**Justus Liebig University Giessen**
Department of Economics and Business Studies

*Doctoral Thesis*

# Essays on Financial Integration, Inequality, and Geoeconomics

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*Für meine Eltern,*
*meine Frau, Jooa Kim*
*und meine Tochter, Dahyeon Yun*

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# I. The Distributional Effects of Capital Inflows

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# The Distributional Effects of Capital Inflows*

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**Abstract**

In this paper, I document empirical evidence indicating that an exogenous capital inflow shock increases income inequality in advanced economies and reduces income inequality in emerging market economies. To examine the effects of capital inflows on income inequality within countries, I estimate a panel structural vector autoregression model with annual data for 53 countries over the period 1990 to 2020. I identify the structural shocks in capital inflows using sign restrictions, thereby distinguishing the exogenous shocks driven by global financial conditions from other shocks. The findings of this paper indicate a remarkable difference in the results between advanced economies and emerging market economies, suggesting that the distributional effects of capital inflows vary depending on the economic conditions in the recipient countries. With respect to income class, a capital inflow shock is primarily associated with an increase in the income share of the high-income class in advanced economies and of the low-income class in emerging market economies. It is necessary for policymakers to pay attention to the distributional effects of capital inflows and to design tailored policy frameworks to address them.

**Keywords**: Capital inflows; income inequality; panel VAR; sign restrictions
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## I.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. In contrast, income inequality within countries has increased in many developed and developing countries. It has fueled intense debates about the causes and dynamics of within-country inequality, which has also induced concerns among policymakers and academic researchers.

Much of the literature suggests that *globalization* or *international integration* is one of the main drivers contributing to the rise in income inequality within countries.[1] International integration refers to the process of deeper integration of a domestic economy into global markets. This can be represented by an increase in transactions between domestic and foreign economic agents. The evolution of international integration has developed in two dimensions: trade integration and financial integration. The latter is characterized by an increase in cross-border 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 the volume of cross-border capital flows.

Many economists posit that financial integration (opening the capital account and an increase in capital flows) improves the economic welfare of recipient countries. These benefits are mainly 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 flows can hurt the macroeconomic stability of a recipient country and trigger a financial crisis (Caballero, 2016; Ghosh et al., 2016). The question then arises as to how international financial integration affects income inequality within countries. In recent years, economists have paid more attention to the distributional effect of financial integration. However, in contrast to trade integration, the distributional effects of financial integration have been studied relatively little. The theoretical predictions and empirical evidence on the impacts of capital flows on income inequality are still ambiguous.

[1] Some researchers argue that skill-biased technological change (SBTC) has been a primary cause of the rise in income inequality. However, SBTC does not account for the evolution of inequality after the 1990s. For a comprehensive review of this, see Card and DiNardo (2002).

[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 percent of all capital market liberalization episodes (224 episodes over the period 1970-2010) occurred during the 1990s, and 27.2 percent occurred in the 2000s. Furthermore, the advancement of information and communication technology since the 1990s has made external financial transactions more effortless than ever, which may also be a significant contributing factor in the accelerated integration of international financial markets.

To fill this research gap, I empirically investigate the distributional effects of capital inflows. I employ a panel structural vector autoregression (SVAR) model with a country-level panel dataset and identify exogenous capital inflow shocks driven by global financial conditions rather than domestic economic conditions, using sign restrictions proposed by Faust (1998), Canova and Nicolo (2002) and Uhlig (2005).

The present paper contributes to the empirical literature on the relationship between financial integration and income inequality in three ways. First, I address the issue of the endogeneity of capital inflows and the potential reverse causality between capital inflows and income inequality.[3] In line with the existing literature on the drivers of capital inflows, I distinguish between capital inflows driven by global *push* factors and those driven by domestic *pull* factors.[4] To identify an exogenous capital inflow shock, which is driven by push factors, I adopt sign restrictions following the seminal work of Uhlig (2005). It is assumed that an exogenous capital inflow shock is associated with an increase in global GDP and a decrease in the U.S. and domestic interest rates, while a domestic shock (i.e., pull factors) is not.

Second, I investigate the dynamic impact of the evolution of financial integration using a *de facto* measure of financial integration, namely capital inflows, and a panel SVAR approach. Most of the existing literature has focused on changes in income inequality before and after the episode of capital account liberalization, using *de-jure* measures that reflect the institutional and regulatory environments for cross-border capital transactions. However, in addition to discontinuous policy and institutional changes, the continuous movements of capital flows also matter in explaining the distributional effects of financial integration. To the best of my knowledge, this paper is also the first to employ a VAR model to study the dynamic effects of capital inflows on income inequality. A VAR model is an appropriate approach when there is a lack of theoretical foundation for the dynamic relationships among economic variables. Furthermore, impulse response functions based on the estimated coefficients of a VAR model enable us to investigate the dynamic response of variables when a structural shock emerges in a multivariate system. Because of these advantages, the VAR analysis method is one of the most popular empirical approaches in modern macroeconomics. Additionally, I incorporate panel analysis methodology, including country-specific fixed effects, into a VAR framework to capture time-invariant factors that may affect endogenous variables.

[3]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.

[4]For a comprehensive overview of the drivers of capital inflows, see Koepke (2019).

Finally, I find that the distributional effect of capital inflows differs significantly between advanced economies (AE) and emerging market economies (EME). While AE have integrated into the global market earlier, EME have opened their capital accounts later, resulting in a relatively low degree of financial integration. Moreover, there are significant differences in the level of economic development between AE and EME. For these reasons, I conduct the analysis separately by dividing the sample countries into AE and EME. I show that the exogenous capital inflow shocks increase income inequality in AE, while the shocks have the opposite effect in EME.

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

## I.2 Literature review

The theoretical foundations for the distributional effects of capital inflows and their transmission mechanisms are still not well-established compared to those of trade integration. The existing literature presents conflicting views on the impacts of international financial integration on income inequality. It is, therefore, challenging to theoretically predict how capital inflows affect income inequality.

The prevailing view among economists is that financial integration fosters economic growth. It is widely accepted that capital inflows stimulate domestic investment, enable an efficient allocation of resources, and contribute to technological development. As a result, the incomes of the low-income class may grow faster than the average income, which eventually leads to a more even income distribution.[5] On the other hand, a surge in capital inflows increases the volatility of the macroeconomy and the likelihood of a financial crisis in the recipient countries. The outbreak of a financial crisis can initially affect the rich through a decline in asset prices and corporate insolvencies. However, the economic recession caused by the financial crisis may disproportionately hit the low-income class and exacerbate income inequality (de Haan & Sturm, 2017; Furceri & Loungani, 2018). This asymmetric effect is because the low-income class has poorer working conditions than the high-income class, and they typically have no buffer to absorb the income shock.

[5]The relationship between economic growth and income inequality may appear differently depending on the level of economic development. Based on long-term historical data, Kuznets (1955) posits an inverted U-shaped relationship between income inequality and economic development. This argument suggests that income inequality increases in the early stage of economic growth and decreases during the mature phase of economic development.

Improvements in financial conditions are another transmission channel for the distributional effects of capital inflows. Increased international capital movements reduce borrowing costs in recipient countries, improve access to financial resources for the low-income class, and enhance the efficiency of domestic financial institutions. In the presence of financial market imperfections, such as information asymmetries and high transaction costs, easing financial market regulations and constraints can mitigate these imperfections and benefit low-income households (Beck et al., 2007; Bumann & Lensink, 2016). By contrast, Jaumotte et al. (2013) argue that the benefits of easing restrictions on access to international finance and improving domestic financial conditions accrue primarily to capital owners and the high-income class. Capital inflows make it easier and cheaper for the high-income class to lend for production or investment, allowing them to earn higher future incomes than those with difficulty accessing finance. The low-income class usually has less access to financial markets than the richer classes. It is worth noting that the effects of the improved financial condition may vary depending on the degree of economic and institutional development in recipient countries (Erauskin & Turnovsky, 2019).

Another channel is via the capital-skill complementarity (Asteriou et al., 2014; Larrain, 2015; 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). The more FDI flows into a country and accumulates in the economy, the greater the demand for high-skilled labor in sectors with a high capital-skill complementarity, given that FDI is generally directed towards relatively high-skill sectors in the host country. These effects are related to the wage gap between high-skilled and low-skilled workers.[6]

Several recent empirical studies show that income inequality, primarily measured by the Gini coefficient, increases with capital account liberalization. Larrain (2015) find that opening the capital account increases wage inequality between high-skilled and low-skilled workers using panel data for 20 developed countries from 1975 to 2005. This study provides empirical evidence that the distributional effects of capital account opening are more pronounced in industries with high dependence on external finance and strong capital-skill complementarity. 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. They argue that financial liberalization can lead to a more equal income distribution if financial depth (measured by domestic credit to the private

[6] It is worth noting that much literature also points out that the distributional effects of FDI vary depending on the economic conditions in recipient countries and the types of FDI. For a comprehensive review of this issue, see (Eichengreen et al., 2021).

sector over GDP) is high. On the contrary, the empirical findings of de Haan and Sturm (2017), based on panel data for 121 countries from 1975 to 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, using panel data for 149 countries from 1970 to 2010. Their study also argues that capital account liberalization weakens workers' bargaining power over employers, thereby reducing the wage share. Furceri et al. (2019) 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. 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, 2024). They document that financial integration drives income inequality.

The empirical studies discussed above mainly focus on changes in income inequality before and after the episode of capital account liberalization, using *de-jure* measures such as the Capital Account Openness Index (KAOPEN Index) developed by Chinn and Ito (2006, 2008), or the Financial Reform Index from Abiad et al. (2010). Few studies are using *de-facto* measures, such as the sum of total external liabilities and assets. However, the continuous movement of capital flows, along with the discontinuous institutional and regulatory changes associated with cross-border capital transactions, is also an important aspect of financial integration. In this paper, I focus on the dynamic impact of cross-border capital inflows on income inequality using the *de-facto* measure, which allows for an assessment of the evolution of international financial integration.

The endogeneity of capital inflows is a critical issue in examining the macroeconomic impacts of capital inflows. Capital inflows are affected by the various domestic conditions of a recipient country, such as output, institutional quality, and political stability. To address this endogeneity problem, some studies employ an Instrumental Variable (IV) estimation. Liu et al. (2023) instrument capital flows using the movements in the U.S. two-year Treasury yields as an instrumental variable. Dorn et al. (2018) and Lang and Tavares (2024) also investigate the effects of international financial 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 the relevant variables as endogenous and identify an external shock to capital inflows using sign restrictions, which allows me to address the endogeneity of capital inflows.

# I.3 Empirical methodology

## I.3.1 Data

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

The multivariate model includes six variables: (i) the year-over-year (YoY) difference in the U.S. 10-year Treasury yields that serve 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 YoY movements in domestic long-term interest rates (typically 10-year government bond yields), and (vi) the annual change rate of the Gini coefficient.[7] The last variable, which captures the change in income inequality within countries, is the key 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.[8]

To measure income inequality, I employ the Gini coefficient, which is the most commonly used in the existing literature. The Gini coefficient allows us to compare income inequality between countries and to see how income inequality within a country has changed over time. A large body of literature studying the cause of income inequality uses data 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, 2020), primarily based on WIID but supplemented

[7]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 I.4.2.

[8]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.

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 coefficients for 198 countries from 1960 to 2021.[9]. 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 primarily use the market Gini to rule out the impact of the government's redistribution policies.

It's worth noting that there are a few caveats to using the Gini coefficient. First, the Gini coefficient measures the overall level of income inequality; therefore, it is not sufficient to explain the detailed changes in income inequality. Other measures, such as income share by income class, can provide more pertinent information for income distribution. To address this limitation, I also 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. The 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. The second issue concerns the underlying raw data used to compile the Gini coefficient. The two databases mentioned above are widely used in empirical research because they cover numerous countries and time series. However, the problem is that some countries (mainly developed countries) use income data to measure the Gini coefficient, while others (mainly developing countries) use consumption expenditure data due to limited data availability. Another issue with the Gini coefficient is that some countries use household data while others use individual data. These problems could limit cross-country comparability. To mitigate these issues, I incorporate country-specific fixed effects into the estimation model and conduct a robustness check using the Gini coefficient from the other database.

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 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 External Wealth of Nations Marks (EWN) dataset from Lane and Milesi-Ferretti (2007) and its update (Lane & Milesi-Ferretti, 2018). The database contains

[9] I use Version 9.4 of the database, published in November 2022, for estimation (`https://fsolt.org/swiid/`)

estimates of international assets and liabilities for 212 countries from 1970 to 2021. These estimates are 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 World Development Indicator (WDI) database of the World Bank.[10] Data on long-term nominal interest rates were obtained from the CEIC database. Table I.1 shows descriptive statistics of included domestic variables.[11]

Table I.1: Descriptive statistics

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

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

It is important to note that there are significant differences between AE and EME with regard to their level of economic development and degree of international financial integration. This raises the question of whether it is reasonable to examine these two groups together. Table I.2 shows differences between the two groups in 1990 and 2020. In 1990, the GDP per capita of AE was approximately seven times

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

[11] As shown in Table I.1, variables include large variability with some outliers. To address the potential bias due to extremely large or small observations, I conduct a robustness check using the winsorized data in Section I.4.2.

that of EME, on average. Even as of 2020, despite the rapid growth of EME over the period, the GDP per capita of AE was around \$40,000 on average, nearly five times that of EME. The degree of international financial integration is quantified through two types of measures: the KAOPEN Index, which is a *de jure* measure, and total external liabilities over the GDP, which is a *de facto* measure. As revealed by the KAOPEN Index, AE have sustained high levels of financial openness over the period. In contrast, EME were financially less open in 1990 and have remained at lower levels of financial openness than AE in 2020, despite the steady increase in capital account openness over the past decades. The volume of external liabilities shows a broadly similar evolution to that of the KAOPEN Index. In 1990, the external liabilities as a percentage of GDP were considerably higher in AE than in EME. This differential expanded further over the period. These notable differences provide a rationale for examining the distributional effects of capital inflows by dividing the sample countries into two groups.

Table I.2: 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 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 WDI database of the World Bank, the Chinn and Ito (2006) database (`https://web.pdx.edu/~ito/Chinn-Ito_website.htm`), and the EWN database from Lane and Milesi-Ferretti (2018).

### I.3.2 Panel SVAR model

To examine the distributional effects of capital inflows, I estimate the following reduced-form panel VAR model of order $p$:

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

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

In a dynamic panel model, classical ordinary least square (OLS) estimation with individual fixed effects no longer provides unbiased estimators (Nickell, 1981). This is because the fixed effects are correlated with the regressors, which are the lag values of the dependent variables. To address this problem, I employ a generalized method of moments (GMM) approach proposed by Arellano and Bover (1995) and apply the forward orthogonal transformation to equation (I.1). The forward orthogonal transformation removes the forward mean, i.e., the mean of all the future observations for each country and year.[12] 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.[13]

The key challenge in studying the impulse response of variables in a VAR system is to identify structural shocks, 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:

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

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 (I.1) of the reduced form panel VAR is linearly connected with the vector of orthonormal structural shocks $\epsilon_{i,t}$ in equation (I.2) as follows:

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

[12]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 $\alpha_{i,t} = \sqrt{T_{i,t}/(T_{i,t}+1)}$.

[13]In addition, this transformation procedure minimizes data loss in unbalanced panel datasets. For balanced panels, the first difference and forward orthogonal transformation yield the same estimated results when the instruments are fixed (Arellano & Bover, 1995).

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

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

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

The most popular strategy for identification is to impose zero restrictions on the matrix $\Theta$. This approach transforms the reduced-form panel VAR into a *recursive* form since the matrix $\Theta$ obtained by the Cholesky decomposition of the covariance matrix $\Sigma_u$ 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 shocks using the recursive strategy might be arbitrary.

