Dependent variable is labor's share				Dependent variable is Gini coefficient			
$LABSH_{t-5}$	0.269	0.066	0.167	0.905**				
	(0.399)	(0.143)	(0.235)	(0.132)				
$GINI_{t-5}$					−0.043	−0.117	0.917***	0.646
					(0.546)	(0.220)	(0.137)	(0.450)
$TRADEOPEN_{t-5}$	−0.032	−0.040***	−0.039**	−0.001	0.617	0.718	−0.095	0.219
	(0.018)	(0.011)	(0.014)	(0.006)	(1.511)	(1.200)	(0.459)	(1.118)
$FINOPEN_{t-5}$	0.059	0.057	0.076**	0.010	−2.344	−1.343	−0.633	−2.329
	(0.038)	(0.109)	(0.037)	(0.018)	(2.813)	(4.883)	(0.850)	(2.900)
Hansen *J*-test	[0.44]	[0.35]	[0.23]	[0.50]	[0.79]	[0.75]	[0.39]	[0.89]
AR(2) test	[0.10]	[0.12]	[0.11]	[0.10]	[0.25]	[0.18]	[0.11]	[0.11]
Wald χ^2 (prob.)	0.000	0.000	0.000	0.000	0.598	0.779	0.000	0.064
No. of inst.	22	21	22	22	12	11	8	10
No. of obs.	480	480	480	480	240	240	144	192
No. of countries	24	24	24	24	24	24	24	24

Notes: ***, **, and * denote the significance at 1%, 5%, and 10% levels, respectively. First, Sys-2 GMM-IV (TL, FL) shows the regression results for the variables using as instruments in estimations including both trade openness and financial openness. While in model (1), the standard instruments include these given variables instead of GMM-style instruments, in model (2), only the trade openness is used as a standard instrument. Second, Sys-2 GMM-IV (TL) represents the case that the regression results for given variables using as instruments in GMM-style for including only trade openness. Third, Sys-2 GMM-IV (FL) represents the case that the regression results for given variables using as instruments in GMM-style for including only financial openness. All models perform the two-step system-GMM developed by Arellano and Bover (1995) and Blundell and Bond (1998) and extended by Roodman (2006), with robust standard errors and by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix. Hansen *J*-test statistics are provided to specify whether the overidentifying restrictions are valid or not. The collapsing method is used in all GMM models to get rid of from too-many-instruments problem and to provide the orthogonality conditions. The joint null hypothesis implies that the instruments are valid instruments, which means that they are uncorrelated with the error term and that the excluded instruments are correctly excluded from the estimated equation. The results of AR(2) are also provided to check for autocorrelation in case of proper maximum lag distance.

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TABLE A7 Exponential regression results of fractional-response fixed-effects models (extensions to time span, 1985–2014 period)

	Annual data, 1985–2014							
	Logit models				Probit models			
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
$LABSH_{t-5}$	1.466***				0.909***			
	(0.249)				(0.154)			
$GINI_{t-5}$			2.645***				1.586***	
			(0.391)				(0.235)	
$TRADEOPEN_{t-5}$	−0.010	−0.103	0.020	0.104	−0.005	−0.063	0.011	0.061
	(0.032)	(0.068)	(0.035)	(0.066)	(0.019)	(0.042)	(0.020)	(0.039)
$FINOPEN_{t-5}$	−0.106***	−0.134**	0.017	0.046	−0.066***	−0.083**	0.010	0.027
	(0.031)	(0.048)	(0.018)	(0.038)	(0.019)	(0.030)	(0.011)	(0.023)
No. of obs.	600	600	600	600	600	600	600	600
No. of countries	24	24	24	24	24	24	24	24
R-squared	0.263	0.142	0.464	0.130	0.263	0.142	0.465	0.128

Notes: ***, **, and * denote significance at the 1%, 5% and 10% levels, respectively. Robust standard errors are in parentheses.

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