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Capital Inflows and Income Inequality: Evidence from Panel VAR Approach

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Abstract

In this paper, I document empirical evidence that an external shock in capital inflows leads to an increase in income inequality in advanced economies and causes a decline in income inequality in emerging market economies. I estimate a panel VAR model with an annual dataset on 53 countries over the period 1990-2020 to study the effects of capital inflows on income inequality within countries. To distinguish the external capital inflow shocks driven by global financial conditions from other shocks, I identify the structural external shocks to capital inflows using sign restrictions. The analysis is performed separately in advanced and emerging market economies since the two groups show significant differences in the level of economic development and the degree of capital market openness. The results are statistically and economically significant. By income class, a capital inflow shock increases primarily the income share of the rich in advanced economies and the poorest half in emerging market economies. These empirical findings suggest that capital inflows have different impacts on income inequality across countries, and policymakers should pay attention to the possibility of adverse distributional effects of capital inflows.

Keywords: Capital inflows, Income inequality, Panel VAR, Sign restrictions
JEL codes: D63, F32, F38

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1 Introduction

Over the past decades, global inequality between countries has decreased since many developing countries have rapidly grown due to globalization and the spillover of technology and knowledge. By contrast, income inequality within countries has risen in many developed and developing countries. It has sparked intense debates about the causes and dynamics of within-country inequality, which has also induced concerns among economists and policymakers.

Among the various factors for the rising income inequality within countries, much of the literature discusses globalization or international integration as one of the main factors.[1] International integration means a deeper integration of a domestic economy into the global market and can be represented by growing transactions between domestic and foreign actors (or residents and non-residents). Progress in international integration has developed in two dimensions. One is trade integration, and the other is financial integration, characterized by an increase in the magnitude of international capital movements. International financial integration has evolved relatively recently compared to trade integration, especially since the early 1990s.[2] Since then, many countries have seen massive growth in cross-border capital movements.

According to many economists, capital inflows lead to an increase in economic welfare in recipient countries. These benefits from foreign capital are associated with boosting economic growth (Prasad et al., 2005; Henry, 2007), consumption smoothing (Kose et al., 2009), and enhancing the efficiency of domestic financial markets. On the other hand, surges and subsequent reversals of capital inflows can hurt macroeconomic stability of a recipient country and trigger a financial crisis (Caballero, 2016; Ghosh et al., 2016). In recent years, economists have paid more attention to the distributional effect of capital inflows, i.e., how increased capital inflows affect income inequality within countries. In contrast to trade integration, however, the distributional effects of financial integration have been studied relatively little. Theoretical predictions and empirical evidence about the impacts of capital inflows on income inequality are still ambiguous.

[1] Researchers skeptical of the role of international integration in this debate argue that Skill-biased technological change (SBTC) has caused chiefly an increase in income inequality. However, SBTC does not explain the evolution of inequality after the 1990s. This paper does not discuss it in detail. See Card and DiNardo (2002) for details.

[2] The opening of the capital account and the associated deregulation mainly took place in the 1990s. Furceri and Loungani (2018) document that 44.6 % (100 episodes) of all capital market liberalization episodes (224 episodes over the period 1970-2010) occurred during the period in the 1990s, and 27.2 % (61 episodes) occurred in the 2000s. Furthermore, the development of information and communication technology since the 1990s, making external financial transactions more effortless than ever, may also be a major factor in the accelerated integration of international financial markets and the rapid growth in cross-border capital flows.

In this paper, I investigate the distributional effects of capital inflows, driven by changes in global financial conditions, *not* driven by domestic economic conditions. For this purpose, I employ a panel vector autoregression (VAR) model with a country-level panel dataset and identify external shocks to capital inflows using sign restrictions proposed by Faust (1998), Canova and Nicolo (2002) and Uhlig (2005).

The present paper contributes to the literature on the macroeconomic impacts of cross-border capital inflows, particularly distributional effects, in several ways. First, to the best of my knowledge, this paper is the first to employ a panel VAR approach to study the dynamic effects of capital inflows on income inequality. The existing literature investigates the impact of financial integration on income inequality using dynamic panel regressions (Asteriou et al., 2014; Bumann & Lensink, 2016; de Haan & Sturm, 2017; Furceri & Loungani, 2018; Furceri et al., 2019; Erauskin & Turnovsky, 2019; Li & Su, 2021), but not a panel VAR for that purpose. A VAR model is best suited when there is a lack of theoretical information about the dynamic relationships between economic variables or no theoretical consensus exists to guide the model specification. Also, a VAR model is an appropriate methodology for addressing the potential endogeneity of variables by treating the interesting variables as endogenous. Moreover, a panel VAR model incorporating a panel analysis methodology into a VAR framework allows for including country-specific fixed effects, which capture time-invariant factors that may affect endogenous variables in the multivariate system.

Second, I identify exogenous capital inflow shocks, thus recognizing that capital flows are driven by both global "push factors" and domestic "pull factors". In order to identify an exogenous capital inflow shock, I adopt sign restrictions following the seminal work of Uhlig (2005). I assume that an exogenous capital inflow shock raises world GDP and lowers the U.S. and domestic interest rates, while a domestic (i.e., "pull factors") shock does not. The most popular approach for identifying structural shocks in a VAR model is a recursive methodology that uses the Cholesky decomposition of the variance-covariance matrix of the error terms in the estimated VAR. However, the fact that the ordering of the dependent variables can affect the outcome and that the ordering can be arbitrary is a primary criticism of this methodology. By using sign restrictions, this paper avoids arbitrary ordering of the variables and identifies the orthogonal shock of capital inflows to other variables.

Finally, I divide countries into two groups, advanced economies (AE) and emerging market economies (EME), according to the International Monetary Fund (IMF) classification of countries, and compare the distributional effects of capital inflows for these two countries groups. While AE have tended to embrace international financial integration earlier, EME have opened their capital account late and integrated into the international financial market to a relatively low degree. Furthermore, AE and EME show significant differences in the level of economic development. Therefore, it may not be reasonable to analyze the two groups jointly. Many relevant studies have focused only on one of these groups or all countries without distinguishing countries. Studying both groups under an identical analytic framework provides a more informative understanding of the distributional effects of capital inflows.

The remainder of the paper proceeds as follows: Section 2 provides a brief overview of the related existing literature. Section 3 then describes data and methodology. The empirical results are presented and discussed in Section 4, and Section 5 concludes.

2 Review of the literature

The theoretical basis for the distributional effects of capital inflows and its transmission mechanisms is still not well-established compared to the effects of trade integration. The existing literature has conflicting views - positive or negative - about the impacts of international financial integration on income inequality. It is, therefore, challenging to predict how capital inflows will affect income inequality within a country.

The mainstream view argues that international financial integration boosts economic growth. According to the widely accepted view of economists, capital inflows stimulate domestic investment, enable efficient use of resources, raise employment, and contribute to technological development and human capital accumulation. As a result of economic development, the incomes of the poor may grow faster than the average total income. Eventually, incomes are distributed more evenly.[3]

On the other hand, a surge in capital inflows increases the volatility of the macroeconomy and the likelihood of a financial crisis in recipient countries. The outbreak of a financial crisis can initially affect the rich through a fall in asset prices or corporate insolvencies. However, the economic recession caused by the financial crisis may disproportionately hit people with low incomes in the middle and long term and exacerbate income inequality (de Haan & Sturm, 2017; Furceri & Loungani, 2018). This uneven effect on the low-income class is because the low-income class typically has poorer working conditions than the high-income class, and they commonly have no buffer to absorb the income shock.

[3] The relationship between economic growth and income inequality may appear differently depending on the degree of economic development. Based on long-term historical data, Kuznets (1955) argue that a development pattern of income inequality has an inverted U-shaped relationship with economic development. This argument means that income inequality increases in the early stage of economic growth and decreases in the mature phase of economic development.

Another channel through which capital inflows affect income inequality relates to improvements in domestic financial conditions. Increased international capital movements reduce borrowing costs in recipient countries, improve access to financial resources for low-income households, and enhance the efficiency of domestic financial institutions. Under specific circumstances where financial market imperfections exist, such as information asymmetry and high transaction costs, easing financial market regulations and constraints will mitigate these imperfections and bring more benefits to the low-income class (Beck et al., 2007; Bumann & Lensink, 2016). By contrast, Jaumotte et al. (2013) argue that the benefits of reduced restrictions on access to international finance and improved domestic financial conditions will primarily accrue to capital owners and high-income households. Capital inflows make it easier and cheaper for wealthy economic agents to lend for production or investment, so they may earn higher future incomes than those with difficulty accessing financing. Low-income classes usually have less access to financial markets than richer classes, especially in developing countries.

This channel can have different effects depending on the domestic economy's and institutions' degree of development (Erauskin & Turnovsky, 2019). Improved access to international financial markets in highly institutionalized countries can reduce income inequality by spreading the benefits evenly across income groups. Conversely, in countries with weak institutional development, the benefits of capital inflows can only accrue to the rich by lowering the cost of foreign borrowing.

Another important channel is the mechanism through the complementarity between capital and skills (Larrain, 2015; Asteriou et al., 2014; Li & Su, 2021; Liu et al., 2023). This channel is based on the theoretical discussion of Feenstra and Hanson (1996, 1997) and is mainly associated with inward Foreign Direct Investment (FDI), although it can also be related to portfolio investment. The more FDI flows in a country and consequently accumulates in the economy, the greater the demand for high-skilled labor in sectors with high complementarity between capital and skills since FDI is generally directed to relatively high-skill sectors in the host country. These effects relate to the relative wage inequality between high-skilled and low-skilled workers.

On the other hand, some literature points out that the distributional effects of FDI depend on the economic conditions in recipient countries and the types of FDI. If capital ownership is concentrated in high-income groups, FDI will increase the return to labor relative to capital and contribute to a decline in income inequality (Eichengreen et al., 2021). In the case of greenfield investment in labor-abundant countries, FDI can accelerate wage growth in labor-intensive sectors by raising the demand for low-skilled workers (Cornia, 2011). By contrast, in the case of mergers and acquisitions (M&A) that involve transferring ownership, foreign firms usually impose cuts in employment and consolidations among firms, which leads to adverse distributional effects. (Baldwin, 1995).

Several recent empirical studies suggest that income inequality, primarily measured by the Gini coefficient, has increased with capital account liberalization. Larrain (2015) find empirical evidence that opening the capital account increases wage inequality between high-skilled and low-skilled workers using panel data from 20 developed countries for 1975-2005. Through a sectoral analysis, this study provides evidence that the distributional effect of opening capital accounts is pronounced in industries with high external financial dependency and strong complementarity between capital and skilled labor. Bumann and Lensink (2016) show that capital account liberalization distorts income distribution within countries with poor financial development, using extensive panel data sets with 106 countries from 1973 to 2008. The results also suggest that financial liberalization can lead to a more equal income distribution if financial depth (measured by domestic credit to the private sector over GDP) is high. On the contrary, the empirical results of de Haan and Sturm (2017) based on panel data from 121 countries for 1975-2005 indicate that the distorting effect of financial liberalization on inequality increases with a degree of financial development.

Furceri and Loungani (2018) show that episodes of capital account liberalization are associated with increased income inequality, based on panel data for 149 countries from 1970 to 2010. Their study also argues that capital account liberalization weakens workers' bargaining power over employers, reducing the wage share. Furceri et al. (2019) also find that capital account liberalization has negatively impacted income inequality, particularly in industries heavily dependent on external financing and exhibiting higher substitution elasticity between capital and labor. Li and Su (2021) argue that opening capital accounts in developing countries increases income inequality in the long run by reducing the income share of the poorest half and increasing the income share of the top 10%, a relationship not observed in developed countries. In addition, some studies investigate the distributional effects of financial integration as part of international integration (or globalization), which includes trade integration (Jaumotte et al., 2013; Asteriou et al., 2014; Lang & Tavares, 2018). They document that financial integration has been the driving force of income inequality.

The studies discussed above mainly focus on changes in income inequality before and after the episode of capital account liberalization, using *de-jure* indicators such as the Capital Account Openness Index (KAOPEN Index) developed by Chinn and Ito (2006, 2008), and the Financial Reform Index from Abiad et al. (2008). There are relatively few studies using *de-facto* indicators such as changes in external liabilities and assets. However, along with discontinuous political and institutional changes related to international capital flows, the continuous movements in capital flows can also explain the distributional effects of international financial integration. In this paper, I focused on the dynamic impact of cross-border capital inflows on income inequality.

The endogenous nature of capital inflows is one of the critical issues in examining the macroeconomic impacts of capital inflows. Capital inflows can be affected by the domestic conditions of a recipient country, such as output, institutional quality, political stability, and income inequality.[4] To circumvent this potential problem, some studies employ an Instrument Variable (IV) estimation. Liu et al. (2023) instrument capital flows using the movements in the U.S. two-year Treasury yields. Dorn et al. (2018) and Lang and Tavares (2018) also investigate the effects of international integration on income inequality using an IV approach, though they focus on overall globalization rather than financial integration. In this paper, I employ the panel VAR approach to treat variables as endogenous and identify an external shock in capital inflows using sign restrictions rather than an IV approach to address the endogeneity of capital inflows.

3 Empirical methodology

3.1 Data

I estimate a panel VAR model using a panel of 53 countries with annual observations from 1990 to 2020. The dataset is an unbalanced panel because of missing data in specific years for several countries. The set of 53 countries is divided between 26 AE and 27 EME based on the IMF classification of countries. Table A.1 in the Appendix presents a list of countries included in the sample.

It may not be reasonable to analyze the two groups jointly since AE and EME show significant differences in the level of economic development and the degree of international financial integration. Table 1 shows these differences between the two groups in 1990 and 2020. In 1990, the GDP per capita of AE was about seven times that of EME, on average. Even as of 2020, although there has been rapid growth of EME over the period, the GDP per capita of AE was about $40,000 on average, about five times that of EME.