With zero restrictions, I use sign restrictions proposed by 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, $\Theta$ is defined as the product of the matrix $P$ (obtained by the Cholesky decomposition of $\Sigma_u$) and the orthogonal matrix $Q$. Since $QQ' = I$, $\Theta$ can be represented as follows:

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

It must be checked that the impulse response functions computed with $Q$ satisfy 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 $\tilde{A}_1$, ... , $\tilde{A}_p$ using Normal-Wishart prior in the reduced form parameters, the orthogonal matrix $Q_i$ is also drawn such that the structural parameters satisfy sign and 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 $\tilde{\Sigma}_u$, $\tilde{A}_1$, ... , $\tilde{A}_p$, 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.[14]

### I.3.3 Identification

To some extent, capital inflows are endogenous to economic conditions in recipient countries. Such conditions include fluctuations in output, political stability, and the quality of institutions, which also affect income inequality. Some studies also argue that the level of income inequality can affect capital inflows (Dorn et al., 2018; de Ferra et al., 2021). This argument suggests a potential reverse causality between capital inflows and income inequality. For this reason, in order to examine the distributional effects of capital inflows, it is essential to identify exogenous capital inflow shocks that are orthogonal to domestic economic shocks associated with an increase in capital inflows.

Since the seminal works of Calvo et al. (1993) and Fernandez-Arias (1996), much of the literature distinguishes the drivers of cross-border capital flows between the *push* factors and the *pull* factors (Fratzscher, 2012; Ahmed & Zlate, 2014; Forbes & Warnock, 2012; Koepke, 2019; Byrne & Fiess, 2016). Following the framework of this strand of literature, I identify the structural exogenous shocks to capital inflows in the multivariate system as the shocks driven by global financial conditions, i.e. push factors. The push factors pertain to an unexpected increase in the demand for domestic financial assets by foreign investors, which is unrelated to the domestic economic situation. These exogenous shocks to capital inflows are heavily influenced by monetary policy in mature economies such as the U.S. and related market volatility, as well as the risk aversion of global investors. Furthermore, an unexpected increase in global savings also leads to surges in capital inflows.

[14]In this paper, the number of accepted draws, $N$, is set to 1,000.

Building on the empirical findings of the existing literature on push and pull factors of capital inflows, I impose the sign and zero restrictions listed in Table I.3 on exogenous capital inflow shocks, as well as domestic shocks. As discussed in Uhlig (2005), 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. I impose that exogenous capital inflow shocks are associated with an increase in world GDP and a fall in U.S. and domestic long-term interest rates. I impose no restrictions on the response of the Gini coefficient, which is the main variable of interest.[15] The restrictions are imposed only upon impact. A restriction over two years is too restrictive.

As discussed in Tillmann (2013), Sá et al. (2014) and Hristov et al. (2020), imposing restrictions on long-term interest rates is crucial for distinguishing capital inflow shocks driven by push factors from other shocks, such as a domestic demand or productivity shock.[16] In an open-economy model, an unexpected increase in global savings trigger higher capital inflows into small open economies, which leads to a fall in world and domestic interest rates. Likewise, an expansionary monetary policy pursued by the monetary authorities of mature economies increases capital flows into other economies. A fall in interest rates in mature economies boosts the demand for domestic assets, as these assets become more attractive. This, in turn, leads to an increase in capital inflows, which subsequently lower the domestic interest rate. The response of domestic interest rates to a capital inflow shock driven by pull factors depends on the source of the shock. That is, whether the shock originates from the demand side, the supply side, or the monetary authority, determines the direction of the response. Theoretically, demand shock and monetary policy shock, which are associated with an increase in capital inflows, typically put upward pressure on domestic interest rates. Conversely, a domestic supply shock can exert downward pressure on domestic long-term interest rates. No restriction, therefore, is imposed on the response of domestic long-term interest rates to domestic shocks.[17] Additionally, the capital inflow 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.

[15] This approach is referred to as the "agnostic approach", which does not impose any restrictions on the impulse response of the variables of interest. It relies on data rather than restrictions to investigate dynamic response pathways of the variables of interest (Uhlig, 2005).

[16] Given that central banks primarily determine short-term interest rates, it would be more appropriate to impose restrictions on the response of long-term interest rates. It can be assumed that long-term rates are determined to a greater extent by the market than short-term rates.

[17] 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 I.4.

To ensure the validity of these restrictions, I excluded some large economies (Germany, Japan, and the U.S.) from the sample countries for the estimation of the model. The financial assets of these countries are regarded as global safe-haven assets. Therefore, the movements and economic consequences of a surge in capital inflows into these countries are significantly different from those in other countries, even during the same phase of the global financial cycle.

Table I.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 |

## I.4 Results

This section presents the estimated impulse response functions from the panel SVAR 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 I.4.2. The results reported in this section focus on the impact of exogenous capital inflow shocks ("push" factors) on income inequality and other domestic variables. The reason 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 include various disturbances such as monetary, demand, and supply shocks, domestic shocks can not be accurately identified by imposing the sign restrictions in Table I.3.

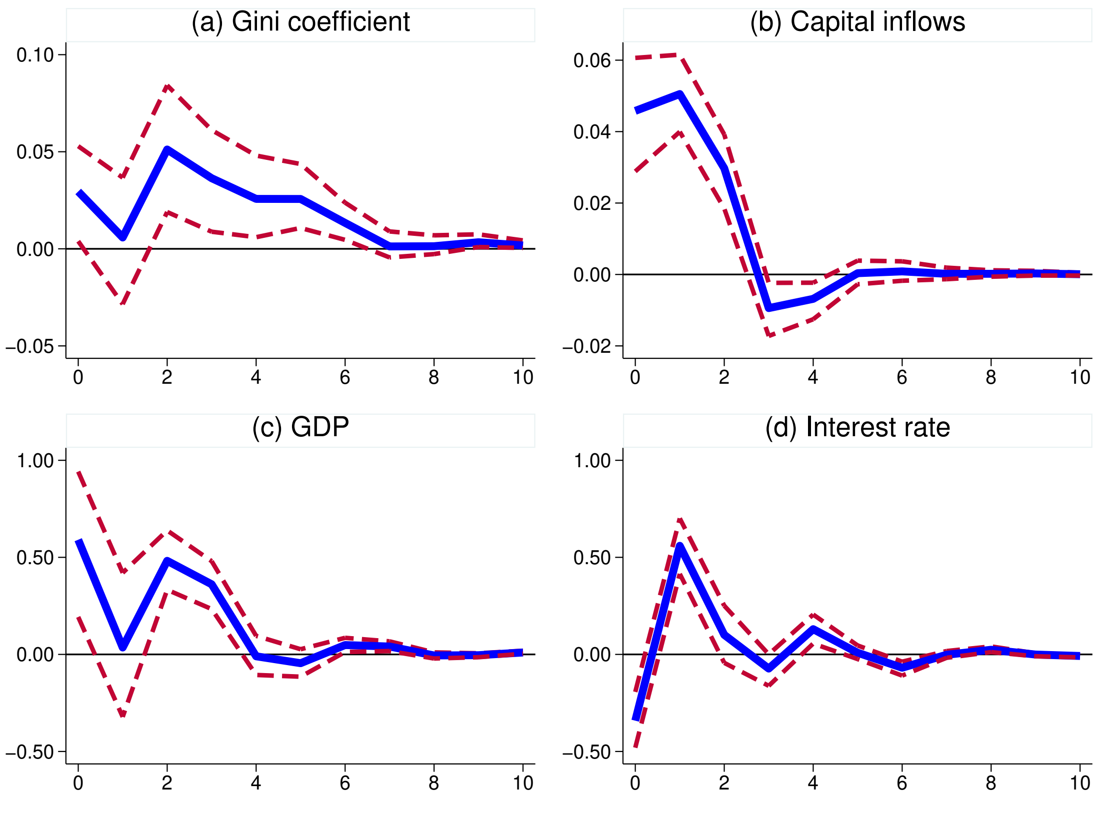
### I.4.1 Baseline results

Figure I.1 shows the impulse response functions to a one-standard-deviation shock in capital inflows over ten years for all sample countries.[18] The solid blue line represents

[18] Assuming identical coefficients across countries, these impulse response functions can be interpreted as the average effect of a capital inflow shock within the sample countries.

the median of impulse response functions across all independent draws that satisfy the sign and zero restrictions in Table I.3. The dashed red lines are the confidence bands formed using the 16th and 84th percentile of all accepted draws. The upper left panel of Figure I.1 shows the impulse response of the Gini coefficient to an external capital inflow shock, where an unexpected increase in capital inflows by about 5% of GDP. As shown in the figure, a capital inflow shock is associated with an increase in the YoY change rate of the Gini coefficient by about 0.05 percentage points two years after the shock. This result is statistically significantly different from zero and also economically significant because income inequality changes very gradually over time. The effects of capital inflow shocks on income inequality last for a long time after two years, although the magnitude of the responses slowly decreases. This result is consistent with the findings of previous empirical studies, which document that capital account liberalization or a surge in capital inflows is associated with an increase in income inequality within countries.

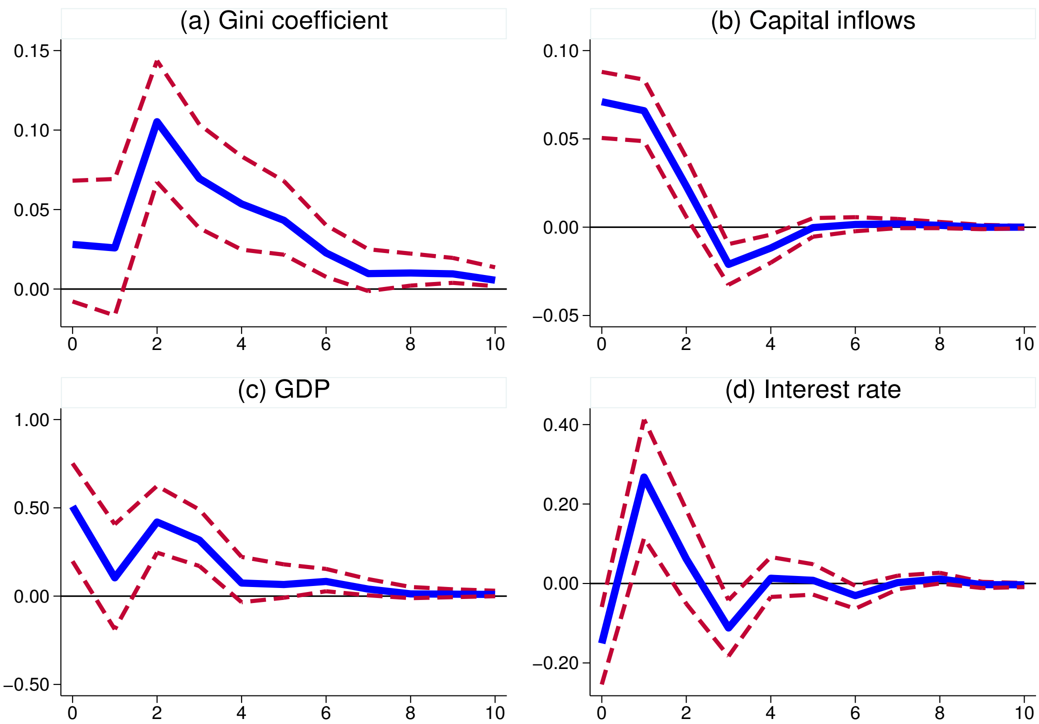
Figure I.1: Impulse responses to an external capital inflow shock in all countries




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figures I.2 and I.3 show the impulse response functions for AE and EME, respectively. The upper left panel of Figure I.2 presents the impulse response of the Gini coefficient for AE to an external capital inflow shock, where an unexpected increase in capital inflows by about 7% of GDP. At the median, a capital inflow shock increases the Gini coefficient, i.e., income inequality. The response of the Gini coefficient's YoY change rate peaks at about 0.1 percentage point two years after 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 I.2: Impulse responses to an external capital inflow shock in AE

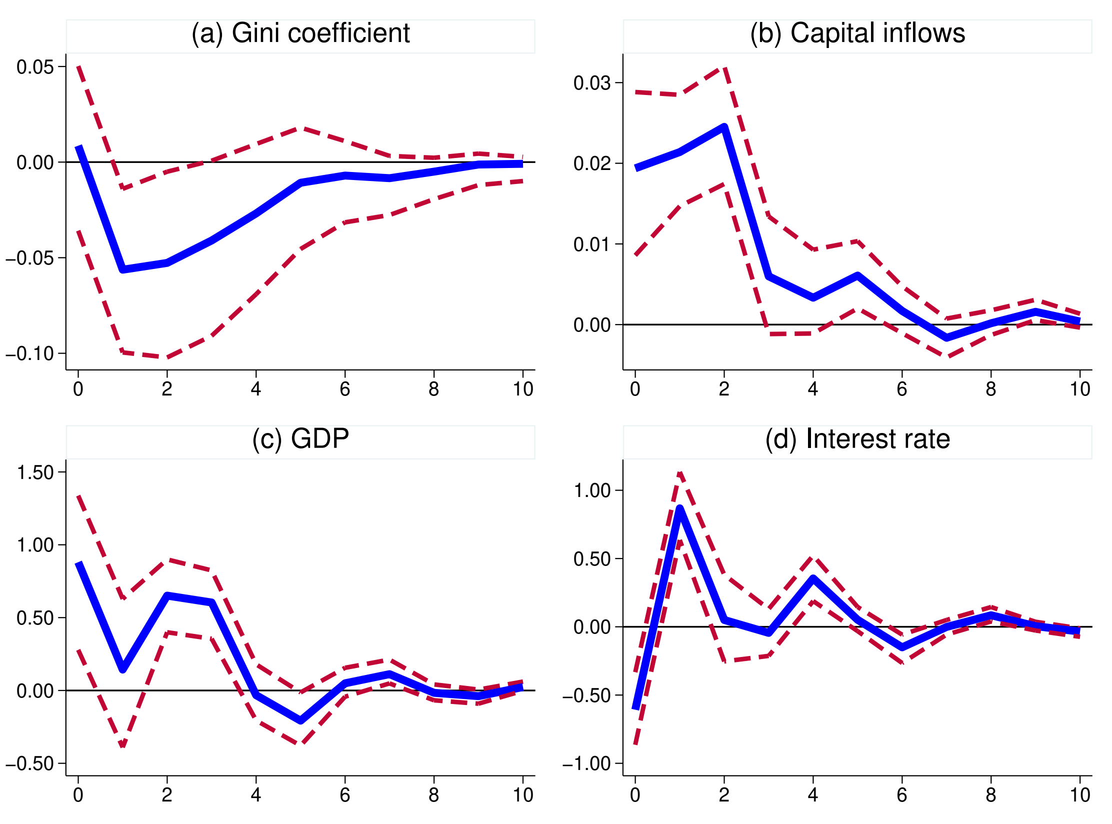



*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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 I.3 presents the results for EME. As shown in the upper left panel, the Gini coefficient falls one year after the shock, where an exogenous increase in capital inflows by about 2% of GDP.[19] 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 significant and also economically significant. Negative responses of income inequality to the shock last for a long time but are no longer statistically significant. Contrary to income inequality, the impulse responses of domestic GDP and long-term interest rates to a capital inflow shock in EME are considerably similar to the previous results for AE.

Figure I.3: Impulse responses to an external capital inflow shock in EME




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

[19]It is less than the magnitude of the capital inflow induced by the identical shock in AE, which is about 8% of GDP. This difference is consistent with data showing that capital inflows typically arise on a larger scale in AE than in EME.

Table I.4 shows the forecast error variance decomposition. At the median, capital inflow shocks explain about 13.6% of the variance in the Gini coefficient in AE and about 10.1% of the variance in the Gini coefficient in EME at a 10-year forecast horizon, respectively.

Table I.4: Forecast error variance decomposition

| | Capital inflow shock | | | |
|---|---|---|---|---|
| | 1Y | 3Y | 5Y | 10Y |
| Advanced Economies | 1.2 | 8.9 | 12.5 | 13.6 |
| Emerging Market Economies | 0.3 | 7.5 | 10.0 | 10.1 |

*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 forecast horizons.

The results presented above 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.

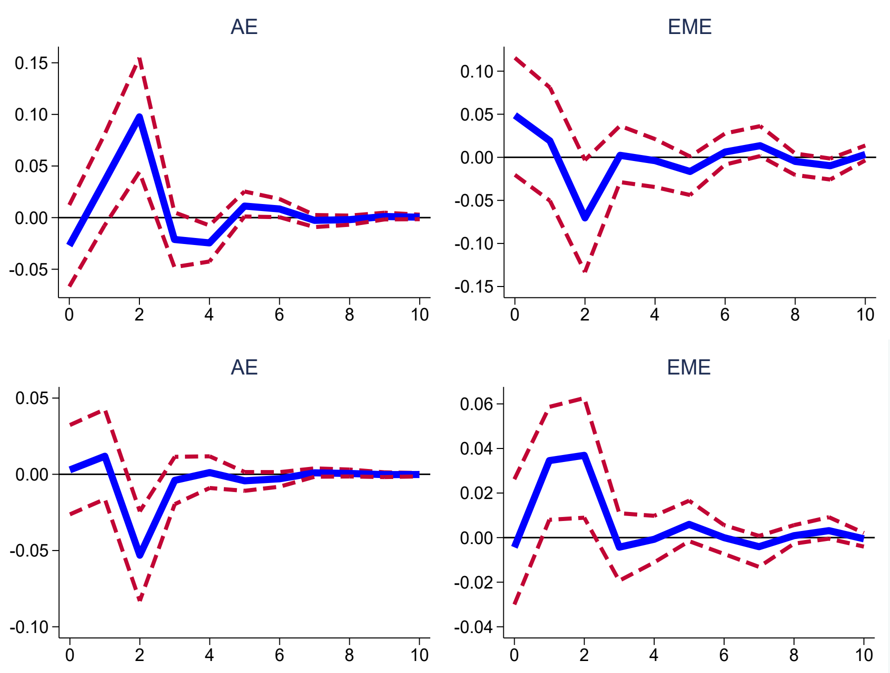
The results using income share data suggest that surges in capital inflows due to external shocks are associated with an increase in the income share of the rich group and a decrease in the income share of the poorest half in AE. By contrast, a capital inflow shock is associated with a decrease in the income share of the rich group and an increase in the income share of the poor group in EME. Figure I.4 shows the impulse response of the income share of the top 10% group and the bottom 50% group.[20] As shown in the upper panel of the figure, a capital inflow shock increases the income share of the top 10% group in AE, while a shock decreases the income share of the top 10% group in EME. The lower panel of the figure shows that the income share of the bottom 50% group responds to an external shock in capital inflows in the opposite direction than above. While a 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 results could support the previous main results using the Gini coefficient.

[20] Figures I.A.1-I.A.4 in Appendix present the impulse response functions for all domestic vari-

Figure I.4: Impulse response of the income share to capital inflow shock

(a) Top 10%

(b) Bottom 50%




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

### I.4.2 Robustness checks

This subsection reports the results of various robustness checks on the baseline results shown in Section I.4.1. Figures I.A.5-I.A.20 in Appendix show the impulse response functions of all domestic variables.

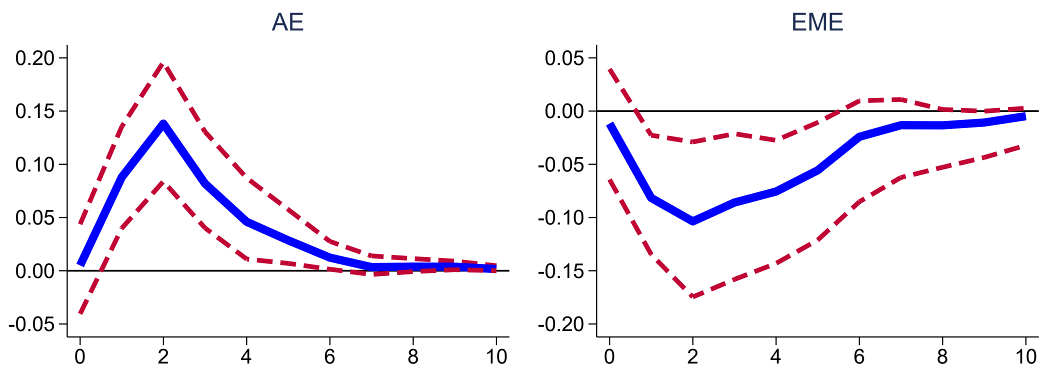
**Alternative measure of income inequality**

Firstly, I test whether the main findings are robust to an alternative measure of income inequality. As discussed in Subsection I.3.1, I use the market Gini from SWIID rather than the disposal Gini, which measures inequality in post-tax and post-transfer income, to rule out the impacts of the government's redistribution policies. Figure I.5 shows that capital inflow shocks have more significant effects on inequality in disposal income than on inequality in market (pre-tax and pre-transfer) income. This result can be attributed to the fact that redistribution policies are unable to moderate the distributional effects of other factors that contribute to an increase or a reduction in income inequality, while they may reduce the overall level of income inequality.

Next, I re-estimate the panel SVAR model using the Gini coefficients from the Estimated Household Income Inequality (EHII) dataset established by the Univer-

ables.

Figure I.5: Robustness check (1.1): disposal Gini




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

sity 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[21] based wage dispersion in industry.[22] This database provides unbalanced panel data with 4,550 annual observations of 154 countries from 1963 to 2015. Figure I.6 shows the impulse responses of the Gini coefficient from EHII. As shown in the left panel, the results for AE are broadly similar to those in the baseline results. An exogenous capital inflow shock increases income inequality, with the peak two years after the shock and the effects gradually decreasing over time. In EME, while the shape of Gini's response differs from that in the baseline results, a capital inflow shock also leads to a statistically and economically significant decline in the change rate of the Gini coefficient.

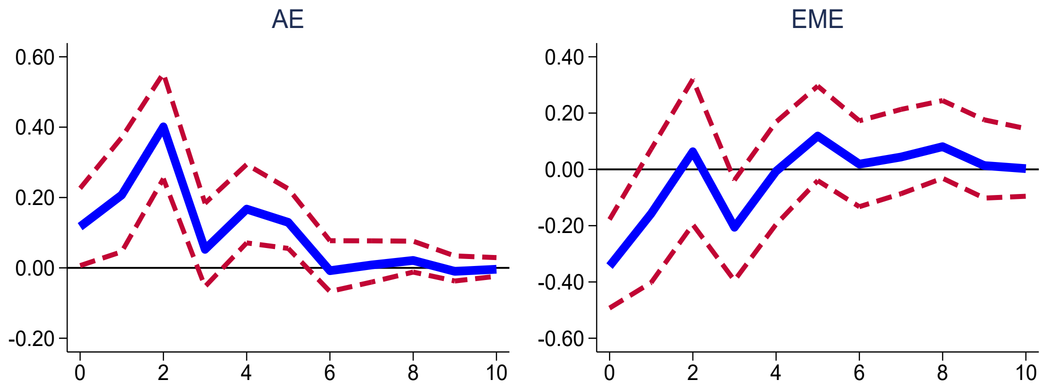
**Different lag lengths**

The estimated impulse response functions can be sensitive to the selected lag length. To test whether the main results are robust to various lag lengths, I re-estimate the panel VAR model with one lag and three lags. The upper panel of Figure I.7 shows the impulse response of the Gini coefficient estimated with lag one. The result in AE is broadly similar to the result in the baseline estimation. A capital inflow shock increases the annual change rate of the Gini coefficient, and these responses of income inequality are statistically significant. In EME, while a capital inflow shock reduces income inequality one year after the shock, this effect is not statistically significant.

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

[22]For a detailed and comprehensive description of the methodology for constructing Gini coefficients from EHII, see Galbraith and Kum (2005).

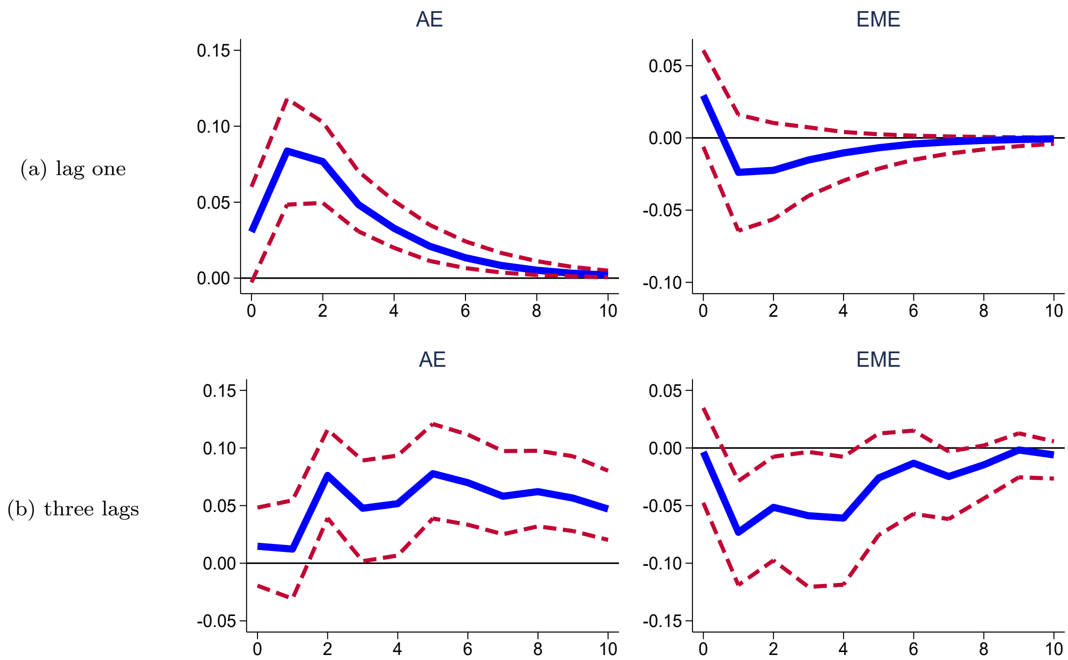
Figure I.6: Robustness check (1.2): Gini from EHII




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

The results in the lower panel of Figure I.7 confirm that the baseline results for AE and EME are robust with three lags. A capital inflow shock leads to an increase in the YoY change rate of the Gini coefficient in AE and a decline in that in EME. The impact of an exogenous capital inflow shock on income inequality is more persistent and statistically significant than in the baseline model.

Figure I.7: Robustness check (2): different lag lengths

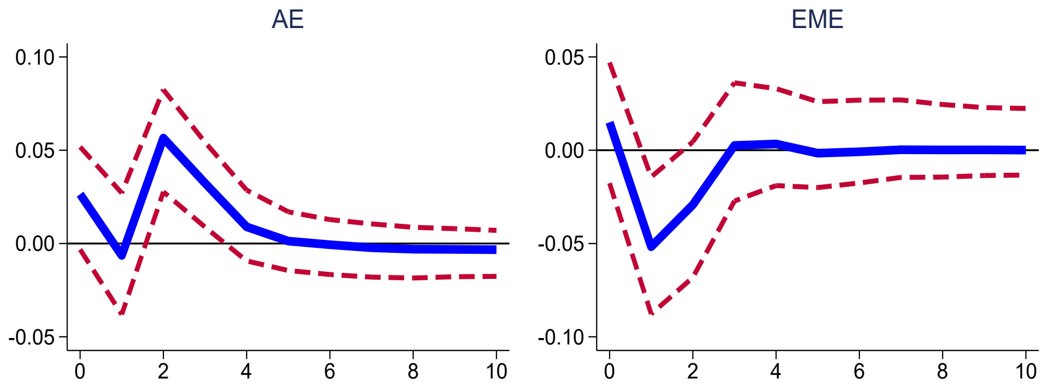



*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

**Using level of variables**

I also test if the baseline results are robust when using the level of some variables rather than the first difference. 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 I.8 shows the 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 I.8: Robustness check (3.1): using long-term interest rates in levels



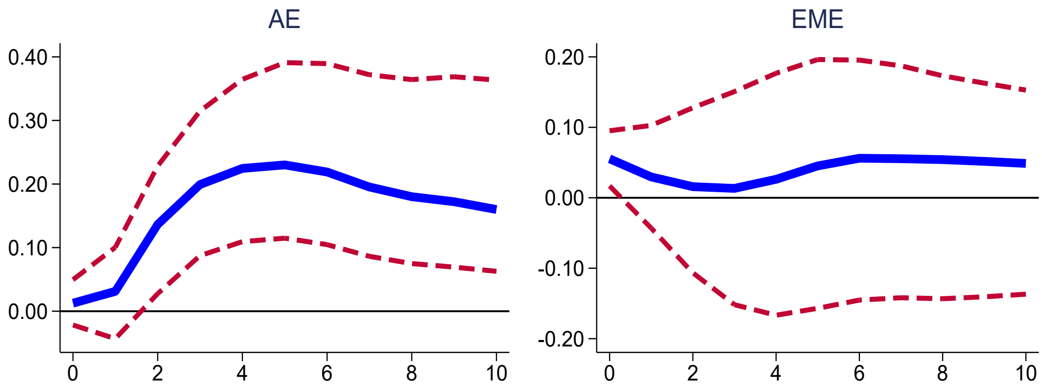

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.9 shows the results using the level of the Gini coefficient. As shown in the left panel, a capital inflow shock has statistically significant and persistent effects on the Gini in AE. In contrast, the right panel shows that the responses of the Gini in EME do not deliver significant outcomes, unlike the baseline results.

**Excluding some countries from the sample**

Although I divide the sample countries into two groups, AE and EME, and include country-specific fixed effects in the model, there may still be cross-country heterogeneity that is not fully controlled. To address the potential bias due to the cross-country heterogeneity, I exclude some countries from the sample. By excluding and regrouping, the sample countries in both groups of AE and EME may become much more homogeneous with respect to the level of economic development and the openness of the capital market than in the baseline estimation. For example, some countries, such as South Korea, were not classified as AE according to the 1990 IMF country classification, but they achieved rapid economic development and are now classified as AE. Other countries in AE, such as the Czech Republic, are transition

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

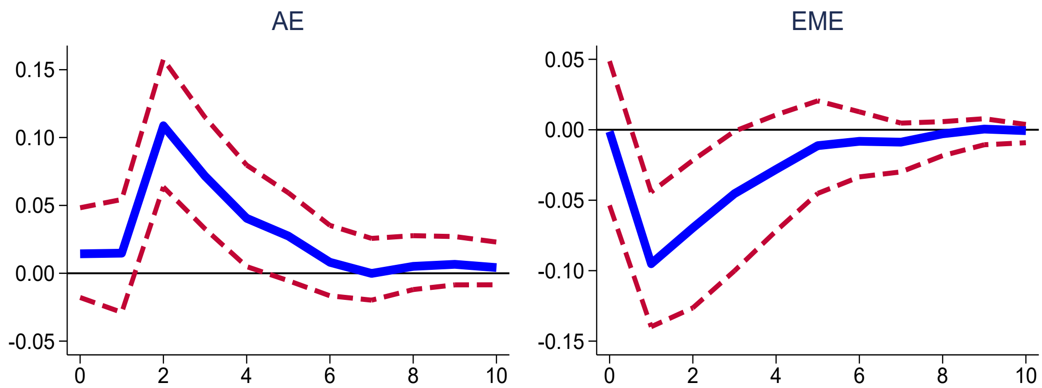



*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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 at a low level of economic development and capital account openness in 2020. The panel VAR model is re-estimated without these countries.[23]

Figure I.10 shows the results of this robustness check. 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. The impulse response of the Gini coefficient in EME is more significant than the baseline results.