[4] de Ferra et al. (2021) argue that countries with higher income inequality experience larger capital inflows than countries with lower income inequality since households in unequal counties borrow more. Dorn et al. (2018) assert that reverse causality may occur because changes in income inequality are likely to influence integration policies.

The degree of international financial integration can be inspected using the *de jure* indicators, such as the KAOPEN Index, and *de facto* indicators, such as total external liabilities. As the KAOPEN Index illustrates, AE have maintained high levels of financial openness throughout the period. On the contrary, EME were financially less open in 1990 and are still less than AE in 2020, despite the steady rise in capital account openness over the past decades. The size of external liabilities is broadly similar to the difference and trend of the KAOPEN Index. The external liabilities, expressed as a percentage of GDP, were higher in AE than in EME in 1990 and increased more greatly in AE than in EME during the period. These differences are the reasons why I separately study the distributional effects of capital inflows by dividing countries into two groups.[5]

Table 1: Economic development and financial integration in AE and EME

		1990		**2020**	
		AE	**EME**	**AE**	**EME**
GDP per capita	*Mean*	28.5	4.0	39.2	8.0
	SD	14.5	2.2	19.4	4.1
KAOPEN index	*Mean*	0.68	0.25	0.98	0.53
	SD	0.30	0.32	0.09	0.32
External liabilities	*Mean*	82.6	55.4	398.1	128.7
	SD	47.4	26.4	460.2	73.2

Notes: GDP per capita is represented in thousands of constant USD in 2015. The KAOPEN index in this table is a normalized index ranging between zero and one. Zero means "least financially open", and one means "most financially open". External liabilities are presented as a percentage of nominal GDP.
Source: The World Development Indicator (WDI) database of the World Bank, the Chinn and Ito (2006) database (`https://web.pdx.edu/~ito/Chinn-Ito_website.htm`), and the External Wealth of Nations Marks II (EWN II) database from Lane and Milesi-Ferretti (2018).

The panel VAR model includes six variables: (i) the year-over-year difference in the U.S. long-term interest rates (the U.S. 10-year Treasury yields) that are used as a proxy for world long-term interest rates, (ii) the growth rate of world GDP, (iii) capital inflows as a percentage of GDP, (iv) the growth rate of domestic GDP, (v) the year-over-year movements in domestic long-term interest rates (typically 10-year

[5] For robustness check in section 4.2, I exclude some countries from the sample in order to reduce the heterogeneity of the countries.

government bond yields), and (vi) the annual change rate of the Gini coefficient.[6] The last variable, which captures the change in income inequality within countries, is this paper's primary variable of interest. The nominal long-term interest rates were converted to real interest rates using inflation rates and inflation expectation in the U.S. case.[7]

The Gini coefficient is the measure of income inequality, which is most commonly used in the literature. A large body of literature studying the cause of income inequality uses the data of Gini coefficients from the World Income Inequality Database (WIID) or the Standardized World Income Inequality Database (SWIID). WIID is a database created by Deininger and Squire (1996) and maintained by the United Nations University-World Institute for Development Economics Research (UNU-WIDER). SWIID is a database compiled by Solt (2009) and Solt (2020), primarily based on WIID but supplemented by several other data sources, such as the Luxembourg Income Study (LIS). In this paper, I employ Gini coefficients from SWIID, which provides the most comprehensive estimates of Gini indices of disposable and market income inequality for 198 countries from 1960 to 2021.[8]

It is important to note that the Gini coefficient is incomplete in detailing movements in income inequality. Other measures, such as income share by income class, may deliver more pertinent information for income distribution. In order to address the limitation of using the Gini coefficient, I employ the data on income shares by income class from the World Inequality Database (WID), which was developed by Piketty and Zucman (2014) and then extended to include the evolution of the national income structure in the long run. WID combines national accounts, fiscal, and survey data to compute longer and more reliable income share series and provides high-quality information for the pre-tax income share of the top 10%, middle 40%, and bottom 50% groups.

Capital inflows refer to the *gross* capital inflows, which means a net increase in domestic financial assets of non-residents, i.e., the foreign investors' purchase of domestic financial assets less the foreign investors' disposal of domestic financial assets. In the present paper, capital inflows are represented as annual changes in

[6] To ensure the stability of the model, I use the first differences in variables for estimation. The level of long-term real interest rates and the Gini coefficient are used for robustness checks in the section 4.2.

[7] It is necessary to compute the *ex-ante* real long-term interest rates by using expected inflation, but data and estimates on expected inflation are not available in many countries. For the U.S. long-term real interest rates, this paper uses the Inflation expectation estimated by the Federal Reserve Bank of Cleveland.

[8] I use Version 9.4 of the database, published in November 2022, for estimation (`https://fsolt.org/swiid/`). The SWIID database provides a disposal Gini based on *post*-tax and *post*-transfer income and a *Market* Gini based on *pre*-tax and *pre*-transfer income. I employ the latter to rule out the impact of the government's redistribution policies.

the total external liabilities as a percentage of nominal GDP. Data on updated external liabilities, which are non-resident holdings of domestic financial assets, is from the EWN II dataset from Lane and Milesi-Ferretti (2007) and its update (Lane & Milesi-Ferretti, 2018).[9] The database contains estimates of international assets and liabilities for 212 countries from 1970 to 2021. This estimate is primarily based on each country's International Investment Position (IIP) statistic and Balance of Payments (BOP) statistic, published and managed by the IMF.

Data on GDP (world and domestic) and inflation rate (GDP deflator) are taken from the WDI database of the World Bank.[10] Data on long-term nominal interest rates were obtained from the CEIC database. Descriptive statistics of all variables are shown in Table 2, which reports statistics separately for AE and EME. [11]

Table 2: Descriptive Statistics

Variables	N	M	SD	Min	Max
AE					
Capital inflows	815	0.18	0.42	-0.73	3.78
Δ lnGDP	804	2.43	3.43	-16.06	21.81
Δ lnGini	807	0.34	0.96	-3.74	7.01
Δ Long-term rates	683	-0.28	2.81	-18.99	27.24
EME					
Capital inflows	827	0.07	0.12	-0.44	1.30
Δ lnGDP	846	3.24	5.70	-59.60	13.31
Δ lnGini	825	0.11	1.03	-4.46	8.72
Δ Long-term rates	484	-0.08	5.35	-32.26	68.19
All countries					
Capital inflows	1,642	0.12	0.31	-0.73	3.78
Δ lnGDP	1,650	2.84	4.75	-59.60	21.81
Δ lnGini	1,632	0.23	1.00	-4.46	8.72
Δ Long-term rates	1,167	-0.20	4.06	-32.26	68.19

Note: N, M, and SD represent the number of observations, mean, and standard deviation, respectively.

[9] I use the December 2022 version (`https://www.brookings.edu/articles/the-external-wealth-of-nations-database/`). The EWN II dataset is based on the classification of residents and non-residents, not on financial asset holders' nationality.

[10] Data for Taiwan is obtained from the IMF World Economic Outlook (WEO) database.

[11] As shown in Table 2, variables include large variability with some outliers. To address the potential bias due to extremely large or small observations, I perform a robustness check using the winsorized data in Section 4.2.

3.2 Panel VAR model

To study the dynamic distributional effect of capital inflows, I estimate the following reduced-form panel VAR model of order p[12]:

$$Y_{i,t} = A_1 Y_{i,t-1} + ... + A_p Y_{i,t-p} + \mu_i + u_{i,t} \tag{1}$$

with the index $i = 1, 2, ..., N$ and $t = 1, 2, ..., T_i$, which represent countries and time periods, respectively. $Y_{i,t}$ is an $(m \times 1)$ vector of endogenous variables for country i at time t. $A_1, ..., A_p$ are $(m \times m)$ matrices of estimated coefficients for lagged dependent variables. μ_i is a vector of country-specific fixed effects, which capture time-invariant unobserved heterogeneity between different cross-sectional units. $u_{i,t}$ is a normally-distributed error term with zero mean, i.i.d. assumption and covariance matrix Σ_u.

In a dynamic panel model, such as the panel VAR model in equation (1), classical ordinary least square (OLS) estimation with individual fixed effects no longer provides unbiased estimators (Nickell, 1981). The bias arises because the fixed effects are correlated with the regressors, which are the lag values of the dependent variables. To avoid this problem, I apply the forward orthogonal transformation, commonly referred to as the "Helmert transformation" (Arellano & Bover, 1995), to equation (1) and employs a generalized method of moments (GMM) approach proposed by Arellano and Bover (1995). The forward orthogonal transformation removes the forward mean, i.e., the mean of all the future observations for each country and year.[13] This procedure preserves the orthogonality between transformed variables and lagged dependent variables. It allows us to use lagged regressors as instruments and to estimate the coefficients in the VAR model by GMM estimation. In addition, this transformation procedure minimizes data loss in unbalanced panel datasets.[14]

The key challenge in applying VAR for studying the responses of variables to the structural shocks is identification, i.e., how to decompose the error $u_{i,t}$ in the reduced form VAR into structural disturbances. The general structural form of panel VAR can be represented in the following form:

[12] The Stata package "pvar" provided by Love and Zicchino (2006) and Abrigo and Love (2016) was used for estimation. I added more programming to the code to impose sign restrictions on the estimated impulse response functions.

[13] The forward orthogonal transformation is given by $Y^{\perp}_{i,t+1} = \alpha_{i,t}(Y_{i,t} - 1/T_{i,t} \sum_{s>t} Y_{i,t})$, where $\alpha_{i,t} = \sqrt{T_{i,t}/(T_{i,t}+1)}$.

[14] For balanced panels, the first difference and forward orthogonal transformation give the same estimated results when the instruments are fixed (Arellano & Bover, 1995).

$$B_0 Y_{i,t} = B_1 Y_{i,t-1} + \ldots + B_p Y_{i,t-p} + \nu_i + \epsilon_{i,t} \tag{2}$$

where B_0 is an $(m \times m)$ matrix representing the contemporaneous reactions of the variables to the structural shocks, $B_1, \ldots, B_p$ are $(m \times m)$ matrices of structural coefficients, and $\epsilon_{i,t}$ is an $(m \times 1)$ vector of the structural shock terms that have zero means, no serial correlation and, no correlation between the individual shocks, i.e., $E(\epsilon_{i,t}\epsilon_{j,t}) = 0$ and $E(\epsilon_{i,t}\epsilon'_{i,t}) = I$. The error term vector $u_{i,t}$ in equation (1) of the reduced form panel VAR is linearly connected with the vector of orthonormal structural shocks $\epsilon_{i,t}$ in equation (2) as follows:

$$u_{i,t} = \Theta\epsilon_{i,t} \tag{3}$$

where $\Theta = {B_0}^{-1}$. The mathematical expression of the problem of identifying simultaneous relationships between variables is to identify the matrix Θ, which describes the effects of the shocks. The ith column of Θ can be defined as the immediate impact on all other variables of the ith structural shock, typically one standard error in size. From the equation (3), the relationship between the covariance matrix Σ_u and the matrix Θ can be written as:

$$\Sigma_u = E(u_{i,t}u'_{i,t}) = E(\Theta\epsilon_{i,t}\epsilon'_{i,t}\Theta') = \Theta\Theta' \tag{4}$$

As the above equation (4) shows, the matrix Θ can be directly derived from the covariance matrix Σ_u. Since there are many degrees of freedom in specifying Θ, imposing some restrictions on Θ is necessary to achieve identification.

The most popular strategy for identification is to impose zero restrictions on the matrix Θ. This approach transforms the reduced-form panel VAR into a "recursive" form since the matrix Θ obtained by the Cholesky decomposition of the covariance matrix Σ_u is a lower triangular matrix. This identification scheme orders the variables according to their perceived degree of "exogeneity" and, therefore, the estimated results crucially depend on ordering the variables. If economic theories do not strongly support the ordering of variables, any order for identification of the economic shock using the recursive strategy might be arbitrary.

With zero restrictions, I use sign restrictions proposed Faust (1998), Canova and Nicolo (2002), and Uhlig (2005) as the identification strategy. Sign restrictions impose a positive or negative sign as a restriction on the variables' responses to the structural shocks. While zero restrictions define that some variables are unaffected by a shock from other variables, sign restrictions combine expected information on how economic variables respond to a structural shock in the system. Using both restrictions simultaneously enables one to capture the shocks better and isolate those from other shocks.

In this identification strategy, Θ is defined as the product of the matrix P (obtained by the Cholesky decomposition of Σ) and the orthogonal matrix Q. Since $QQ' = I$, Θ can be represented as follows:

$$\Sigma_u = \Theta\Theta' = PQ(PQ)' = PQQ'P' = PIP' = PP' \tag{5}$$

It must be checked that Q satisfies a set of sign and zero restrictions since the orthogonal matrix Q is not unique. In order to obtain the plausible matrix Q, this paper uses the algorithm proposed by Fry and Pagan (2011) and Arias et al. (2018). For each independent draw i of $\tilde{\Sigma}_u$ and $\tilde{A}_1, \dots, \tilde{A}_p$ using Normal-Wishart prior in the reduced form parameters, the orthogonal matrix Q_i is also drawn such that the structural parameters satisfy zero restriction. The matrix Q_i is randomly generated using Householder transformations based on QR-decomposition. The algorithm for QR-decomposition is to generate some $(m \times m)$ random matrix W from an $\mathcal{N}(0, I_m)$ distribution and then decompose $W = QR$, where Q is an orthogonal matrix from a uniform distribution and R is a triangular matrix whose diagonal is normalized to positive. Then, the impulse response functions are calculated using $\tilde{\Sigma}_u$, $\tilde{A}_1, \dots, \tilde{A}_p$, and Q_i. The draw i is accepted if it assembles the impulse response functions that satisfy the imposed sign restrictions and is discarded if not. This procedure is repeated until there are N accepted impulse response functions. The posterior median and error bands, the 16th and 84th percentiles that account for parameter uncertainty, of the obtained N impulse responses are subsequently computed and reported.[15]

3.3 Identification

In this paper, the exogenous capital inflow shock is interpreted as an unexpected increase in foreign investors' demand for domestic financial assets unrelated to the domestic economic situation. These external shocks to capital inflows are driven by the change in global financial conditions ("push" factors), which are heavily influenced by US monetary policy and related market volatility, and also the degree of risk aversion of global investors. An unexpected increase in global saving may also lead to surges in capital inflows. In order to adequately identify capital inflow shocks, it is required to identify shocks orthogonal to domestic economic shocks ("pull" factors), which can also lead to an increase in capital inflows.