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



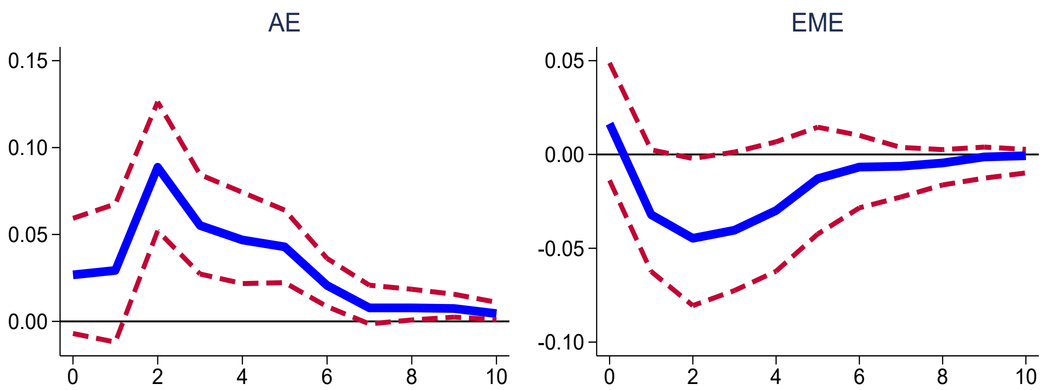

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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

**Addressing the impacts of outliers**

As shown in Table I.1, all domestic variables in the model indicate some variability with outliers. To mitigate concerns about potential bias from outliers, I winsorize the four domestic variables (capital inflows, domestic GDP, domestic long-term real interest rates, and the Gini coefficient) at the 2.5th and 97.5th percentiles. I then re-estimate the model with the winsorized data. Figure I.11 shows that the re-estimated impulse response functions are not significantly different from those of the baseline results.

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




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

### I.4.3 Discussion

In this subsection, I discuss how the empirical findings of this paper contribute to previous debates on the distributional effects of capital inflows. Several empirical studies suggest that capital account liberalization or a surge in capital inflows increases income inequality. Among them, Furceri and Loungani (2018) and Li and Su (2021) provide empirical evidence that capital account liberalization is associated with an increase in income inequality. Liu et al. (2023) find that surges in capital inflows raise income inequality using an IV approach. However, these studies report empirical results for a broad sample of countries without distinguishing between AE and EME, or only for AE (Asteriou et al., 2014; Larrain, 2015). In contrast, I divide sample countries into AE and EME, and find empirical evidence that the distributional effects of capital inflows differ across the two country groups. The main results presented in Subsection I.4.1 reveal that exogenous capital inflow shocks increase income inequality in AE but reduce income inequality in EME.

The adverse impact of capital inflows on income inequality in AE is broadly consistent with the results of previous empirical literature. The results for EME, however, are at odds with the empirical evidence from existing studies. They are more aligned with conflicting theoretical views predicting that capital inflows can *reduce* income inequality. As previously stated in Section I.2, there are contrasting views on the distributional effects of financial integration (capital account liberalization and a surge in capital inflows). The empirical findings presented in this paper can provide support for these two competing hypotheses to some extent simultaneously, implying that the relationship between capital inflows and income inequality may vary depending on heterogeneity across countries. In other words, these findings suggest that there may be a non-monotonic (non-linear) relationship between capital inflows (financial integration) and income inequality. This raises a fundamental question regarding the underlying mechanisms that generate the contrasting outcomes across the two country groups, AE and EME. Why should the response direction of income inequality to capital inflow shocks differ between AE and EME?

The composition of total external liabilities may be relevant to the relationship between capital inflows and income inequality. 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, and trade credits. Figure I.A.21 in the Appendix shows the composition of external liabilities as a percentage of GDP in AE and EME over time. As shown in the figure, there are notable differences in the composition of external liabilities between AE and EME. The proportion of other investments has steadily decreased in both groups since 1990. Conversely, the proportion of portfolio investment has grown mainly in AE (27.3% → 39.2%), and the proportion of FDI has increased significantly in EME (15.0% → 50.3%).

Harms et al. (2024) document a high correlation between income inequality and the equity share of external liabilities.[24] With regard to the financial structure, more finance increases income inequality if finance expands via market-based financing (Brei et al., 2023). In the context of this paper, which sheds light on capital inflows, market-based expanded finance is closely linked to portfolio investments. These empirical findings imply that an increase in the share of portfolio investment might be a contributory factor in the adverse distributional effects of capital inflows in AE.

In contrast, an increase in FDI can lead to a reduction in income inequality under certain conditions, although the evidence on the distributional effects of FDI is

[24]Harms et al. (2024) argue that the observed high correlation between two variables may be attributed to the existence of entry barriers in non-traded goods industries. However, they do not assert a causal relationship between the variables.

mixed (Eichengreen et al., 2021). In particular, in the case of greenfield investment in labor-abundant countries, as distinct from the mergers and acquisitions (M&A), FDI accelerates wage growth for low-skilled workers in labor-intensive sectors (Cornia, 2011) and increases the return to labor relative to capital with greater total investments (Amighini et al., 2017).

However, it is challenging to identify the effect of each subtype of capital inflows separately. While the share of a country's total external liabilities varies across subtypes, the *flows* of these three types on a yearly basis are highly correlated with each other. In order to ensure the orthogonality of shocks across subtypes of capital inflows and to derive reliable estimates of their effects, it is essential to include all three types in the model simultaneously and to identify exogenous shocks to each type separately. The zero and sign restrictions employed in this paper, however, are not sufficient to identify exogenous shocks to these three types of inflows. This is acknowledged as one of the limitations of this paper.

Another factor that could contribute to generating this differential is the degree of financial development. Existing studies argue that the financial development, which is generally measured by the domestic credit to private sectors as a percentage of nominal GDP, may condition the effects of financial integration (Kose et al., 2011).[25] However, there is conflicting empirical evidence on how the impact of financial integration on income inequality is contingent on financial development. Bumann and Lensink (2016) suggest that financial liberalization improves the income distribution in countries where financial development is high. Similarly, Furceri and Loungani (2018) show that the inequality-increasing effects of capital account liberalization are substantially lower in countries with more developed financial systems. In contrast, de Haan and Sturm (2017) find that financial liberalization worsens income inequality with higher financial development. Figure I.A.22 in the Appendix shows the level of financial development in sample countries along two dimensions: financial markets and financial institutions, using the Financial Development (FD) Index from the IMF. With a few exceptions, countries in AE generally have a higher level of financial development than countries in EME. In this regard, the results of this paper align more closely with the findings of de Haan and Sturm (2017).[26]

The following underpinning mechanism can produce these results: In EME, where financial markets are in a relatively underdeveloped state, increased capi-

[25] While there is a certain degree of interconnection between financial development and financial integration (including expanded cross-border capital flows), these two distinct concepts should not be considered as analogs.

[26] It is important to note that the aforementioned literature examines the effects of capital account liberalization on income inequality using the de jure indicators rather than the de facto indicators.

tal inflows benefit low-income households through the easing of funding conditions, improved access to credit, and more efficient allocation of financial resources. It may, in turn, contribute to a reduction in income inequality. On the contrary, as financial development deepens, typically in AE, more foreign capital favors rich households, already well-positioned in the financial markets, by creating more complex financial instruments or services. It may exacerbate income inequality. In this regard, Jaumotte et al. (2013) point out that "[i]f financial flows make resources available to a broader cross-section of the work-force, they would serve to reduce inequality by allowing investment in skills and human capital. However, if they make more financial resources available to those who already have capital and collateral, this would likely exacerbate inequality" (Jaumotte et al., 2013, p.282).

To sum up, the findings in this paper imply that the relationship between capital inflows and income inequality may not be straightforward and could change over time as the economy grows and/or the financial sector develops. While it is beyond the scope of the present paper to complete these remarkable differences in the distributional effects of capital inflow between AE and EME, the findings could provide important directions for future research.

## I.5 Conclusion

In this paper, I examine the effects of capital inflows on income inequality within countries using a panel structural VAR approach. To address the potential endogeneity of capital inflows, I identify exogenous capital inflow shocks employing sign restrictions. Furthermore, I estimate the distributional effects of capital inflow shocks for AE and EME separately, thereby yielding more informative results.

I find that capital inflow shocks worsen income inequality in AE while contributing to a reduction in income inequality in EME. An exogenous shock in capital inflows increases the year-over-year change rate of the Gini coefficient in AE by 0.1 percentage points two years after the shock. By contrast, the annual change rate of the Gini coefficient in EME declines by 0.06 percentage points one year after the shock. Capital inflow shocks account for about 13.6% of the forecast error variance in the Gini coefficient in AE and about 10.1% in EME at a 10-year forecast horizon, respectively. These results are statistically and economically significant. I also conduct further analysis by disaggregating these effects by income class, which supports the main results. The shocks to capital inflows are primarily associated with an increase in the income share of the high-income class in AE and the low-income class in EME.

These empirical findings have two main implications. Firstly, the possibility of adverse distributional effects of capital inflows provides another rationale for the proper management and control policies of cross-border capital movements. Capital flow management policies have primarily focused on macroeconomic instability and the likelihood of financial crises caused by a surge in capital inflows. Nevertheless, the findings of this paper suggest that policymakers should consider the potential adverse impacts of capital inflows on income inequality. In the recent policy paper, IMF (2022) has proposed the inclusion of 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. It is also worth noting that this paper 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" (IMF, 2022, p.8).

Secondly, the distributional effects of capital inflows vary depending on the economic conditions in the recipient countries. Although we have not yet been able to fully picture the distributional effects of capital inflows, the results of this paper suggest that country heterogeneity may play an important role in the distributional effects of capital inflows. These findings also underscore the necessity of tailored policy frameworks: AE may require redistributive policy measures to offset inequality induced by capital inflows, whereas EME must strengthen industrial policies that convert capital inflows into broad labor demand and inclusive growth. Future research using micro-level data, such as firm-level and household survey data, could enhance our understanding of the distributional implications of capital inflows.

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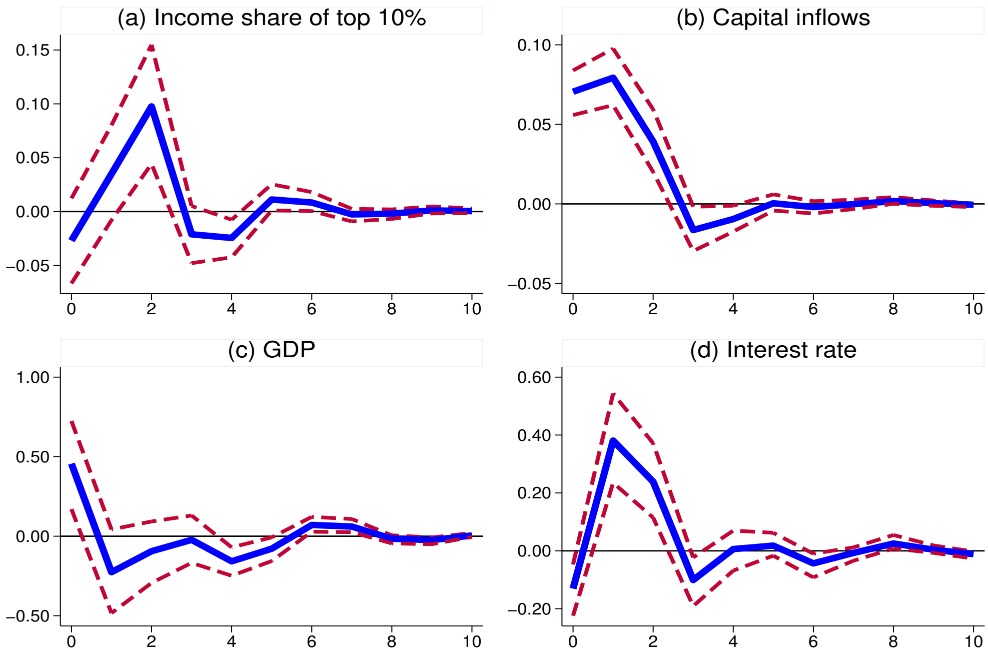
# I.A Appendix

Table I.A.1: Countries in the sample

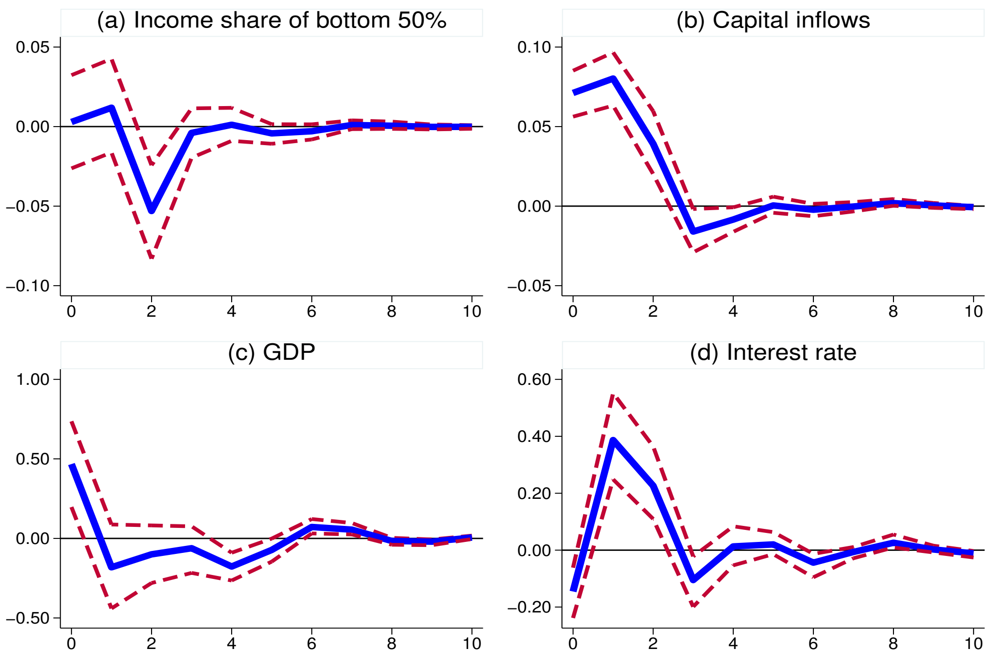
| | | |
|---|---|---|
| Advanced Economies | | |
| Australia | Austria | Belgium |
| Canada | Czech Republic | Denmark |
| Finland | France | Greece |
| Ireland | Israel | Italy |
| Korea | Latvia | Lithuania |
| Netherlands | New Zealand | Norway |
| Portugal | Spain | Slovak Republic |
| Slovenia | 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 I.A.1: Impulse responses to an exogenous capital inflow shock in AE (2)