[15] In this paper, the number of accepted draws, N, is set to 500.

Table 3 lists the restrictions imposed in this paper to identify external shocks to capital inflows. To distinguish these shocks from domestic shocks, I imposed sign and zero restrictions on both types of shocks simultaneously.[16] As discussed in Sa et al. (2014), imposing restrictions on long-term interest rates can be useful for distinguishing capital inflow shocks driven by "push" factors from other shocks. Other things being equal, capital inflows driven by "push" factors will be associated with downward pressure on domestic interest rates, while inflows due to "pull" factors will tend to put upward pressure on domestic interest rates. In an open-economy model, a global increase in demand for domestic financial assets generally lowers domestic real interest rates since domestic residents consume more while domestic investment rises. Another possibility is that foreign monetary authorities, such as the U.S. or EU central bank, make an expansionary monetary policy. Low foreign interest rates make domestic assets more attractive, and corresponding capital inflows drive down the domestic interest rate. Following Sa et al. (2014), it is assumed that an increase in capital inflows driven by external shocks leads to a fall in domestic long-term real interest rates[17], while domestic shocks lead to a rise in domestic long-term real interest rates. However, this restriction on domestic shock may be too restrictive, as it assumes that domestic long-term interest rates rise in response to domestic supply shocks. No restriction, therefore, is imposed on the response of domestic long-term interest rates to domestic shocks.[18] Additionally, in contrast to the external shocks driven by "push" factors, the domestic shocks driven by "pull" factors are supposed to have no impact on world interest rates (the U.S. long-term interest rates) and world GDP.

No restrictions are imposed on the response of the Gini coefficient, the main variable of interest. Uhlig (2005) called this approach the "agnostic approach", which does not impose any restrictions on the impulse response function of the interesting variables. This approach relies on data rather than restrictions to investigate dynamic response pathways of the interesting variables. In addition, to ensure the validity of this restriction, some large economies were not included in the sample for the estimation of the model.[19] As the financial assets of these countries are recognized as global safe-haven assets, the movements and economic effects of capi-

[16] Uhlig (2005) argues that including other shocks in the model using sign restriction ensures that the shocks of interest truly capture their exogenous component and not an endogenous response to other innovations in the system.

[17] As the central bank largely controls the short-term interest rates, the restrictions are imposed on the long-term interest rates rather than the short-term interest rates. One can assume that the long-term interest rates are determined in the market.

[18] Even if a positive sign restriction is imposed on the response of domestic long-term real interest rates to domestic shocks, the estimated results are qualitatively and quantitatively not different from the results reported in Section 4.

[19] Germany, Japan, and the United States

tal inflows in these countries are significantly different from those of other countries, even in the identical phase of the global financial cycle. The restrictions are imposed only upon impact. A restriction over two years is too restrictive.

Table 3: Sign restrictions

	Global shock (Push factors)	Domestic shock (Pull factors)
U.S. long-term rates	−	**0**
World GDP	+	**0**
Capital inflows	+	+
Domestic GDP	+	+
Domestic long-term rates	−	unrestricted
Gini coefficient	unrestricted	unrestricted

4 Results

This section presents the estimated impulse response functions from the panel VAR model and the results of robustness checks, and briefly discusses the empirical findings. I estimate the model with two lags for considering sluggish movements of income inequality. As a robustness check, I re-estimate the model with different lag lengths in the section 4.2. The results reported in this section focus only on the impact of external capital inflow shocks ("push" factors) on income inequality (and other domestic variables). The first reason for doing this is that this study mainly aims to investigate the distributional effect of capital inflows driven by changes in global financial conditions unrelated to domestic conditions. Furthermore, since domestic shocks summarize various disturbances such as monetary, demand, and supply shocks, domestic shocks may not be accurately identified simply by imposing the sign restrictions in Table 3.

4.1 Baseline results

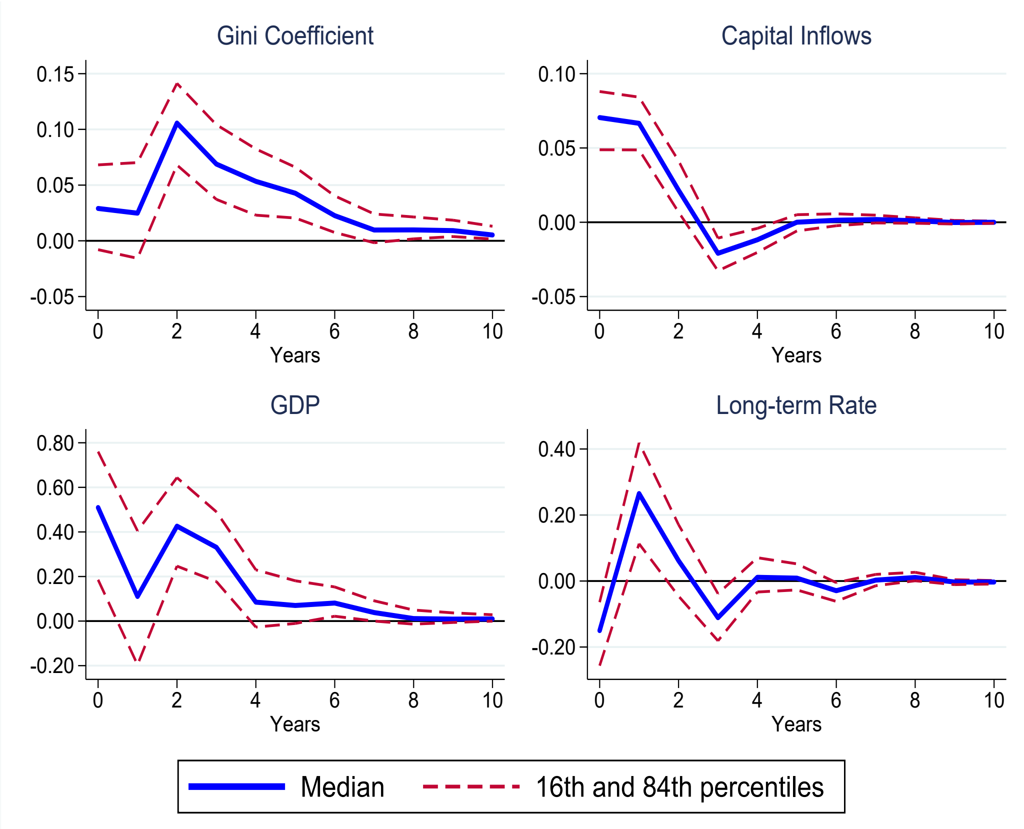
Figures 1 and 2 show the impulse response functions for a one-standard-deviation external shock in capital inflows over ten years for AE and EME, respectively.[20] The solid blue line represents the median of impulse response functions across all independent draws that satisfy the sign and zero restrictions in Table 3. The dashed

[20] Figure A.1 in the Appendix presents the impulse response functions for all countries.

red lines are the confidence bands formed using the 16th and 84th percentile of all accepted draws.

The upper left panel of Figure 1 presents the impulse response of the Gini coefficient for AE to an external capital inflow shock that leads to an unexpected increase in capital inflows by about 7% of GDP. At the median, a capital inflow shock leads to an increase in the Gini coefficient, i.e., income inequality. The response of the Gini coefficient's year-over-year change rate peaks at about 0.1 percentage point two years after the occurrence of the shock. This result is statistically significantly different from zero and also economically significant because income inequality changes very gradually over time. The Gini coefficient also reacts positively to a shock in capital inflows upon impact and after one year, but these responses are not statistically significant. The estimated impulse response of the Gini coefficient suggests that an external shock in capital inflows has a delayed effect on income inequality. Moreover, The effects of capital inflow shocks last for a long time after two years, although the magnitude of the responses slowly decreases.

Figure 1: Impulse responses to an external capital inflow shock in AE



As shown in the lower left panel, a capital inflow shock also leads to an output expansion in recipient countries. The growth rate of GDP rises immediately by about 0.5 percentage points as a response to the shock. The response of GDP is also long-lasting and statistically significant. The movements in the long-term real interest rates show a negative response at the time of the capital inflow shock.

Figure 2 presents the results for EME. As shown in the upper left panel, the Gini coefficient falls one year after the shock that leads to an unexpected increase in capital inflows by about 2% of GDP.[21] At the median, a capital inflow shock leads to a decline in the annual change rate of the Gini coefficient by about 0.06 percentage points. This result is statistically significantly different from zero and also economically significant. Negative responses of income inequality to the shock last for a long time but are not statistically significant (confidence bands include zero). Contrary to income inequality, the impulse response functions of GDP and long-term interest rates to a capital inflow shock in EME are considerably similar to the previous AE results.

Figure 2: Impulse responses to an external capital inflow shock in EME

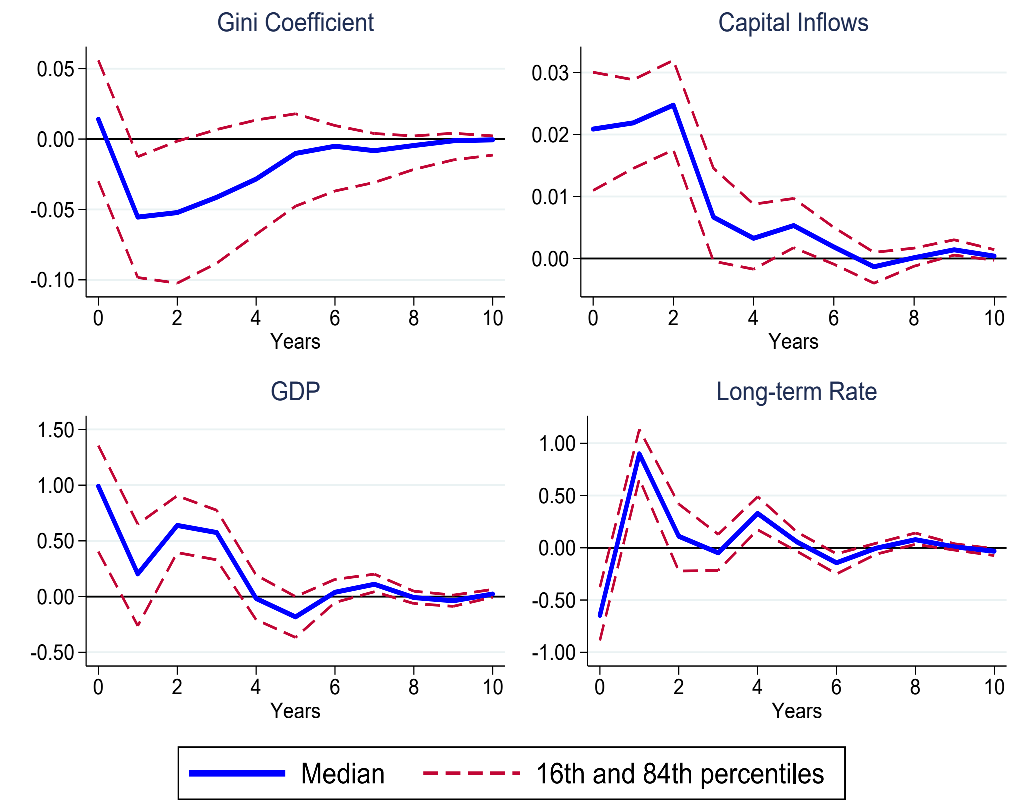


Table 4 shows the forecast error variance decomposition. At the median, capital inflow shocks explain about 14.6% of the variance in the Gini coefficient in AE and about 10.2% of the variance in the Gini coefficient in EME at a 10-year forecast horizon, respectively.

[21] It is less than the magnitude of the capital inflow induced by the identical shock in the case of AE. This difference is consistent with data showing that capital inflows typically arise on a larger scale in AE than in EME.

Table 4: Forecast error variance decomposition

	Capital inflow shock			
	1Y	3Y	5Y	10Y
AE	1.3	9.7	13.5	14.6
EME	0.4	7.5	10.0	10.2

Note: The table shows the percentage of the forecast error variance of the Gini coefficient explained by the capital inflow shocks over 1-, 3-, 5- and 10-year time horizons.

However, the above-presented results cannot explain the distributional effects of capital inflows by income class because the Gini coefficient consists of aggregated data across all income classes. Even if the Gini coefficient increased, it can not be accurately captured whether this was due to an increase in relative income for the upper-income class, a decrease in the relative income for the lower-income class, or both. For this reason and to gain further insights into the impact of capital inflow shocks on income inequality, I re-estimate the panel VAR model by replacing the dependent variable with the income shares of the top 10% and bottom 50% group from the WID, respectively.