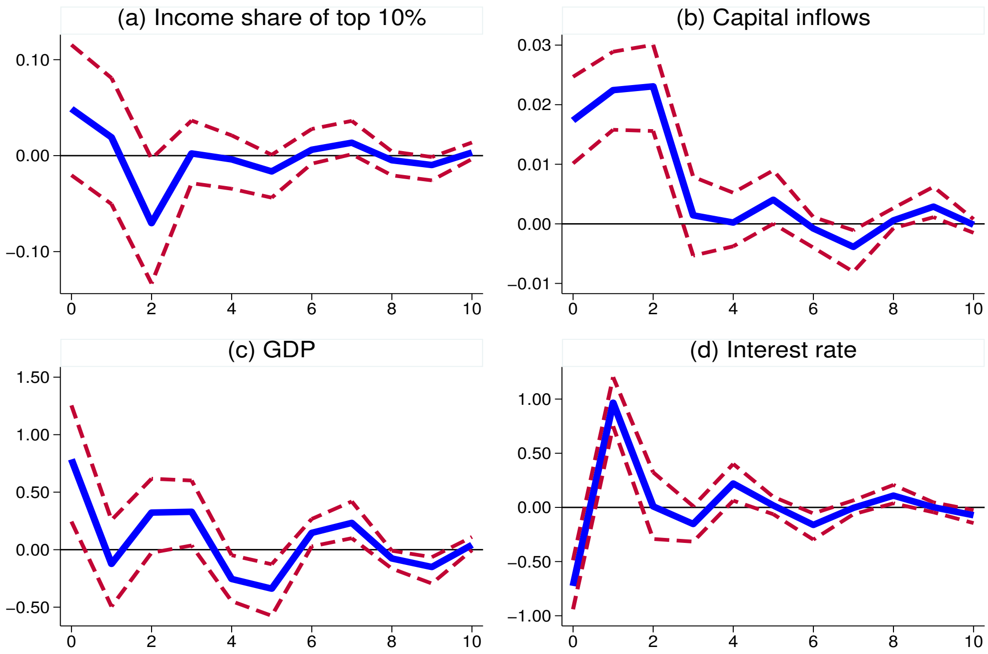

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.2: Impulse responses to an exogenous capital inflow shock in AE (3)



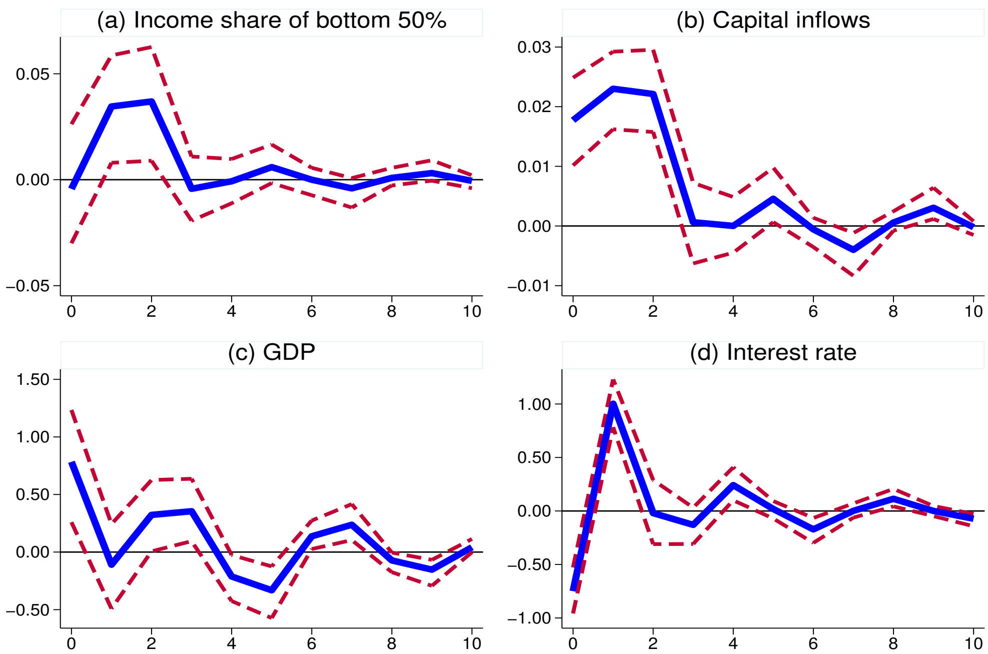

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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



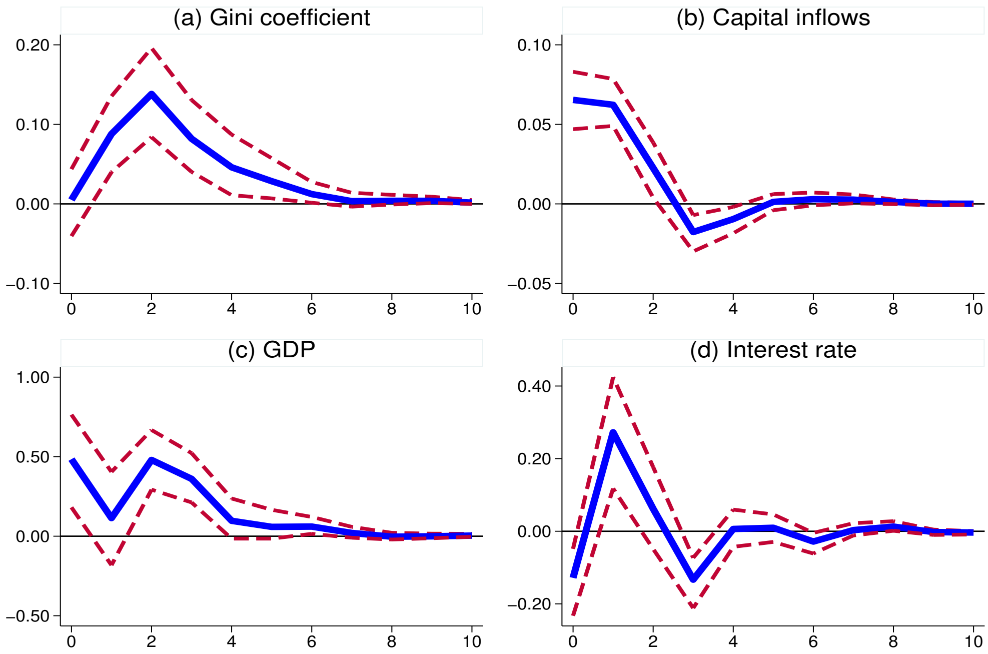

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.4: Impulse responses to an exogenous capital inflow shock in EME (3)



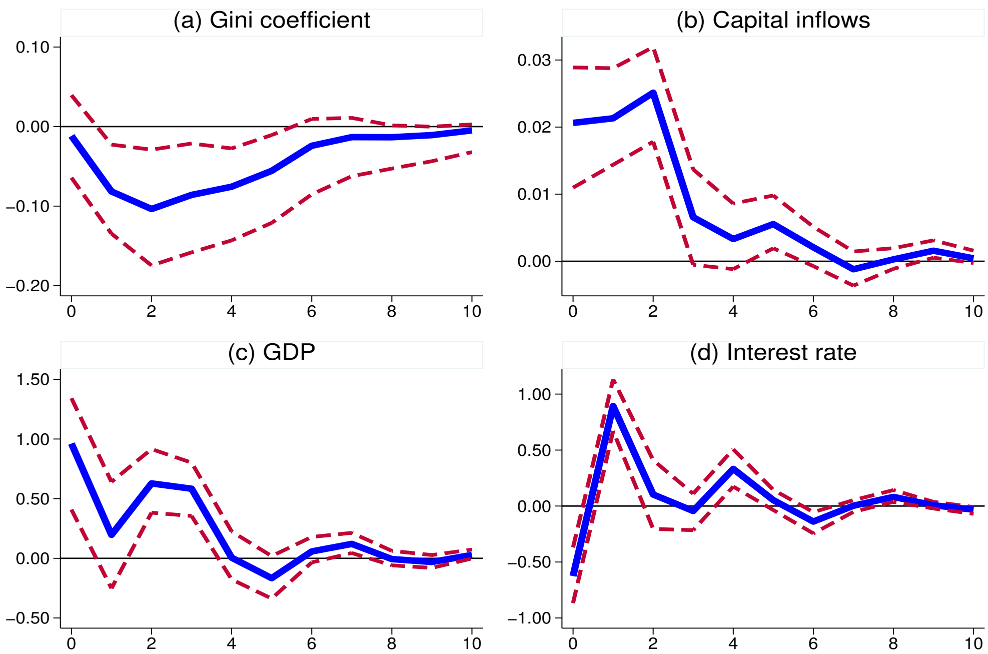

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.5: Robustness check (1.1): disposal Gini in AE



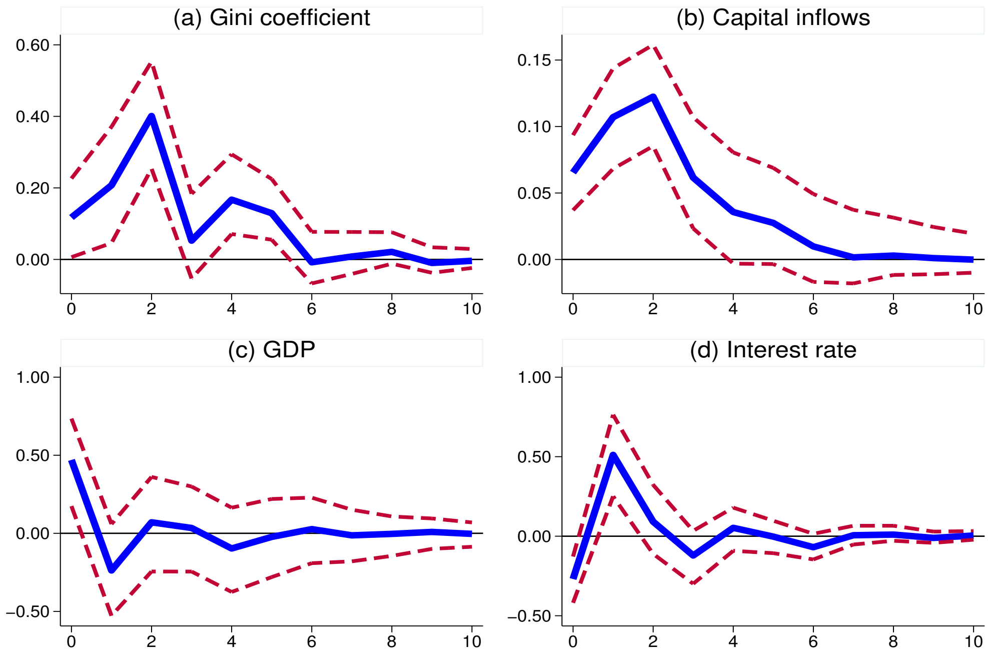

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.6: Robustness check (1.1): disposal Gini in EME



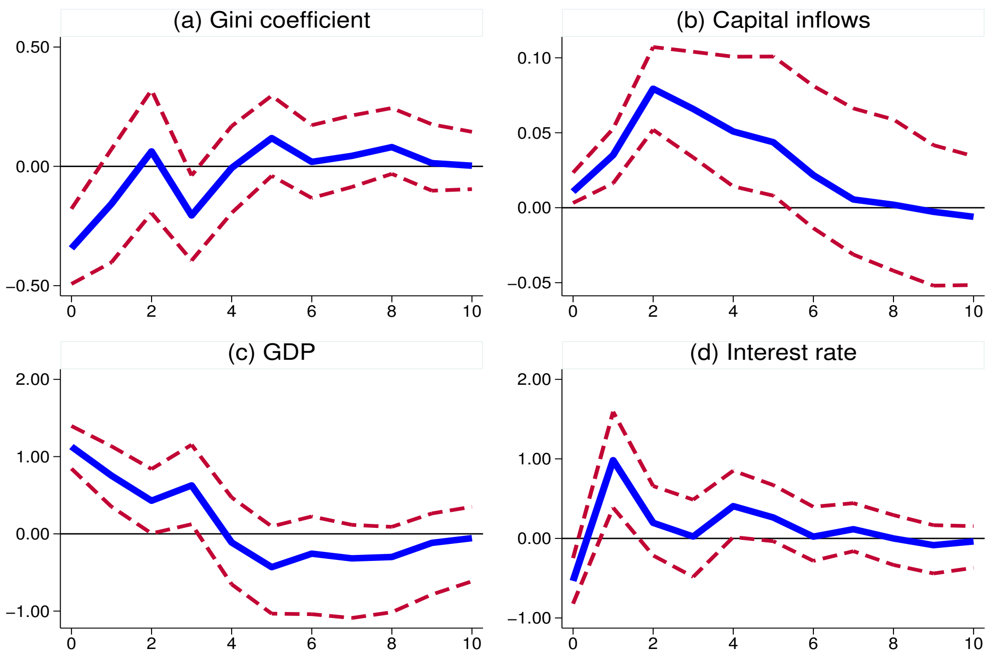

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.7: Robustness check (1.2): Gini from EHII in AE



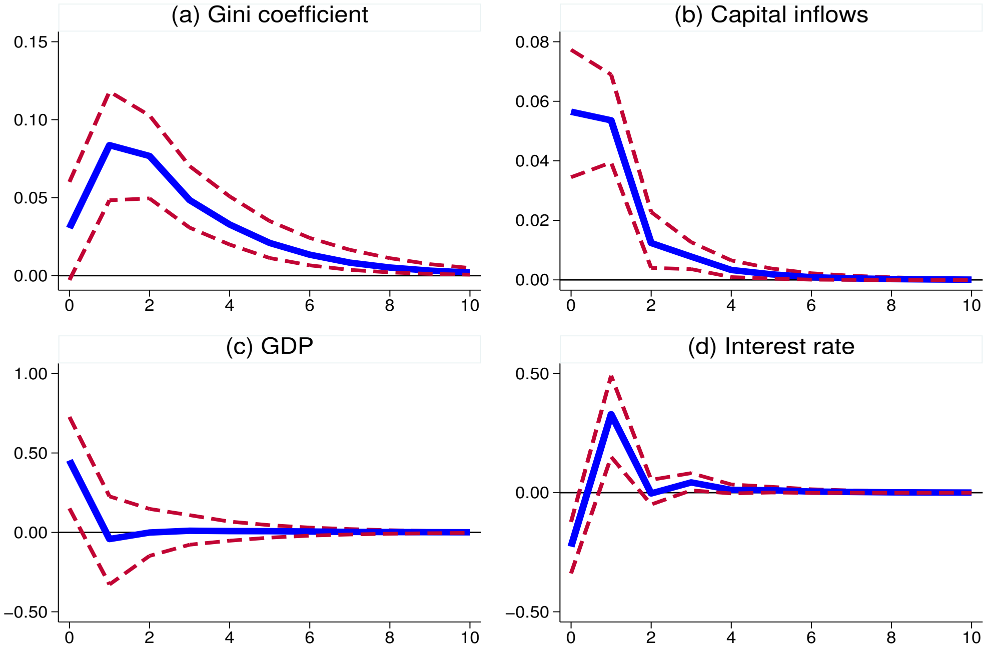

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.8: Robustness check (1.2): Gini from EHII in EME