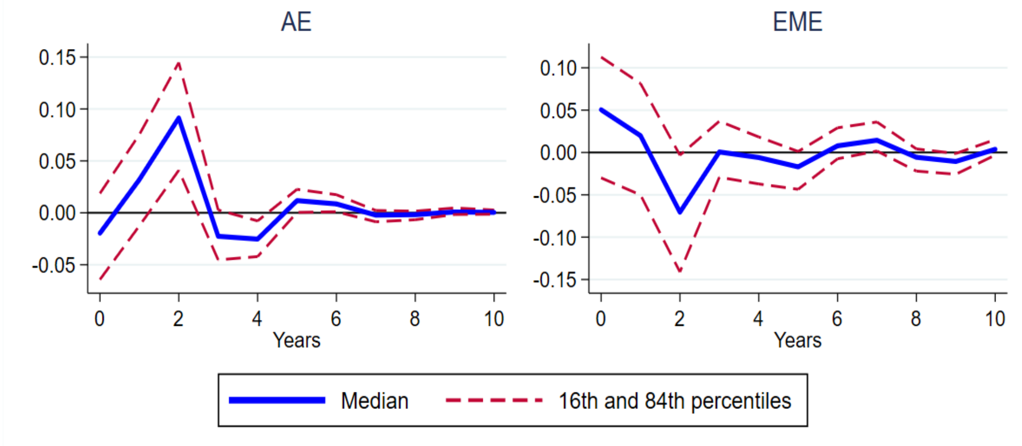
The results using income share data suggest that surges in capital inflows due to external shocks are associated with an increase in the income share of the rich group and a decrease in the income share of the poorest half in AE. By contrast, a capital inflow shock is associated with a decrease in the income share of the rich group and an increase in the income share of the poor group in EME. In AE, an increase in the income share of the rich group dominates the overall distributional effects of a capital inflow shock. In EME, the magnitude of the positive response of the income share of the poor group is more significant than the negative response of the rich group.[22]

Figure 3 shows the impulse response of the income share of the top 10% group. A capital inflow shock leads to an increase in the income share of the rich group in AE after two years, and this response is statistically significant. After that, the response of the top 10% group reverts to zero, repeatedly going up and down. Conversely, the income share of the top 10% group in EME reacts negatively to a shock in capital inflows after two years.

As shown in Figure 4, the income share of the bottom 50% group responds to an external shock in capital inflows in the opposite direction than above. While a

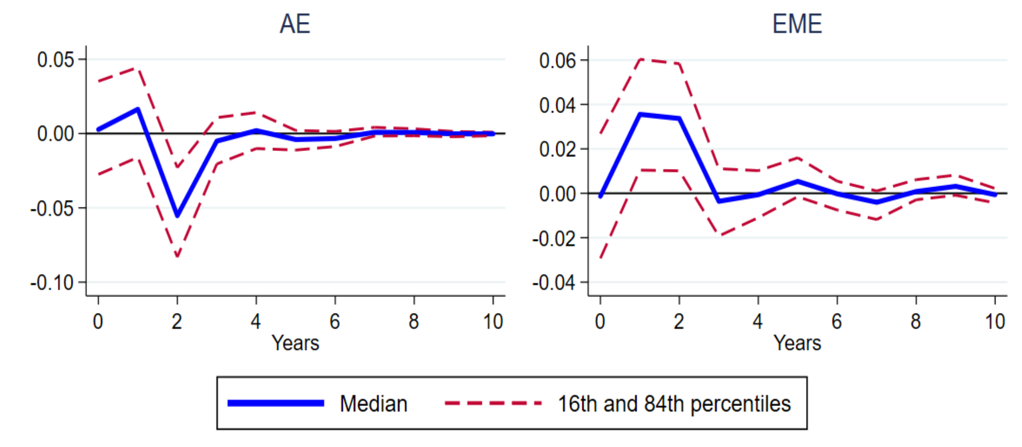
[22] Figures A.2, A.3, A.4 and A.5 in the Appendix show the impulse response functions for all domestic variables.

Figure 3: Impulse response of the income share of top 10% group to capital inflow shock



capital inflow shock leads to a decline in the income share of the poorest half after two years in AE, a shock increases those of the poorest half after one and two years in EME. These responses of the income share of the bottom 50% group in AE and EME are statistically significant.

Figure 4: Impulse response of the income share of bottom 50% group to capital inflow shock



4.2 Robustness checks

In this section, I conduct various robustness checks for baseline results in Figures 1 and 2 and report the impulse response of income inequality to capital inflow shocks in AE and EME separately. Figures in the Appendix report the impulse response functions of all variables.

4.2.1 Alternative measure of income inequality

Firstly, I test if the main findings are robust to using an alternative measure of income inequality that is also used in numerous relevant empirical literature. The panel VAR model is re-estimated using the other Gini coefficients from the Estimated Household Income Inequality (EHII) dataset established by the University of Texas Inequality Project (UTIP). The EHII dataset uses the econometric method to establish the relationship between the Deininger-Squire Gini coefficient and the Theil Index[23] based wage dispersion in industry.[24] The database includes 4,550 annual observations from 154 countries from 1963 to 2015.

Figure 5: Robustness check (1) - Gini from EHII

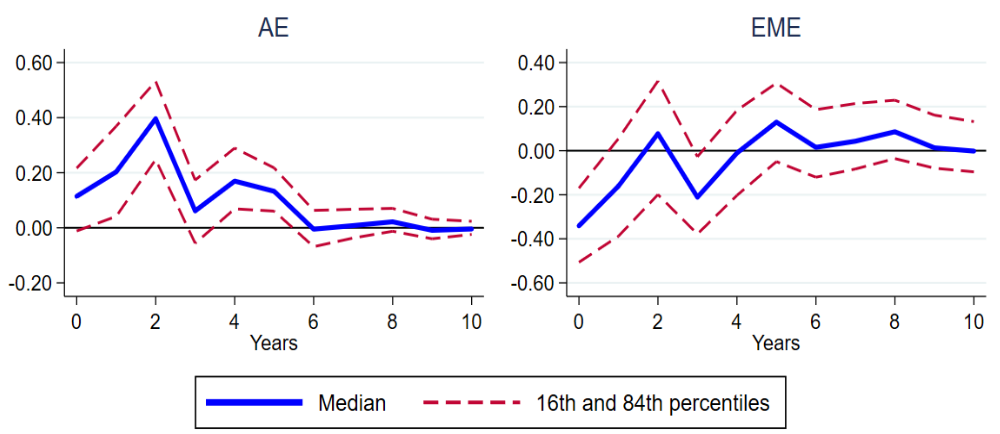


Figure 5 shows the impulse responses of income inequality re-estimated using the Gini coefficient from EHII. As shown in the left panel, the results for AE using other data on income inequality are broadly similar to those in the baseline results estimated with the Gini from SWIID. Over time, the effect of the capital inflow shock gradually decreases significantly. In the case of EME, while the shape of Gini's response is different from those in the baseline results, a capital inflow shock

[23] The Theil Index is calculated based on wage data from the industry statistics of the United Nations International Development Organization (UNIDO).

[24] Galbraith and Kum (2005) describe the methodology for constructing Gini coefficients from EHII.

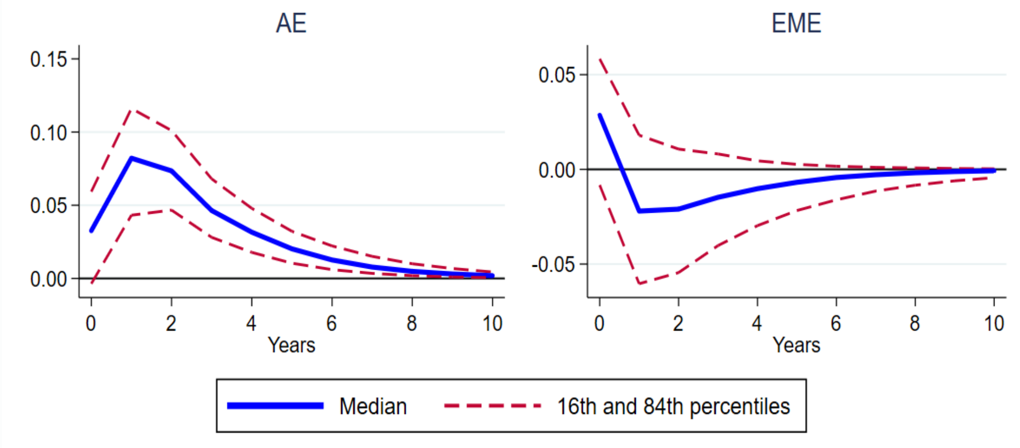
also leads to a statistically and economically significant decline in the change rate of the Gini coefficient.

4.2.2 Different lag lengths

The estimated impulse response functions can be sensitive to the selected lag length. To test the estimated model's robustness in various lag lengths, I re-estimate the model with lag 1 and 3.

Figure 6 shows the impulse response of the Gini coefficient to a capital inflow shock with lag one. The result in AE is broadly similar to the result in the baseline estimation. A capital inflow shock leads to an increase in the annual change rate of the Gini coefficient over time, and these responses of income inequality are statistically significant. In EME, a capital inflow shock reduces income inequality over periods after one year but relinquishes its statistical significance.

Figure 6: Robustness check (2.1) - lag one

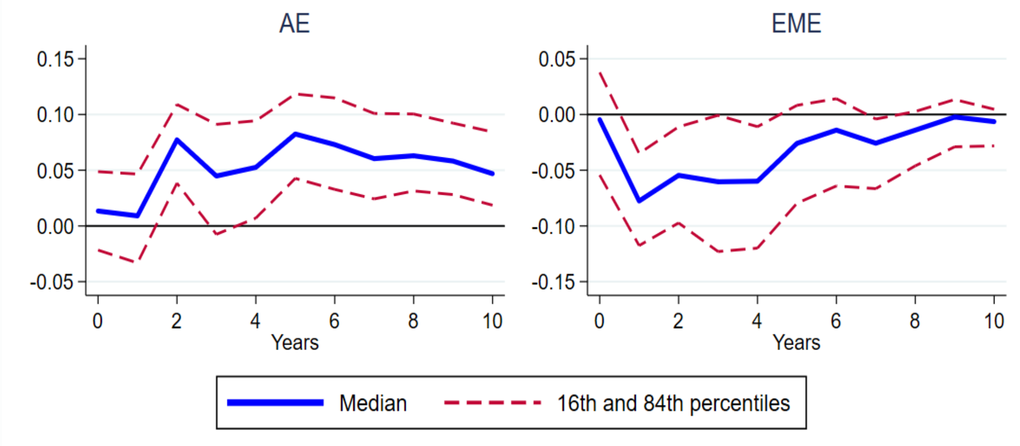


As shown in Figure 7, the baseline results for AE and EME are robust with lag three estimates. Similarly, a capital inflow shock leads to an increase in the year-over-year change rate of the Gini coefficient in AE and a reduction in that in EME. The impact of an external shock in capital inflows on income inequality is more persistent and statistically significant than in the case of two lags in both countries groups.

4.2.3 Using level of variables

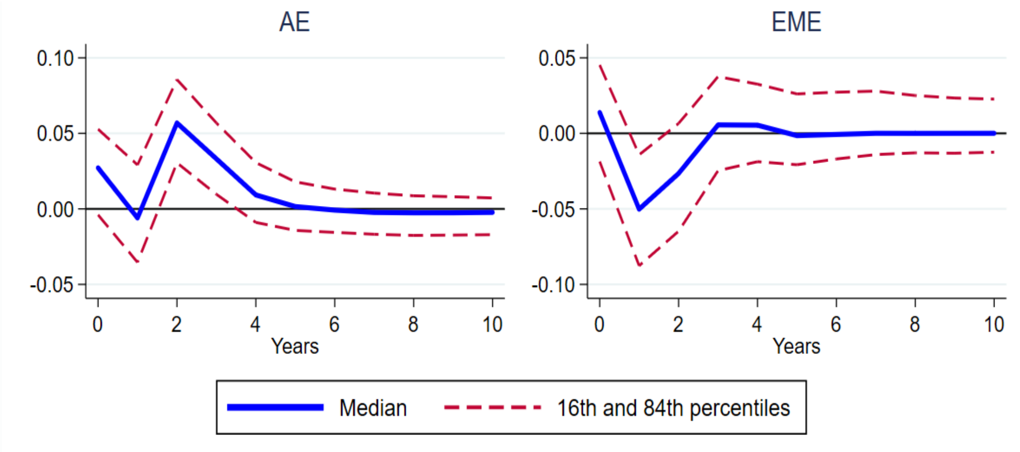
I also test if the impact of capital inflow shocks on income inequality in the baseline results is robust using the level of some variables rather than the first difference in the variables. In particular, I re-estimate the model with (i) the level of long-term interest rates and (ii) the level of the Gini coefficient. Figure 8 shows the

Figure 7: Robustness check (2.2) - three lags



re-estimated impulse responses of the Gini coefficient using the level of US and domestic long-term interest rates. While the magnitude of the responses becomes relatively smaller in AE and EME, the direction and statistical significance of the responses remain unchanged.

Figure 8: Robustness check (3.1) - using long-term interest rates in levels

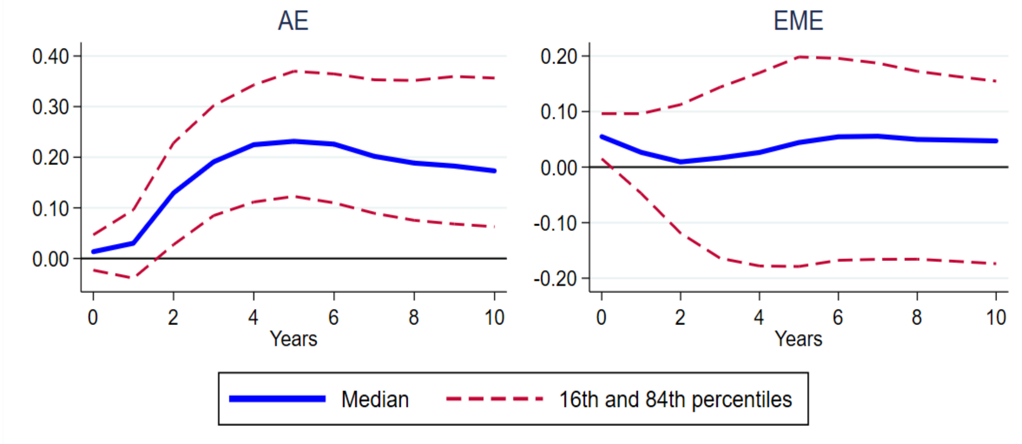


The results obtained using the level of the Gini again indicate a statistically significant and persistent effect of a capital inflow shock on the Gini in AE (the left panel in Figure 9). In contrast, as shown in the right panel, the responses of the Gini in EME do not deliver significant outcomes, unlike the baseline results.