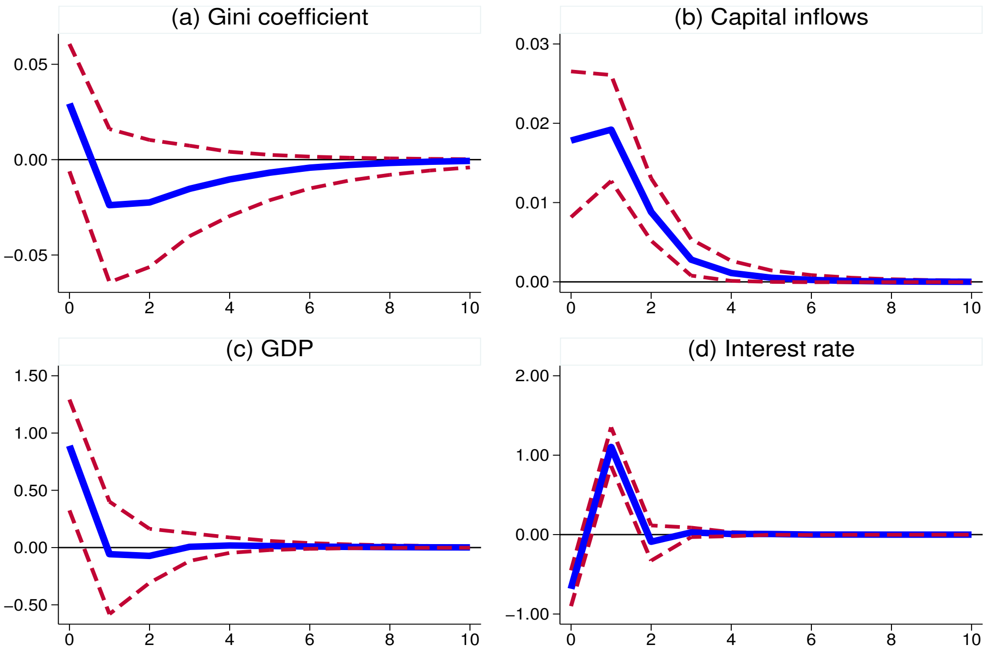

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.9: Robustness check (2.1): lag one in AE



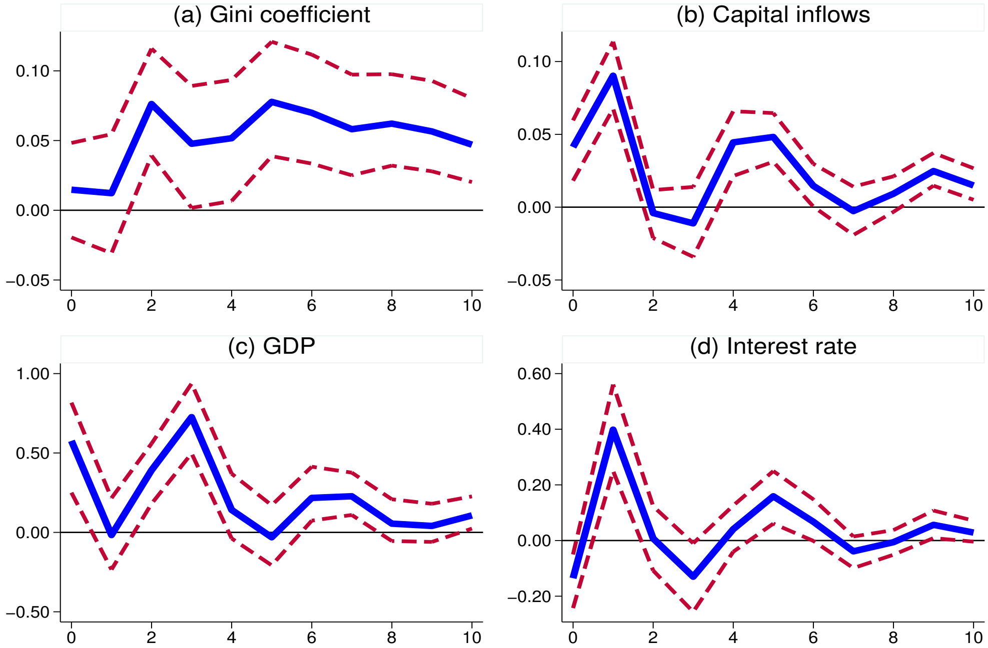

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.10: Robustness check (2.1): lag one in EME



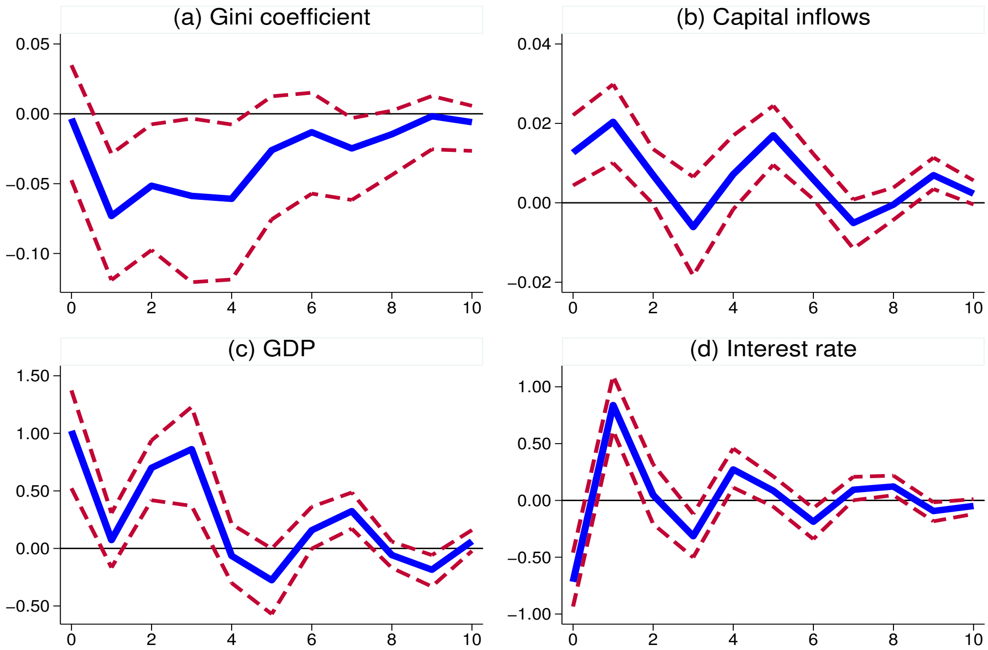

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.11: Robustness check (2.2): three lags in AE



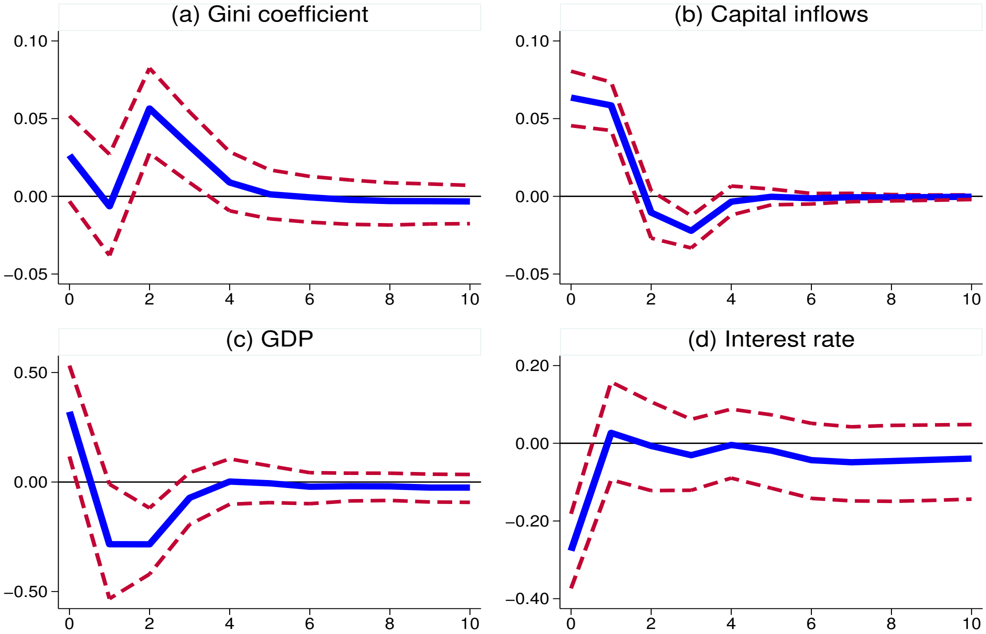

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.12: Robustness check (2.2): three lags in EME



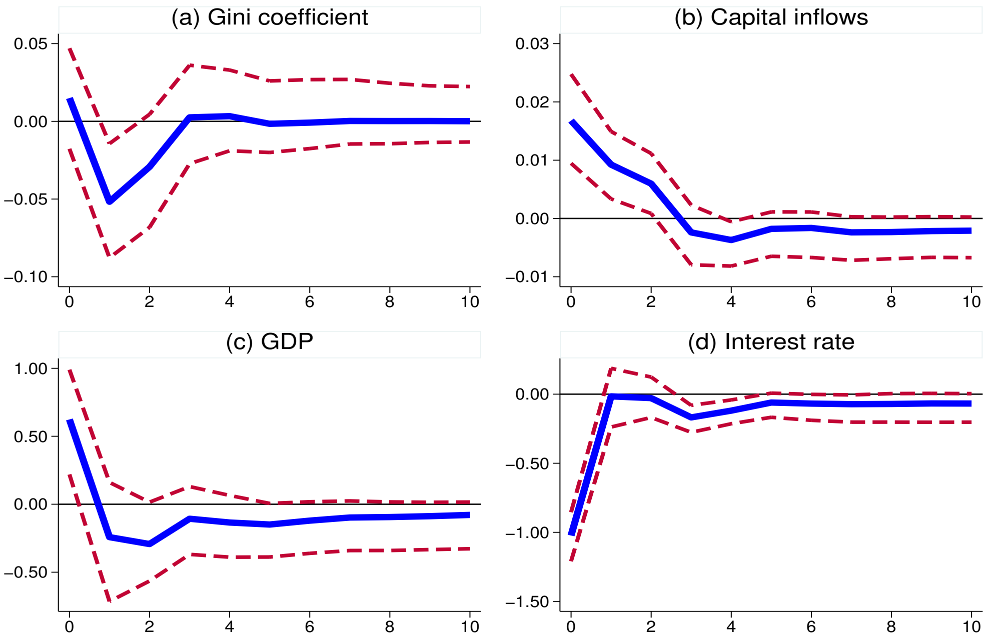

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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



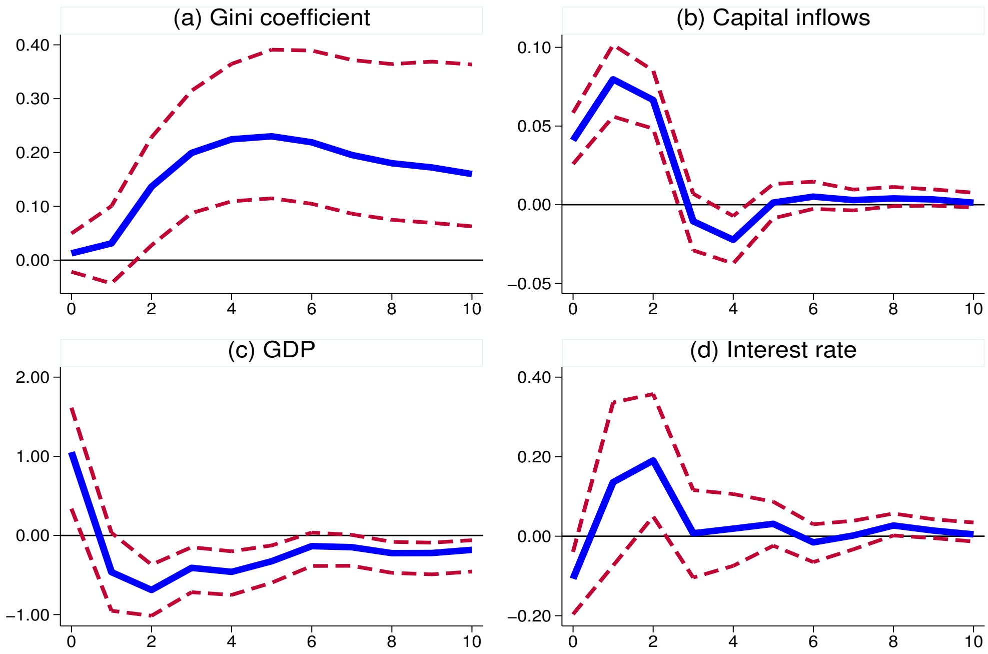

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.14: Robustness check (3.1): using long-term interest rates in levels in EME



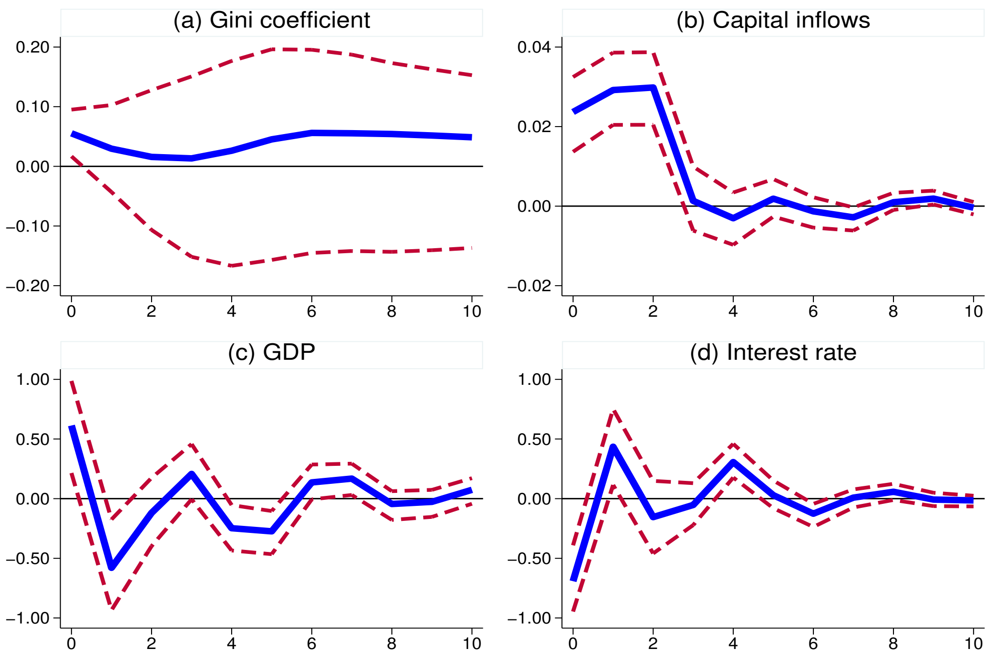

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.15: Robustness check (3.2): using the Gini in level in AE



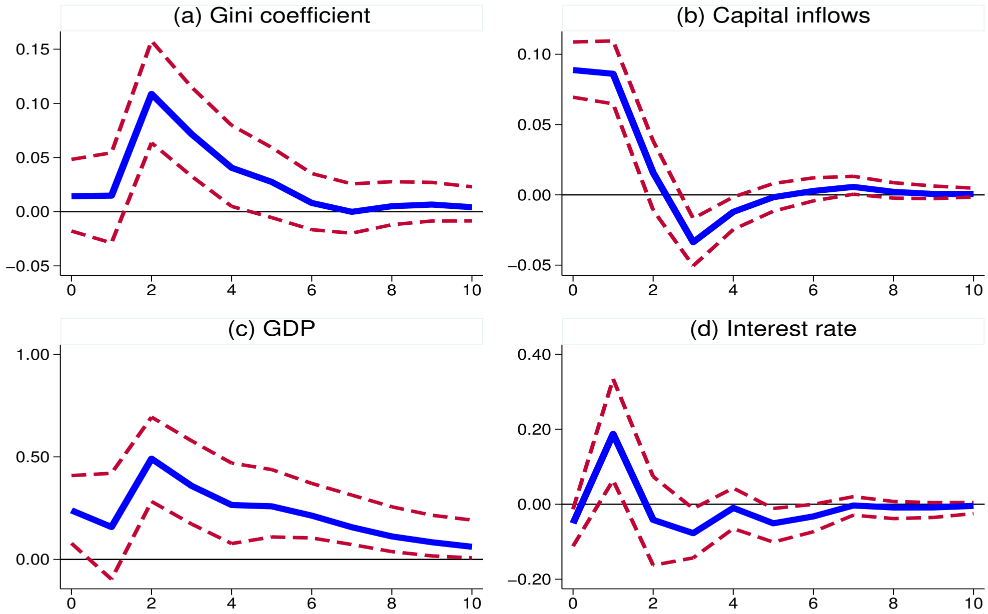

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.16: Robustness check (3.2): using the Gini in level in EME



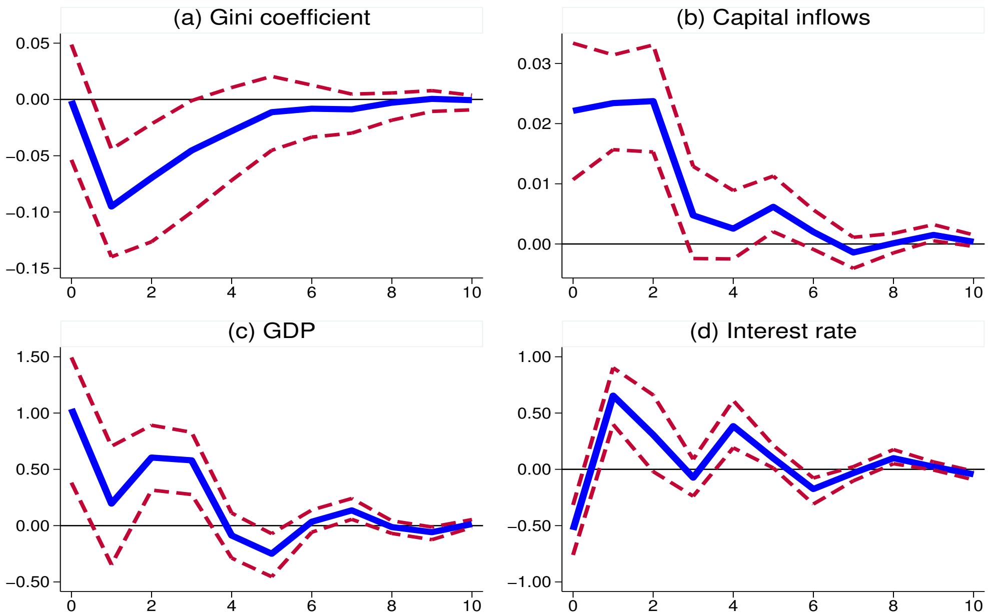

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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



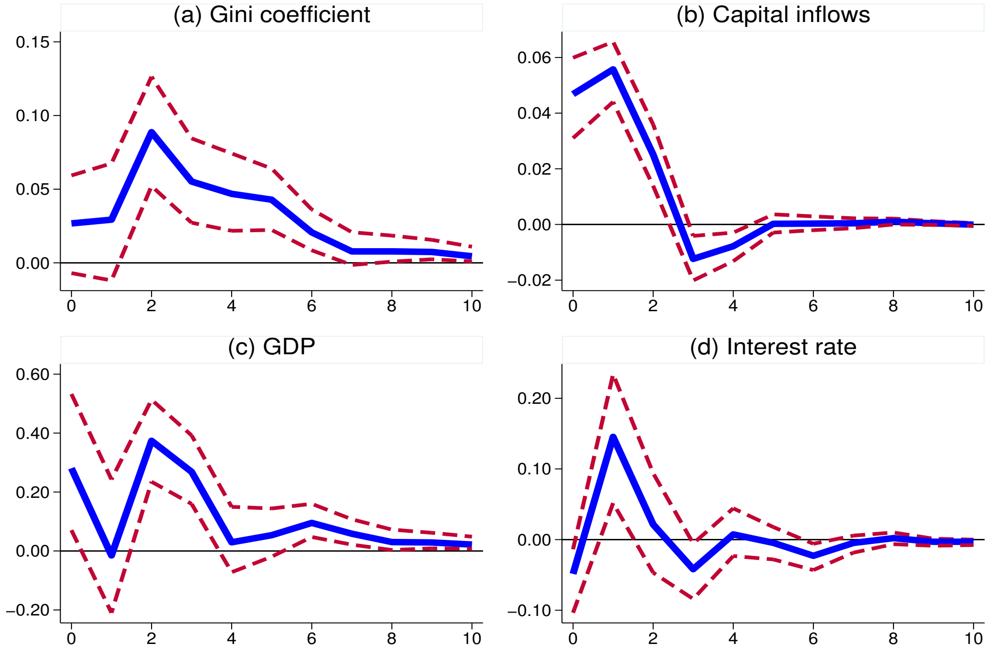

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.18: Robustness check (4): excluding some countries from the sample in EME



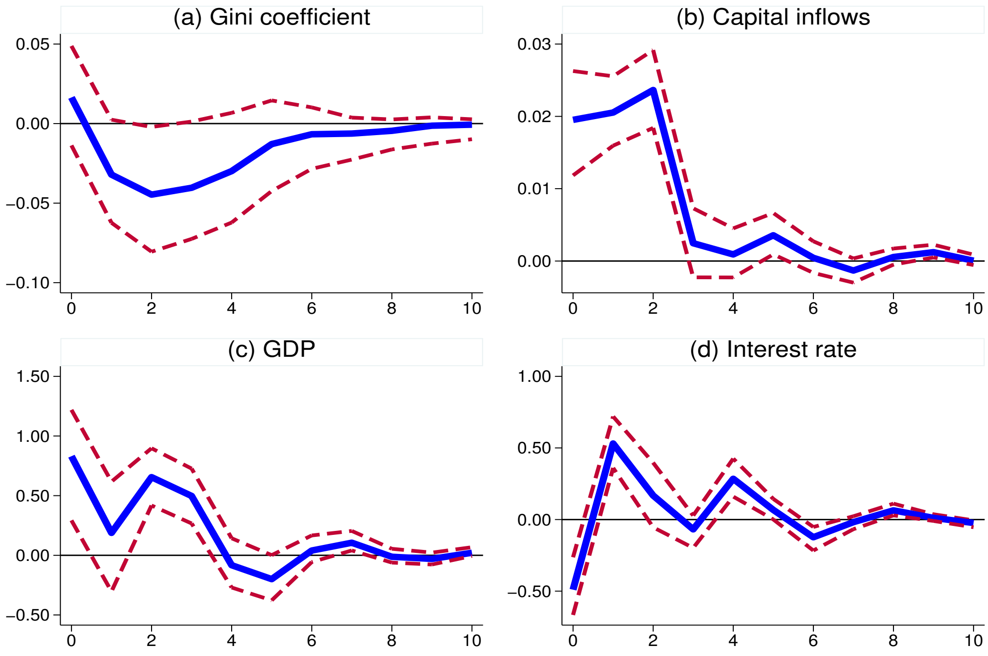

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

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




*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

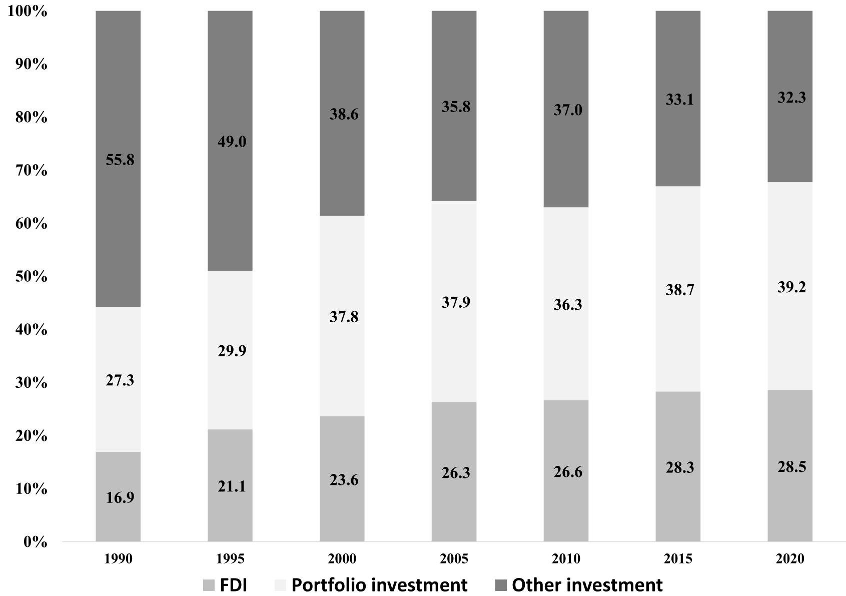
Figure I.A.20: Robustness check (5): addressing the impacts of outliers in EME



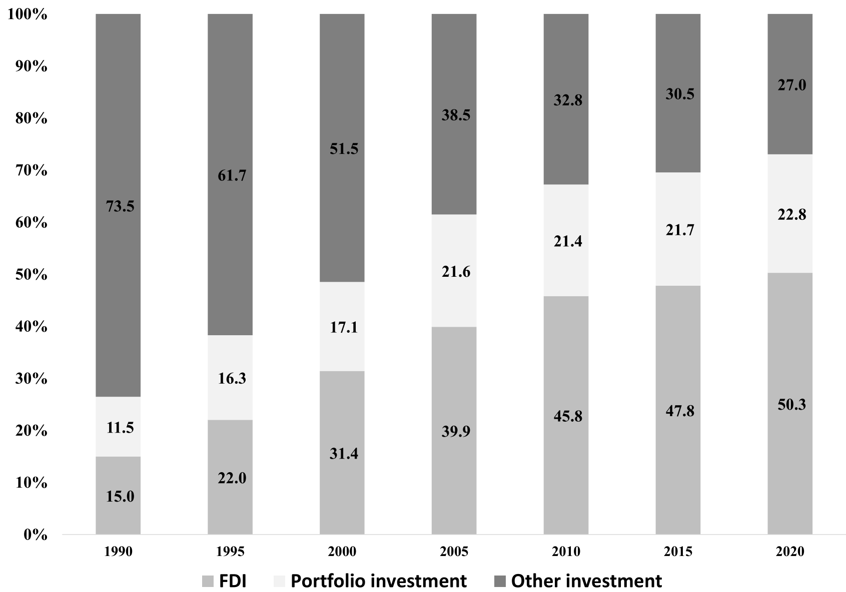

*Note*: The solid blue line represents the median response and the dashed red lines are the 16th and 84th percentiles of 1,000 independent draws that satisfy sign restrictions.

Figure I.A.21: Composition of external liabilities, 1990-2020

(a) AE

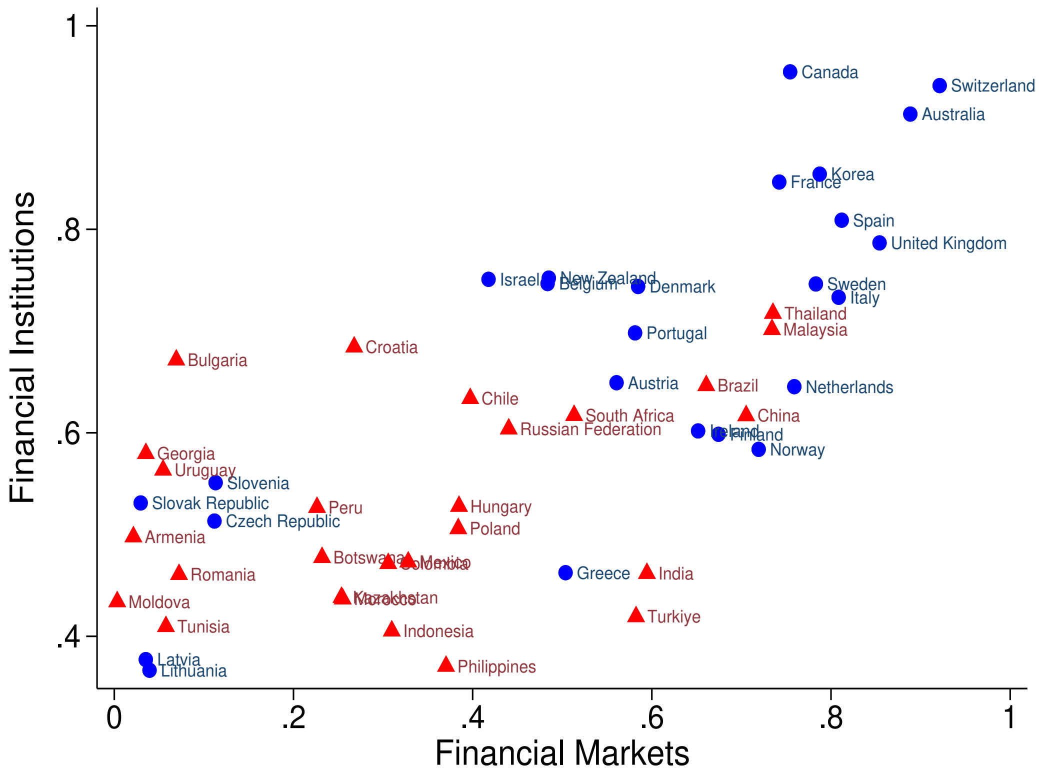



(b) EME




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

Figure I.A.22: The FD index by country




*Note*: The blue circles represent the countries in AE, and the red triangles represent those in EME. The FD index has a broader coverage and provides nine indices, the multi-dimensional measures for financial market and institution development in terms of depth, accessibility, and efficiency. For more details on the index, see Svirydzenka (2016). *Source*: IMF

# II. Financial Integration and Wage Inequality: Evidence from European Countries

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# Financial Integration and Wage Inequality: Evidence from European Countries

Jinyeong Yun†

**Abstract**

In this paper, I examine the impact of financial integration on wage inequality using an unbalanced panel for 20 European countries over the period 1999 to 2021. Focusing on between-firm wage inequality arising from firm heterogeneity, I find that financial integration, measured by the sum of external assets and liabilities, increases wage inequality within industries. The effects are more pronounced in the medium term rather than in the short term. Furthermore, the direction of financial integration, i.e., whether it involves an increase in external assets or liabilities, matters: inward financial integration (or capital inflows) significantly increases wage inequality, while outward (or capital outflows) does not. I also provide empirical evidence on the channel of distributional effects of financial integration. The financial integration shocks widen the capital intensity (the capital-to-labor ratio) gap between firms within industries, which in turn widens the labor productivity gap. Through this channel, financial integration may boost wages in large firms more than in small and medium-sized enterprises (SMEs). Additionally, I show that the impact of financial integration on wage inequality depends on the degree of external financial dependence of an industry. The level of financial development may mitigate the adverse impact of financial integration on wage inequality. However, this effect is less definitive.

**Keywords**: Financial integration; wage inequality; firm heterogeneity
***JEL* codes**: D63, F36, F38, J31

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## II.1 Introduction

Increased income inequality within countries has emerged as a major concern of our time, as income inequality has risen in many developed and developing countries over the past few decades. Given that wages account for the largest share of household income, it is critical to investigate the drivers of wage inequality for a better understanding of the cause of overall income inequality.[1] The widening wage inequality can significantly reduce the purchasing power of a large portion of households, undermining aggregate consumption and economic growth. Rising overall income inequality and wage inequality can also fuel many social problems, such as social unrest, distrust of government institutions, and political polarization. Consequently, it is now crucial for policymakers and academic researchers to understand the factors that contribute to rising wage inequality.

A growing body of literature has recently underscored the role of international financial integration as a key driving force of rising income inequality. Several empirical studies have argued that financial liberalization (or capital account opening) is positively associated with increased income inequality (Jaumotte et al., 2013; Asteriou et al., 2014; de Haan & Sturm, 2017; Furceri & Loungani, 2018; Erauskin & Turnovsky, 2019; Furceri et al., 2019; Li & Su, 2021; Erauskin & Turnovsky, 2022). Similarly, some research has also found empirical evidence on the adverse distributional effects of cross-border capital flows (Liu et al., 2023; Yun, 2025). However, there are only a few studies in the literature that address the impact of international financial integration on wage inequality rather than overall income inequality.