4.2.4 Excluding some countries from the sample

I exclude some countries from the sample in order to address the possible bias due to the heterogeneity of countries in the baseline results. By re-grouping, the countries in both groups of AE and EME may become much more homogeneous regarding the level of economic development and the openness of the capital market than the case

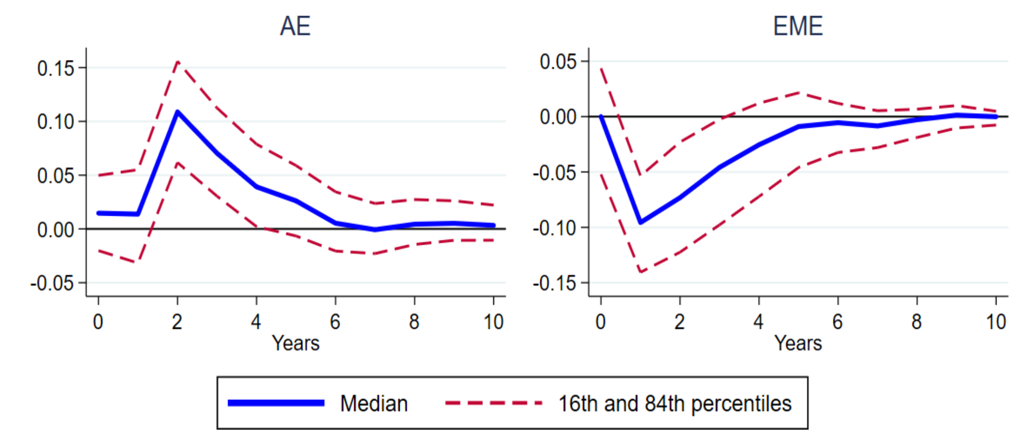
Figure 9: Robustness check (3.2) - using the Gini in level



in baseline estimation. Some countries, such as Korea, were not AE according to the 1990 IMF country classification, but they achieved rapid economic development during the period and were classified as AE. Other countries, such as the Czech Republic, in AE are transition economies that were gradually integrated into Western Europe during the period after the mid-1990s. These transition economies were also not classified as AE in 1990. On the other hand, some countries in EME are still too low in economic development and capital account openness. The panel VAR model is re-estimated without these countries.[25]

Figure 10 reports the results of this exercise. While the magnitude of the response of the Gini coefficient in AE to a shock in capital inflows is not significantly different from those for the baseline estimation, that of the Gini coefficient in EME is more significant than the baseline results.

Figure 10: Robustness check (4) - excluding some countries from the sample

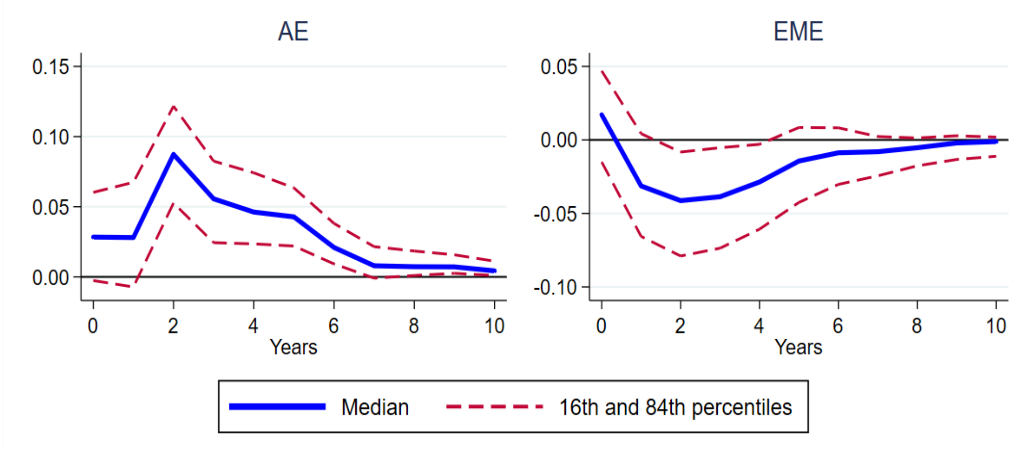


[25] AE: Czech Republic, Israel, Korea, Latvia, Lithuania, Slovak Republic, Slovenia, and Taiwan. EME: India, Indonesia, Moldova, Morocco, Philippines, and Tunisia.

4.2.5 Addressing the impacts of outliers

As shown in Table 2, all domestic variables in the model indicate some variability with outliers. To address this issue, I also test for the impact of outliers on the estimated results by winsorizing the four domestic variables — capital inflows, domestic GDP, domestic long-term real interest rates, and the Gini coefficient — at the 2.5% and 97.5% levels. The re-estimated results reported in Figure 11 show that the impulse response functions tend to be close to those of baseline results, and the differences in the impulse response functions are insignificant.

Figure 11: Robustness check (5) - addressing the impacts of outliers



4.3 Discussion

In this section, I briefly discuss how the empirical findings of this paper can contribute and add some insight into previous debates about the distributional effects of capital inflows.

The existing literature argues that capital account liberalization or a surge in capital inflows increases income inequality. Bumann and Lensink (2016), de Haan and Sturm (2017) and Furceri and Loungani (2018) document that capital account liberalization is associated with an increase in the Gini coefficient. Liu et al. (2023) find that surges in capital inflows raise income inequality using an IV approach. However, these studies report estimation results for all countries (or only AE (Asteriou et al., 2014; Larrain, 2015)) without distinguishing between AE and EME. On the other hand, the results of this paper suggest that the relationship between capital inflows and income inequality may differ in the two groups.

The main results using the Gini coefficient suggest that an external shock in capital inflows leads to an increase in income inequality in AE and a decrease in income inequality in EME. The adverse impact of capital inflow on income inequality in AE is consistent with the results of previous empirical studies using the *de jure* indicators.[26] However, the results in EME are inconsistent with empirical evidence from existing studies and are more in line with the mainstream theoretical views, which predict that capital inflows may reduce income inequality.

By employing another measure of income inequality, the income share of the rich group and the poorest half, I find that capital inflow shocks lead to an increase in the income share of the top 10% group in AE and the bottom 50% group in EME. These findings do not support the empirical results of Li and Su (2021) that capital account liberalization is associated with a decline in the income share of the poorest 50% and an increase in that of the richest 10% in developing countries.

The empirical findings of this paper suggest that capital inflows have different impacts on income inequality across countries. Then, why does the response direction of income inequality to capital inflow shocks differ between AE and EME? There are some candidates for the explanation of the reasons. As shown in Table 1 above, two country groups show significant differences in the level of economic development and the degree of financial integration. Many existing relevant studies point out that the level of economic, political, and institutional development of a recipient country may influence the relationship between capital inflows and income inequality.[27] Another candidate for explanation is the degree of international financial integration. While AE have tended to go deeper into international financial integration (i.e., greater openness of the capital account and larger size of external liabilities), EME have integrated into the international financial market to a relatively low degree. This fact implies that capital inflows can improve income inequality in the early stages of financial integration but worsen income inequality as financial integration progresses.

Additionally, the composition of total external liabilities may matter for the relationship between capital inflows and income inequality.[28] External liabilities are divided into the following categories: (i) FDI, (ii) portfolio investment, including equity and portfolio debt, and (iii) Other investment, including bank loans, non-resident deposits, trade credits. The composition of external liabilities shows notable differences between AE and EME.[29] While the proportion of other investments has steadily decreased in both groups since 1990, the proportion of portfolio

[26] The estimated impulse response functions for all countries presented in Figure A.1 in the Appendix are also consistent with other empirical studies.

[27] Some studies (Bumann & Lensink, 2016; de Haan & Sturm, 2017) also argue that financial depth, measured by the ratio of credit to the private sector over GDP, moderates the relationship.

[28] Harms et al. (2023) point out the high correlation between income inequality and the equity share in external liabilities and argue that entry barriers in non-traded goods industries may drive both variables. However, they do not assert a causal relationship between the variables.

[29] Figure A.20 in the Appendix shows the evolutions over time in the composition of external liabilities as a percentage of GDP in AE and EME.

investment has grown mainly in AE (27.3% → 39.2%), and the proportion of FDI has increased significantly in EME (15.0% → 50.3%). However, the inflows of these three types show high correlations with each other, and the estimation results using subcategories are quantitatively and qualitatively very close to the main results using total capital inflows. Therefore, the distributional effects of the subtypes of capital inflows can not be separately estimated in the framework of this paper.

It is beyond the scope of the present paper to complete these remarkable differences in the impacts of capital inflow shocks on income inequality between AE and EME, and is left for future research.

5 Conclusion

The increase in income inequality within countries over the past decades has sparked intense debates among economists about the cause and dynamics of this phenomenon. International financial integration, represented as an increase in cross-border capital flows, has received relatively less attention than other factors, such as trade integration and Skill-biased technology change (SBTC). The empirical findings of this paper suggest that movements in international capital flows driven by global financial conditions should also be treated as one of the important axes of this discussion.

In this paper, I examine the effects of capital inflows on income inequality within countries using a panel VAR model. To identify external shocks in capital inflows, I employ sign restrictions. By performing an analysis with this approach, I address the potential endogeneity of capital inflows. Furthermore, I separately perform the model estimation in AE and EME to obtain more accurate insights.

The empirical results suggest that capital inflow shocks worsen income inequality in AE while contributing to reducing income inequality in EME. An external shock in capital inflows increases the year-over-year change rate of the Gini coefficient in AE by about 0.1 percentage point two years after the shock. By contrast, the annual change rate of the Gini coefficient in EME falls by about 0.06 percentage points one year after the shock. Capital inflow shocks explain about 14.6% of the forecast error variance in the Gini coefficient in AE and about 10.2% in EME at a 10-year forecast horizon, respectively. These results are statistically and economically significant and robust to several robustness checks. Disaggregating these effects by income class, I also find that the shocks to capital inflows are primarily associated with an increase in the income share of the rich in AE and the low-income class in EME. The empirical findings of this paper imply that even if capital inflows foster the economic growth of recipient countries, the benefits from the output expansion may be unevenly distributed across income classes.

The possibility of adverse distributional effects of capital inflows delivers another rationale for policies of proper management and control of cross-border capital movements. In addition to macroeconomic instability and the possibility of a financial crisis, policymakers should consider the potential adverse impacts of capital inflows on income inequality. To this end, capital flow management policies at the individual national level and cooperation between countries, including international organizations, are required. Furthermore, policymakers should pay much attention to these policy implications because it may lead to an increase in domestic political unrest. If skepticism about the welfare effects of international financial integration is widespread and income inequality increases, populism can also spread (Rodrik, 2018). These changes in the political environment may reverse the existing integration trend and negatively impact the economic benefits generated by the integration. In this context, the recent IMF policy paper (IMF, 2022) proposes to include the distributional effects of international financial integration as one of the core issues within the Institutional View on Capital Flow Management Measures (CFMs) of the IMF (IMF, 2012). It should also be noted that IMF (2022) states that the topics "need further research and could not be addressed in this review. ... because the analytical foundation to propose policy changes is insufficient at this time".

From a theoretical and empirical perspective, the following questions remain open: Do capital inflows improve or exacerbate income inequality? Through what channels do capital inflows affect income inequality within countries? Although we cannot yet fully picture the distributional effects of capital inflows, the empirical findings of this paper provide important directions for future research.

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A Appendix

Table A.1: Countries in the sample

Advanced Economies		
Australia	Austria	Belgium
Canada	Czech Republic	Denmark
Finland	France	Greece
Ireland	Israel	Italy
Korea	Latvia	Lithuania
Netherlands	New Zealand	Norway
Portugal	Spain	Slovak Republic
Slovenia	Sweden	Switzerland
Taiwan	United Kingdom	
Emerging Market Economies		
Armenia	Botswana	Brazil
Bulgaria	Chile	China
Colombia	Croatia	Georgia
Hungary	India	Indonesia
Kazakhstan	Malaysia	Mexico
Moldova	Morocco	Peru
Philippines	Poland	Romania
Russian Federation	South Africa	Thailand
Tunisia	Türkiye	Uruguay

Figure A.1: Impulse responses to an external capital inflow shock in all countries

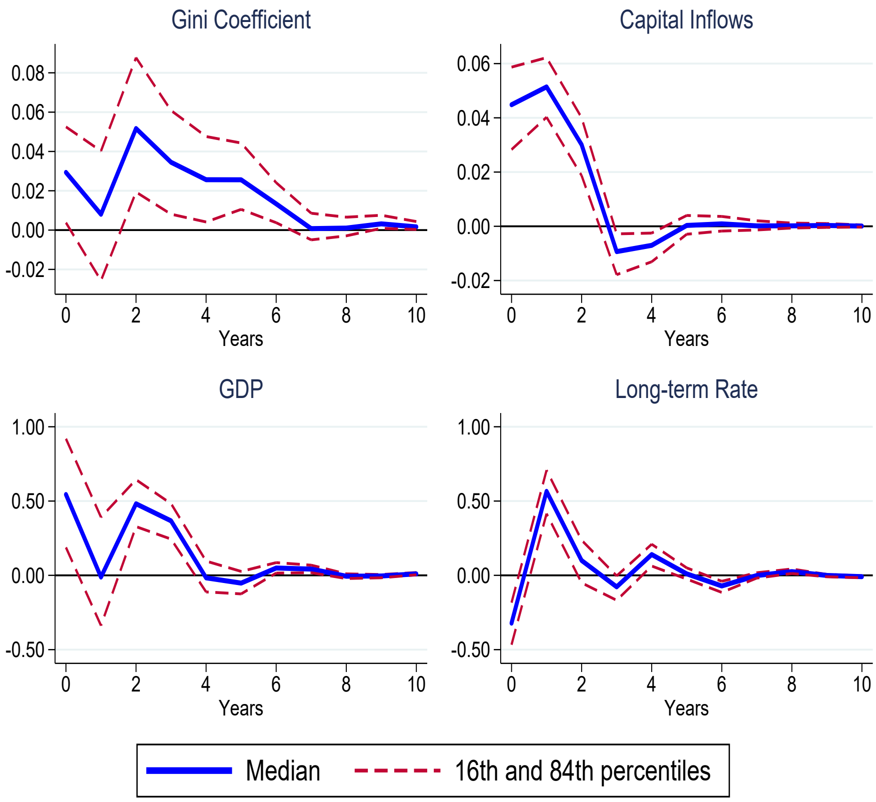


Figure A.2: Impulse responses to an external capital inflow shock in AE (2)

Figure A.3: Impulse responses to an external capital inflow shock in EME (2)

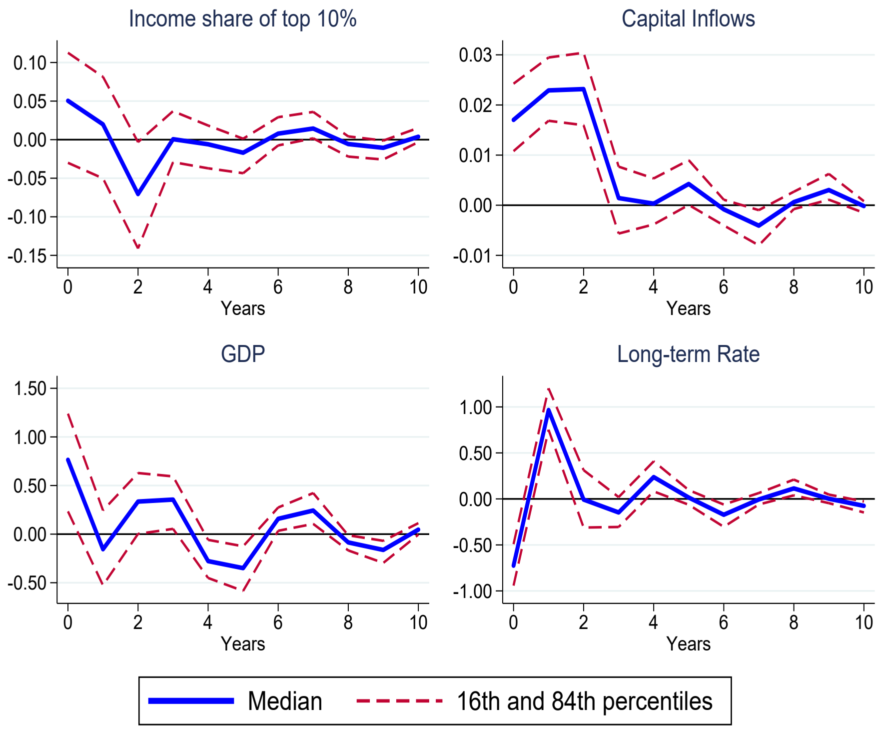


Figure A.4: Impulse responses to an external capital inflow shock in AE (3)

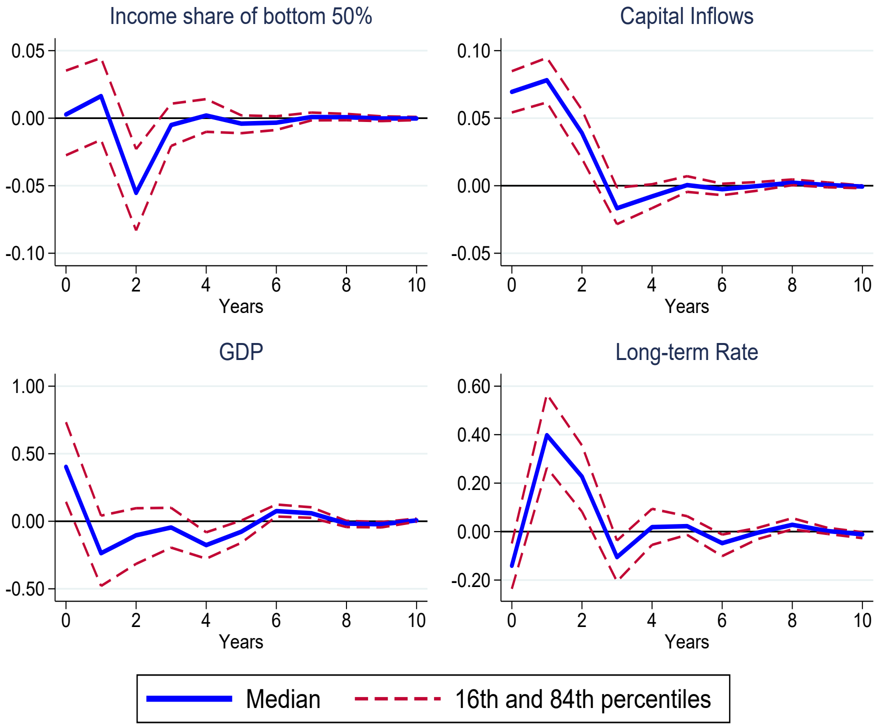


Figure A.5: Impulse responses to an external capital inflow shock in EME (3)

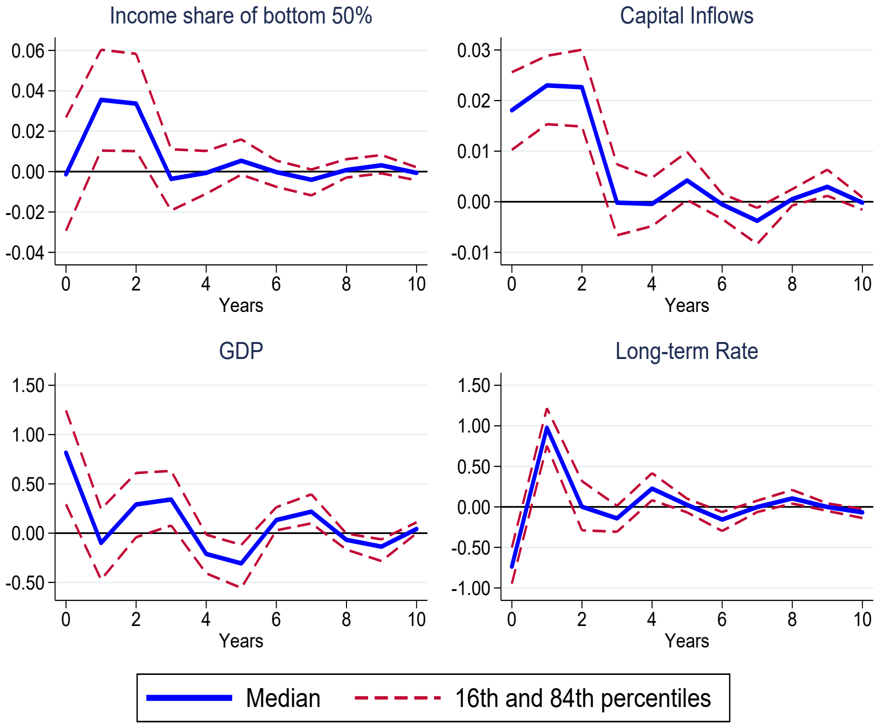


Figure A.6: Robustness check (1) - Gini from EHII in AE

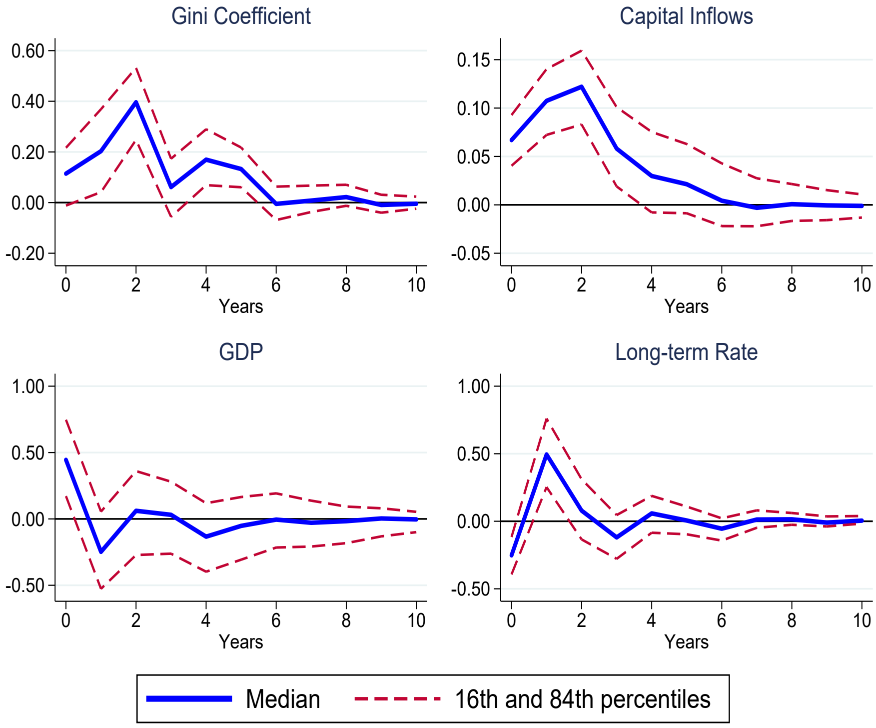


Figure A.7: Robustness check (1) - Gini from EHII in EME

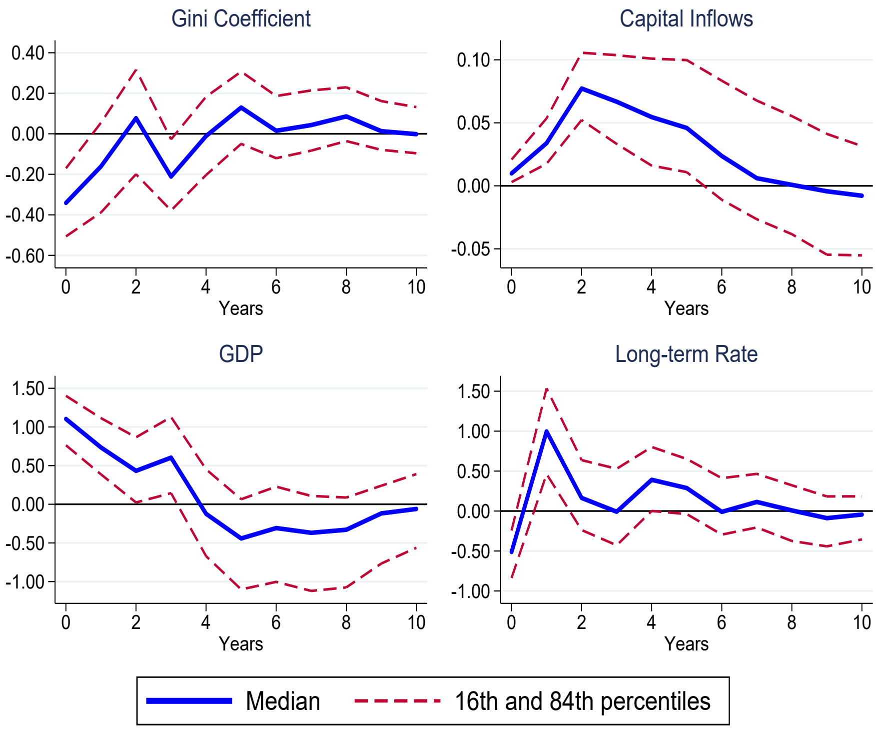


Figure A.8: Robustness check (2.1) - lag one in AE

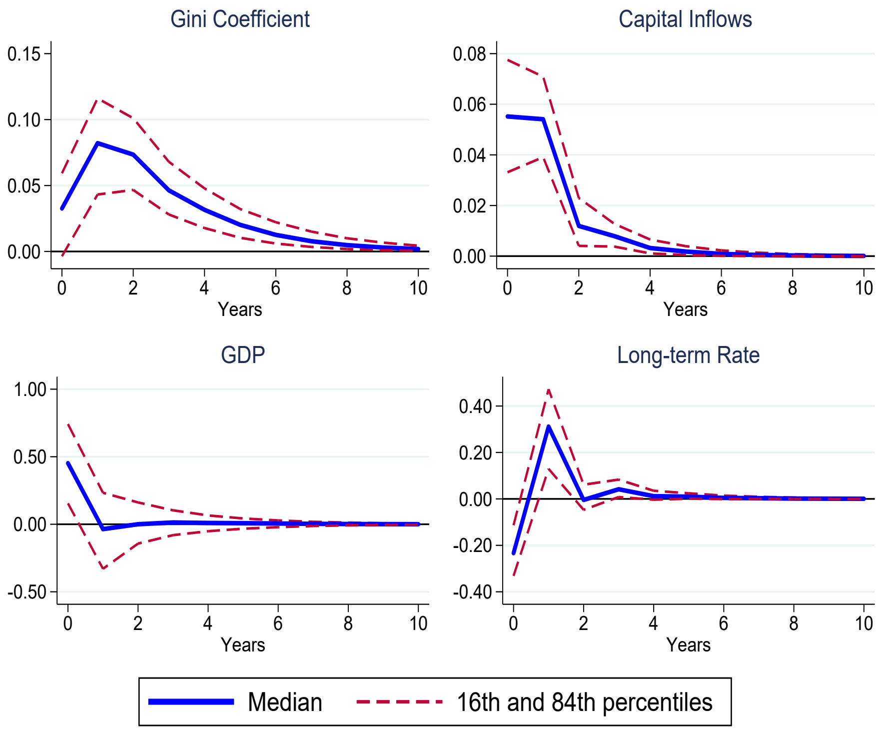


Figure A.9: Robustness check (2.1) - lag one in EME

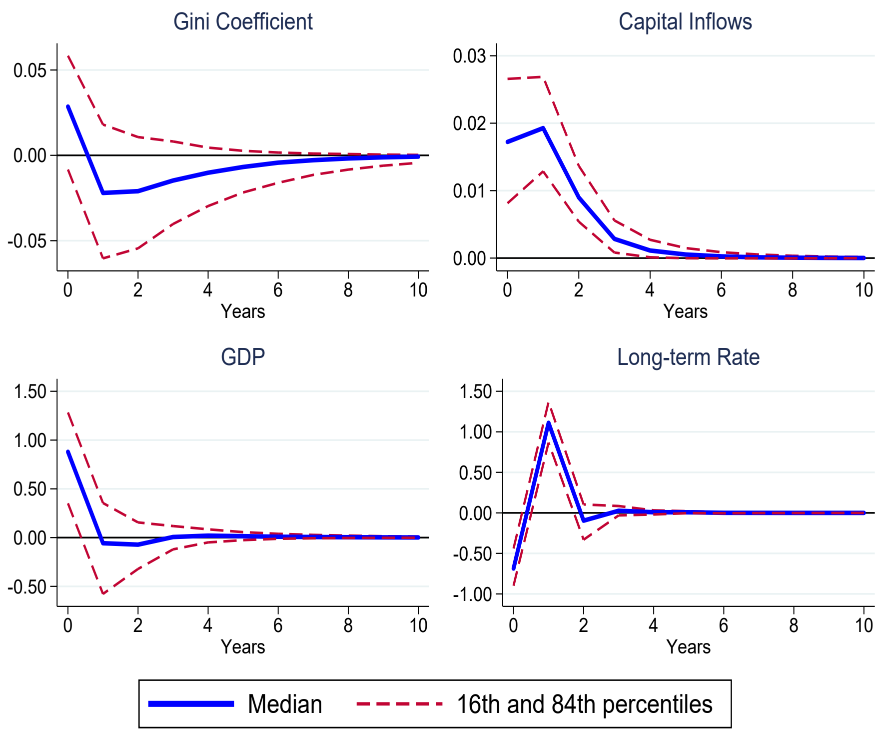


Figure A.10: Robustness check (2.2) - three lags in AE

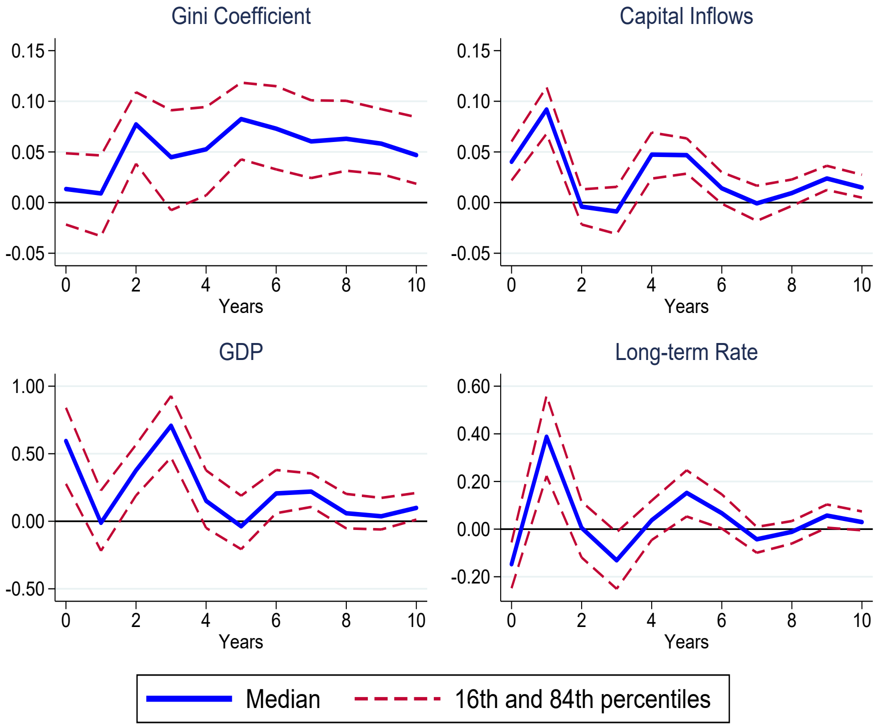


Figure A.11: Robustness check (2.2) - three lags in EME

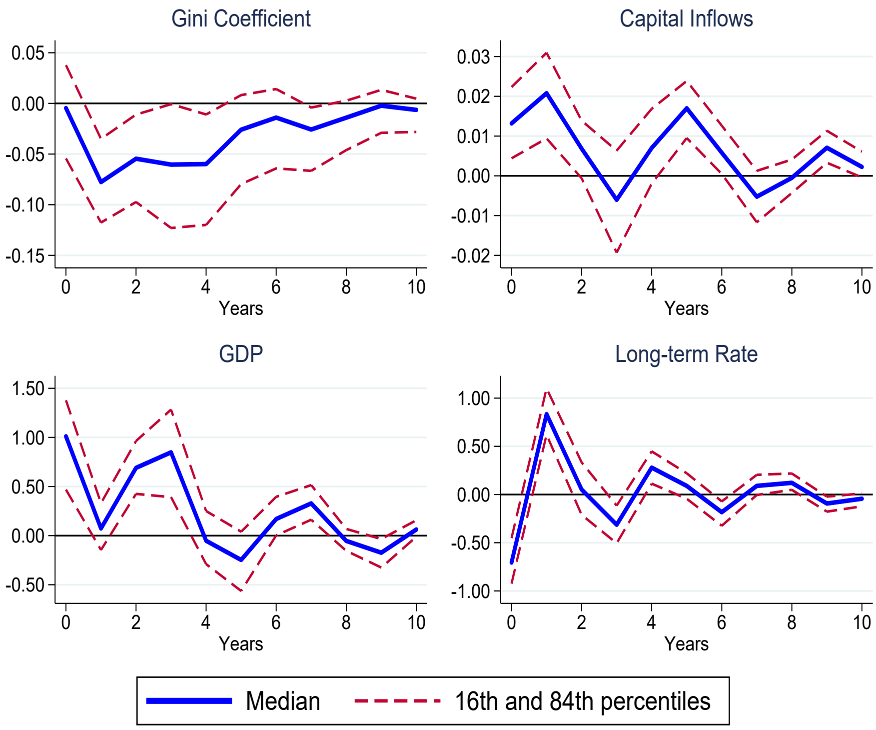


Figure A.12: Robustness check (3.1) - using long-term interest rates in levels in AE

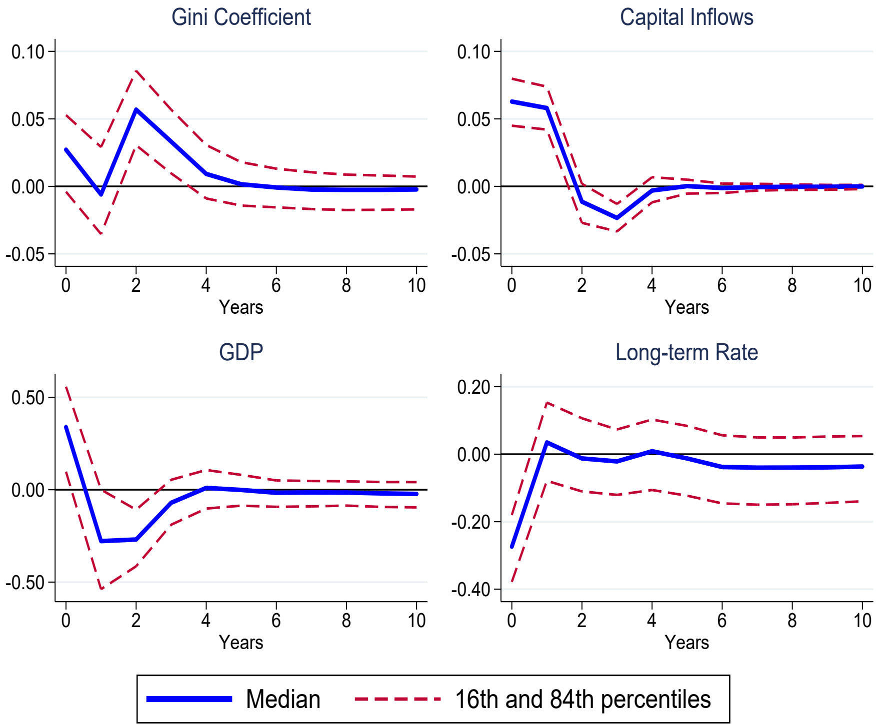


Figure A.13: Robustness check (3.1) - using long-term interest rates in levels in EME

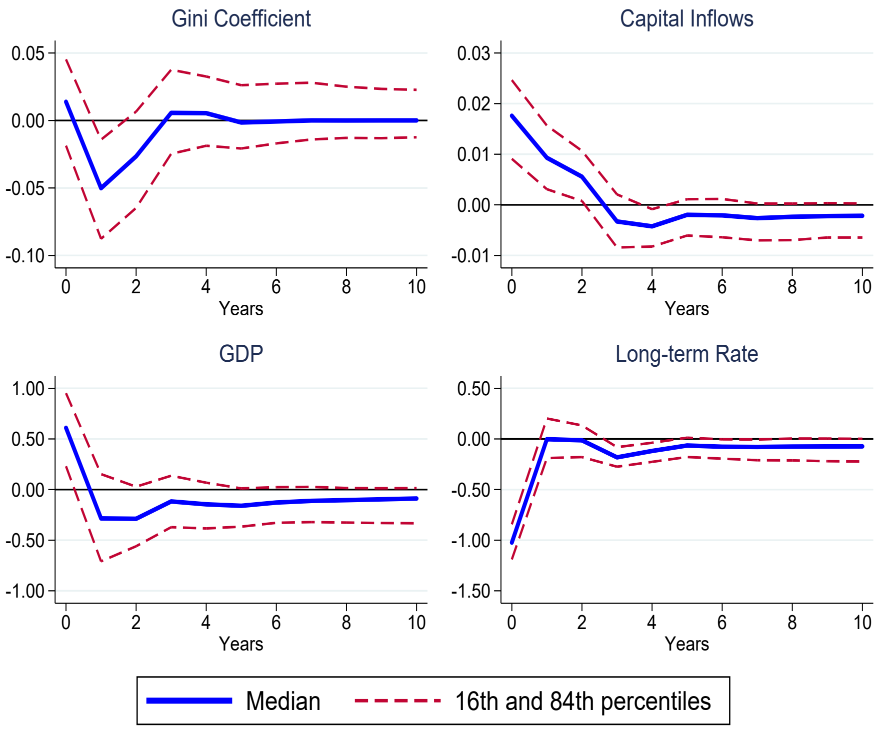


Figure A.14: Robustness check (3.2) - using the Gini in level in AE

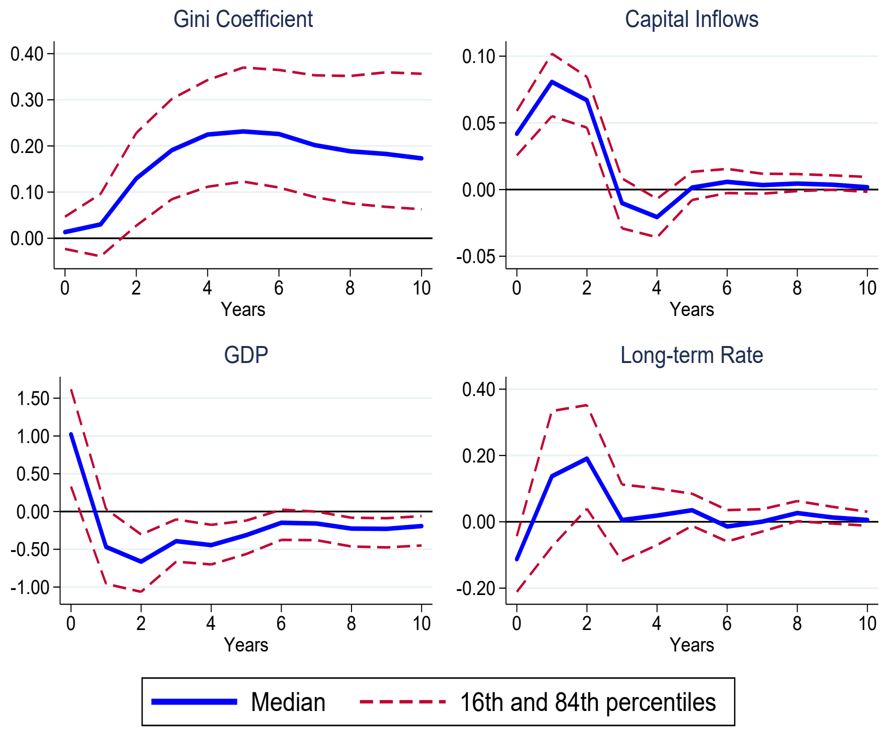


Figure A.15: Robustness check (3.2) - using the Gini in level in EME

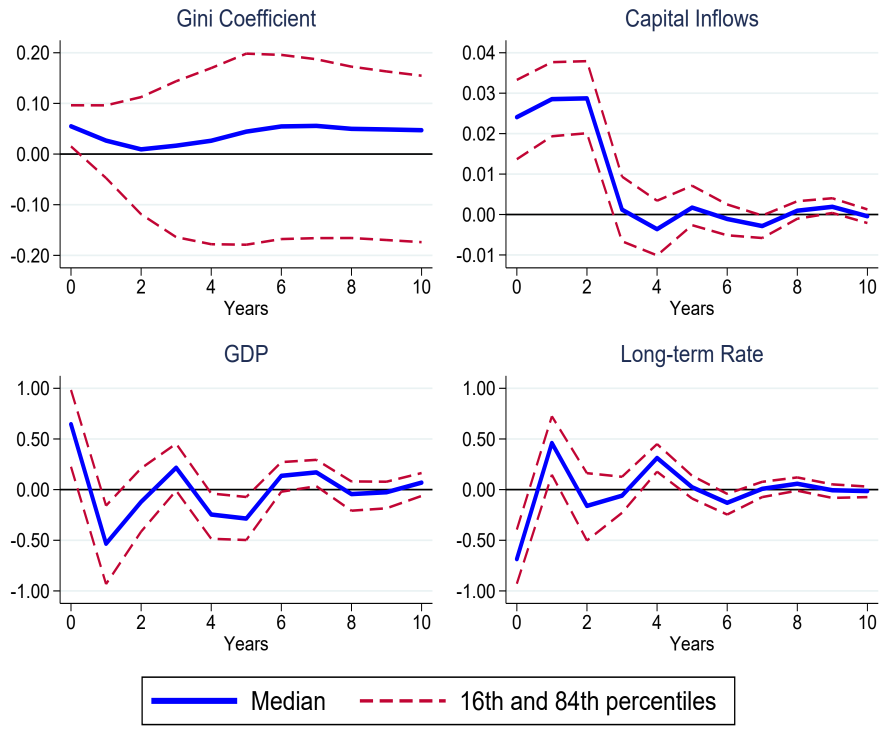


Figure A.16: Robustness check (4) - excluding some countries from the sample in AE

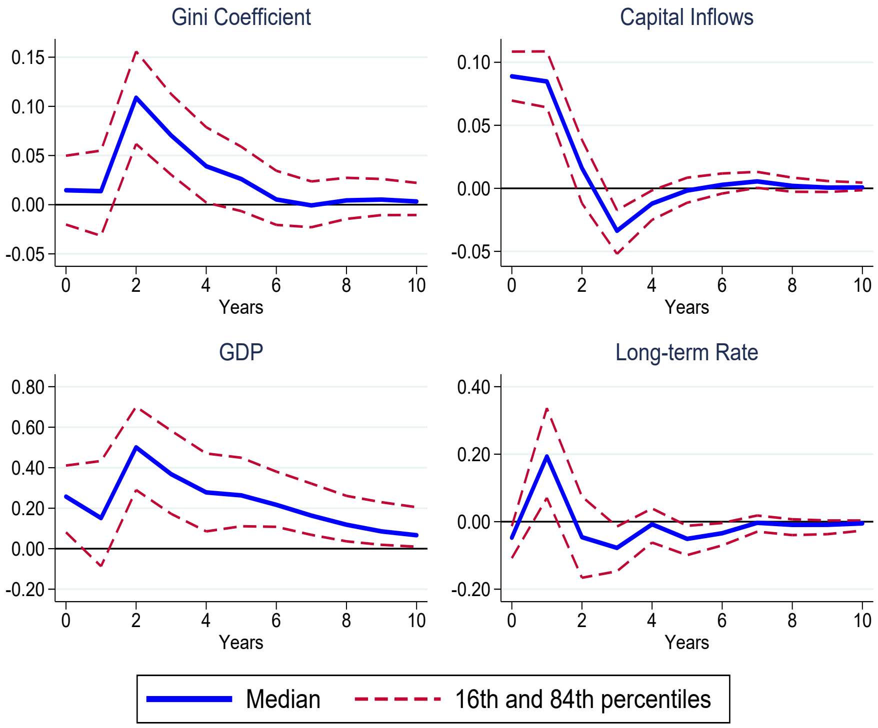


Figure A.17: Robustness check (4) - excluding some countries from the sample in EME

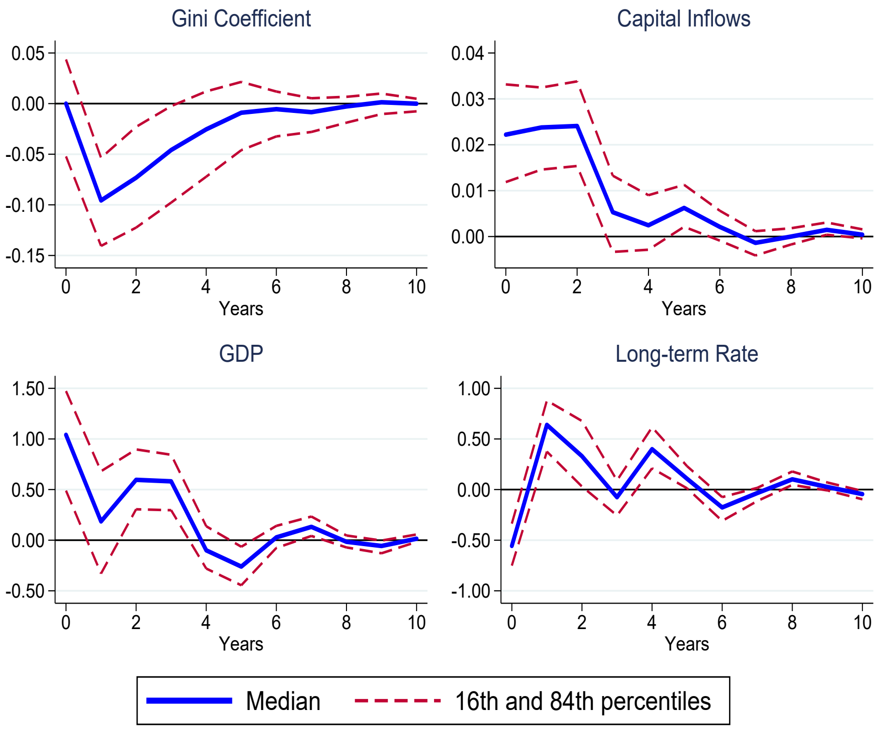


Figure A.18: Robustness check (5) - addressing the impacts of outliers in AE

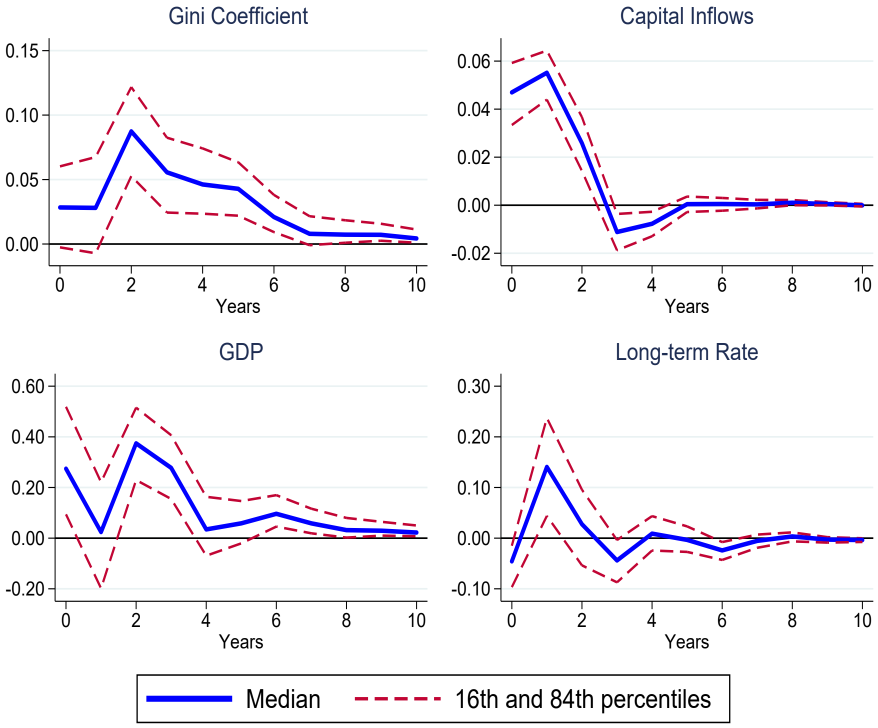


Figure A.19: Robustness check (5) - addressing the impacts of outliers in EME

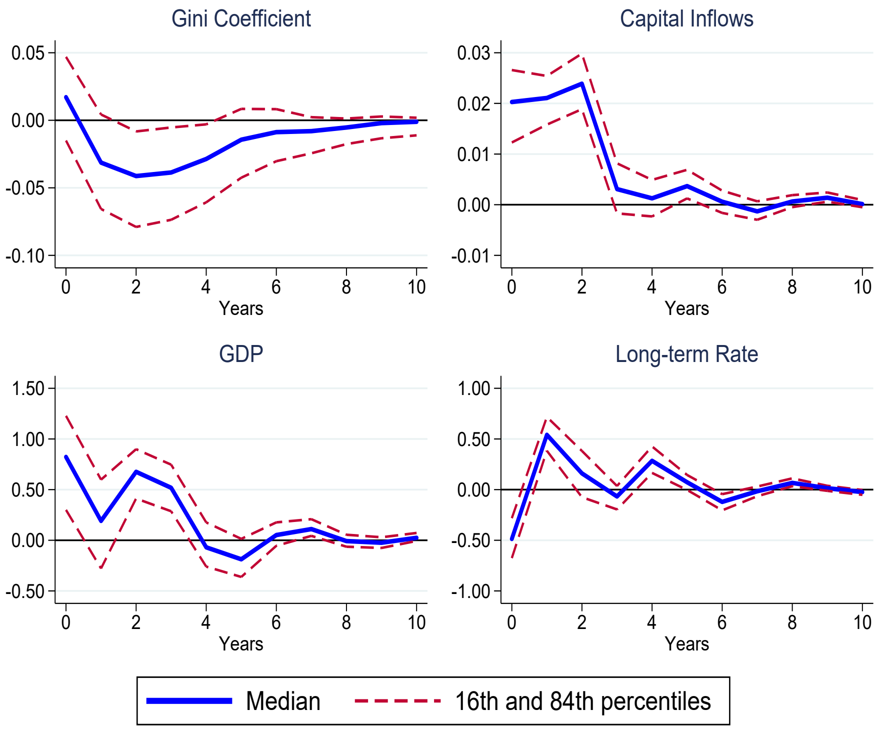
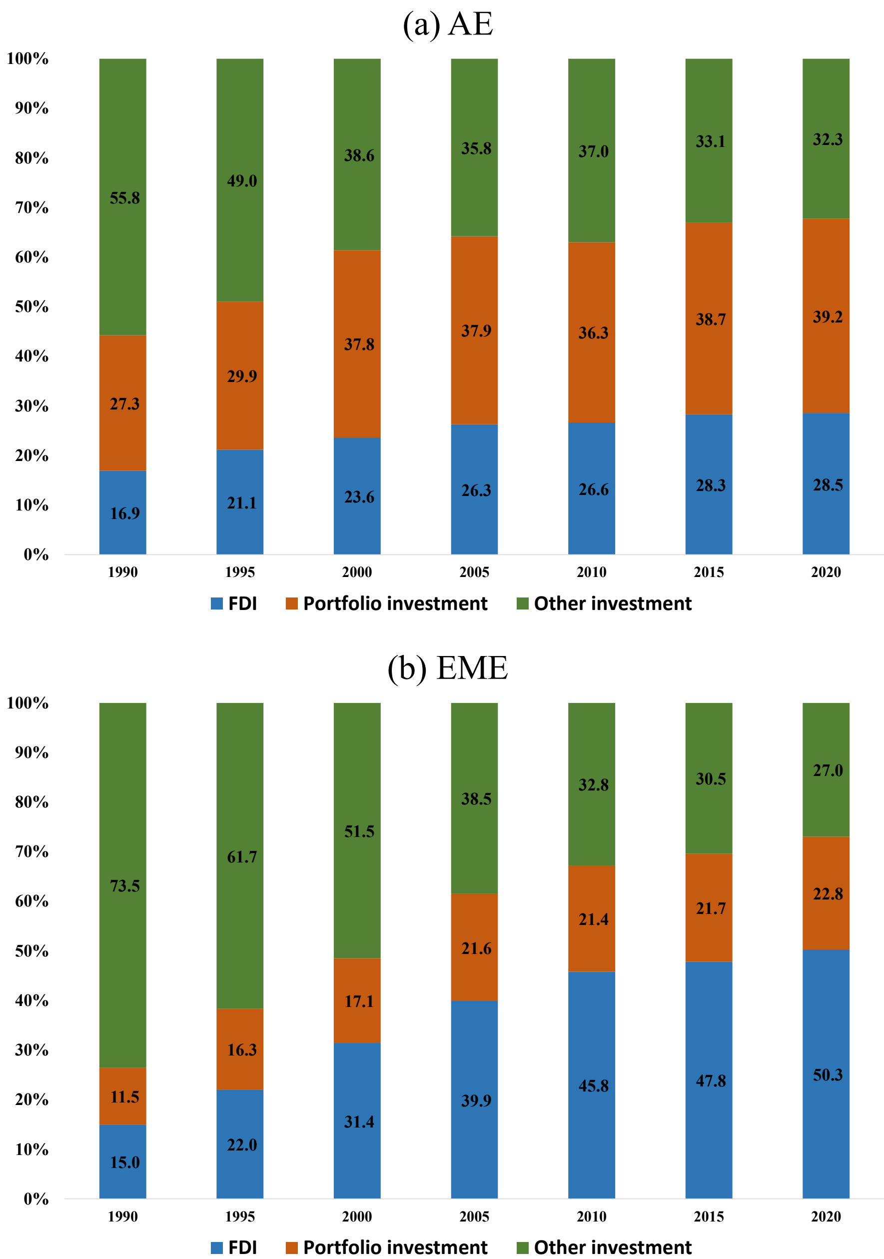


Figure A.20: Composition of external liabilities, 1990-2020



Note: The figure shows averages for countries of the composition of external liabilities as a percentage of GDP.
Sources: EWN II-database from Lane and Milesi-Ferretti (2018) and author's calculations