The existing literature on wage inequality has mainly focused on the widening skill premium between high-skilled and low-skilled workers, which leads to greater wage inequality. On the other hand, since the seminal work of Abowd et al. (1999), which introduced a framework to decompose wage dispersion, there has been a growing body of literature emphasizing the role of firms in explaining wage inequality (Akerman et al., 2013; Card et al., 2013; Alvarez et al., 2018; Song et al., 2019; Criscuolo et al., 2020). These studies suggest that firm heterogeneity contributes to a significant portion of rising wage inequality, meaning that some firms pay employees with similar skill levels higher wages than others. The literature defines and analyzes this as "between-firm wage inequality", distinct from "within-firm wage inequality",

[1] Overall income inequality can be more deeply analyzed by decomposing it into different parts: changes in the distribution of production factors (roughly capital, labor, and land) among households, changes in the remuneration of production factors, and changes in transfers received and taxes paid by households (Cornia, 2011). Of these, changes in the remuneration of production factors can be further decomposed into changes in the relative remuneration of production factors and inequality in the remuneration of the same production factor, including wage inequality.

which reflects differences in workers' characteristics such as education level, work experience, and gender. However, theoretical and empirical answers to what drives between-firm wage inequality remain controversial.

In this paper, I contribute to the debate on the distributional effects of international financial integration, by bridging these two streams of literature. I empirically examine the impact of financial integration on the between-firm wage inequality in European countries over the period 1999 to 2021, using the Competitiveness Research Network (CompNet) dataset. To shed light on the medium to long-term effects of financial integration on between-firm wage inequality, I estimate the impulse response functions of wage inequality to financial integration shocks using local projections proposed by Jordà (2005) with high-dimensional panel regressions. I show that financial integration, measured by the sum of total external assets and liabilities, increases wage inequality within industries in the medium term. Inward financial integration has a more pronounced distributional effect than outward financial integration. Furthermore, I investigate the mechanisms of how financial integration affects between-firm wage inequality. I find that financial integration shocks widen the capital intensity (the capital-to-labor ratio) gap between firms within industries, which in turn widens the labor productivity gap. Through this channel, financial integration may boost wages in large firms more than in small and medium-sized enterprises (SMEs).

The remainder of this paper is structured as follows: Section II.2 provides a brief overview of the relevant existing literature. Section II.3 then describes the data, variable construction, and empirical approach used in the present paper. Next, Section II.4 presents and discusses empirical results. Section II.5 summarizes the main findings and concludes.

## II.2 Conceptual framework and related literature

### II.2.1 Between-firm wage inequality

The classical theory of wage determination suggests that the market wage is determined by the equilibrium between the labor supply of households and the labor demand of firms, which depends on the marginal product of labor. In a competitive labor market, each firm takes this market wage as a given, and firm heterogeneity plays no role in how much firms pay their workers. Based on this basic view of macroeconomics and labor economics, a large strand of existing literature studying wage inequality has long focused on the role of differences in workers' skills, which reflect their education level and work experience.

On the other hand, a growing body of literature has recently argued that firm heterogeneity plays a significant role in wage determination and wage dispersion. Since the seminal work of Abowd et al. (1999), many studies have analyzed wage inequality by separately identifying "within-firm wage inequality" and "between-firm wage inequality". Within-firm inequality refers to dispersion in wages between workers within firms. Between-firm inequality refers to the dispersion in average wages across firms, which mainly arises from firm heterogeneity. With the growing availability of administrative matched employer-employee data, researchers have found that between-firm wage inequality accounts for a substantial proportion of overall wage inequality. Using Swedish manufacturing data, Akerman et al. (2013) document that much of the wage inequality exists within sectors and occupations across workers with similar characteristics. Card et al. (2013) study the contribution of firm heterogeneity in West German wage inequality using administrative data from 1985 to 2009. They argue that increasing firm heterogeneity explains a large share of the rise in wage inequality. Similarly, Alvarez et al. (2018) document that firm heterogeneity accounts for 40 percent of the total wage inequality decrease in Brazil. With a matched employer-employee database for the United States from 1978 to 2013, Song et al. (2019) find that two-thirds of the rise in wage inequality is due to a rise in the dispersion of average earnings between firms. They also suggest that rising between-firm variance is accounted for by a widening gap between firms in the composition of their workers. Criscuolo et al. (2020) show that changes in the dispersion of average wage between firms account for about half of the changes in overall wage inequality, using a linked employer-employee dataset for 14 OECD countries.

### II.2.2 Financial integration

To study the impacts of international financial integration (henceforth "financial integration"), we should first clarify how to define and measure it. The existing literature has used various terms, such as "financial integration", "financial globalization", and "capital account liberalization", to refer to the trend toward more closely coupled financial markets in neighboring countries, regions, and the global economy and increased cross-border capital flows. In this paper, I use "financial integration" rather than any of these concepts.[2] The term "globalization" is often used to describe a phenomenon in many different aspects: economic, political, social, and cultural. "Capital account liberalization" or "Capital account openness"

[2]However, all these concepts are closely related and are sometimes used synonymously in the relevant literature (Gräbner et al., 2021).

is mainly used to refer to a specific point in time associated with the capital account opening or financial deregulations. Most studies use the concept of "capital account liberalization" to focus on changes in economic variables before and after the episode of capital account liberalization (Larrain, 2015; Bumann & Lensink, 2016; de Haan & Sturm, 2017; Furceri & Loungani, 2018; Furceri et al., 2019; Li & Su, 2021). However, the remarkable phenomenon of financial market integration should be understood as a continuous process that accumulates over time rather than a discontinuous process that occurs before and after a specific event at a particular time. The removal of domestic and international regulations restricting capital flows, the reform of laws and institutions, the opening of bond markets, and the opening of equity markets are not simultaneous but rather a series of events. From this perspective, we can adequately capture the entire evolution of this phenomenon by using the term "financial integration" to refer to the closer interconnection of the financial markets.

Existing measures of financial integration can be categorized into two groups: "*de-jure*" and "*de-facto*" indicators.[3] *De-jure* measures are based on a country's institutional and regulatory environment for cross-border capital transactions. On the other hand, *de-facto* measures are outcome-oriented indicators that reflect a country's actual position in global financial markets. As discussed later in Section II.3.1, I use *de-facto*, not *de-jure*, indicators in the analysis to measure financial integration.

### II.2.3 Link between financial integration and wage inequality

Several studies have emphasized the importance of firm heterogeneity for wage inequality, but the causes of rising wage inequality and firm heterogeneity remain unresolved and controversial. In this paper, I aim to study the role of financial integration as one of the drivers for those. The research question of this paper is whether financial integration leads to an increase in between-firm wage inequality. More precisely, this study hypothesizes that financial integration increases wage inequality between large businesses and small and medium enterprises (SMEs). This argument implies that the wages of big firms, which initially pay relatively high wages, may grow faster with financial integration than those of SMEs.

To varying degrees across businesses, firms rely on external capital to finance fixed costs that cannot be covered by retained earnings or internal cash flow from business operations. These fixed costs typically relate to investment in fixed capital equipment, research and development (R&D) expenditures, and marketing and

[3]For a more detailed overview of two types of measures, see Gräbner et al. (2021)

advertising expenses. To meet the needs for these expenditures, firms raise funds from banks and other financial institutions or take credit directly from domestic and foreign investors. In addition, most variable costs, such as the purchases of intermediate input, wages paid to workers, and land or equipment leases, are incurred before sales revenue is realized. Therefore, firms need to finance some of these upfront variable costs through external financing.

Capital account opening and the resulting progress in financial integration (increased cross-border capital flows) provide more opportunities for firms to obtain external financing for operations and investments. These chances are critical for expanding production capacity or improving product quality. The additional external funds allow firms to grow, and firms can pay higher wages to their employees. However, given some market imperfections, increased external finance may be unevenly distributed among firms. Under the constraints of financial markets, increased capital inflows provide more finance to large businesses already paying higher wages to their workers. Big firms publicly traded in financial markets have easier access to external finance than SMEs, while many SMEs are financially constrained. Moreover, global investors tend to invest in big firms because information asymmetries in financial markets prevent global lenders and investors from accurately valuing firms.

As previously noted, only a few studies have examined the relationship between financial integration and wage inequality. Using aggregate and sectoral data in 20 developed countries from 1975 to 2005, Larrain (2015) shows that opening the capital account increases wage inequality, with a focus on inequality between skilled and unskilled workers. Eklou and Foster (2023) reach a similar conclusion that capital account liberalization increases wage inequality using firm-level data in ASEAN5 countries over the period 1995-2019.

## II.3 Empirical methodology

### II.3.1 Data

***Wage inequality*** In order to measure between-firm wage inequality, I use data from the CompNet dataset, which compiles data obtained from administrative and public sources. This dataset provides a variety of indicators computed at the firm level, covering non-financial corporations with at least one employee in European countries. Indicators of the CompNet dataset are aggregated to different levels according to sector definitions or firm characteristics, such as the location of firms. I employ the industry-country-year level of aggregation among these aggregation levels, where industries are classified corresponding to the 2-digit NACE Revision

2 industries. The most recent version of the CompNet dataset, the $9th$ vintage, is an unbalanced panel dataset for 22 European countries from 1999 to 2021.[4] The analysis in this paper focuses on 20 of these 22 countries, excluding Malta and the United Kingdom.[5] It is worth mentioning that all variables in the CompNet dataset are available for two samples, the "all" sample and the "20e" sample, which includes only firms with 20 or more employees. I use the latter to improve the coverage of the samples and comparability between countries.[6]

To measure between-firm wage inequality within industries, I construct a variable using information on the distribution of firms' average wage (labor costs per employee). The CompNet dataset provides statistics of percentiles for all indicators included in the dataset, which allows for computing the dispersion in the firms' average wages. Wage inequality is measured by calculating the ratio of the $90th$ percentile of firms' average wage to the $10th$ percentile (90-10 ratio) and the $50th$ percentile (90-50 ratio).[7]

Table II.1 reports summary statistics for between-firm wage inequality.[8] Wage inequality measured by the 90-10 and 90–50 ratios is, on average, 2.92 and 1.65, respectively, which means that the average wage of the 90th percentile firm is 192% and 65% higher than the average wage of the 10th percentile firm and the median firm in each industry, respectively. As shown in the table, the distribution of wage inequality is long-tailed to the right, indicating the presence of outliers.[9]

[4]For most countries, however, data are only available up to 2020.

[5]Malta's undersized economy, small sample size for individual industries, and enormous external assets and liabilities significantly limit comparability with other countries. For the UK, data at the 2-digit industry level are not available.

[6]The main reason for having two samples is that in some countries, firms are legally obliged to report their balance sheet data only when certain size thresholds are met. For more detail, see the User Guide for the 9th Vintage of the CompNet Dataset, available at `https://www.comp-net.org/fileadmin/_compnet/user_upload/9th_Vintage_User_Guide_final_102023.pdf`

[7]While it can be assumed that workers at the same firm in the same industry have similar characteristics, it is worth noting that these measures of wage inequality do not control for heterogeneity in worker characteristics. Therefore, the measures computed from the data may capture not only between-firm wage inequality but also, to some extent, wage differentials due to worker heterogeneity.

[8]Observations from industries with small sample sizes (less than 20) are excluded because changes in a specific firm may dictate the overall wage inequality within an industry.

[9]To address potential bias due to exceptionally high values of wage inequality, I also conduct a robustness check by excluding some outliers.

Table II.1: Descriptive statistics: between-firm wage inequality

| | Mean | Median | SD | Min | Max | N |
|---|---|---|---|---|---|---|
| 90-10 ratio | 2.92 | 2.45 | 1.76 | 1.19 | 34.93 | 16,243 |
| 90-50 ratio | 1.65 | 1.54 | 0.47 | 1.05 | 17.78 | 16,243 |

*Source*: CompNet dataset and author's calculation

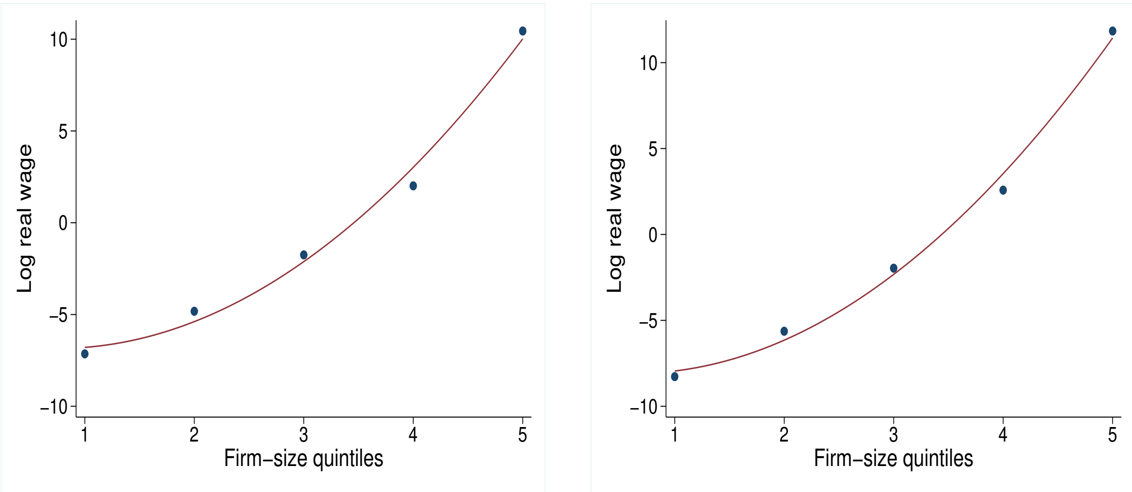
Figure II.A.1 in the Appendix shows the average 90-10 ratio for each country. Wage inequality in Central and Eastern European countries tends to be higher than in Western European countries. In Hungary, Latvia, and Romania, the 90th percentile firms pay their workers about four times higher than the average wage of the 10th percentile firms, while the average 90–10 ratios in Belgium, Finland, and Sweden are about two.

While firms operating within the same industry share similar characteristics, there is substantial heterogeneity among them. Many studies show that firms within industries are heterogeneous in terms of their economic properties, such as productivity, capital intensity, R&D expenditures, and markups. In this paper, I assume that differences in firm size can serve as a representative proxy for these various firm heterogeneities. Figure II.1 presents the difference in average real wages (and median of real wages) across quintiles of firm size within industries, calculated using the full sample over the entire sample period. I adjust real wages by estimating the residuals of a panel regression using country-, year-, and industry-fixed effects. As shown in the figure, real wages increase on average (exponentially rather than linearly) as firm size increases.

***Financial integration*** The degree of financial integration can be measured using the *de jure* and *de facto* indicators. The *de jure* indicator of financial integration most commonly used in the literature is the Capital Account Openness Index (KAOPEN Index) developed by Chinn and Ito (2006, 2008). However, using the KAOPEN Index has some limitations for the analysis in this paper. Based on the KAOPEN index, many sample countries have already fully opened their capital accounts at the start of the sample period. This means that the KAOPEN indices for these countries have the same value over the sample period. As an alternative, I use the *de facto* indicator, the sum of total external assets and liabilities. Data on total external assets and liabilities is taken from the External Wealth of Nations (EWN)

Figure II.1: Real wage by firm size

(a) Mean within each quintile (b) Median within each quintile




*Note*: Log real wage is the residual of a regression using country-, year-, and industry-fixed effects.
*Source*: CompNet dataset and author's calculation

database developed by Lane and Milesi-Ferretti (2007, 2018) and its update.[10] The database provides estimates of external financial assets and liabilities for 212 countries over the period 1970-2022, based on International Investment Position (IIP) statistics and Balance of Payments (BOP) statistics provided and maintained by the International Monetary Fund (IMF).

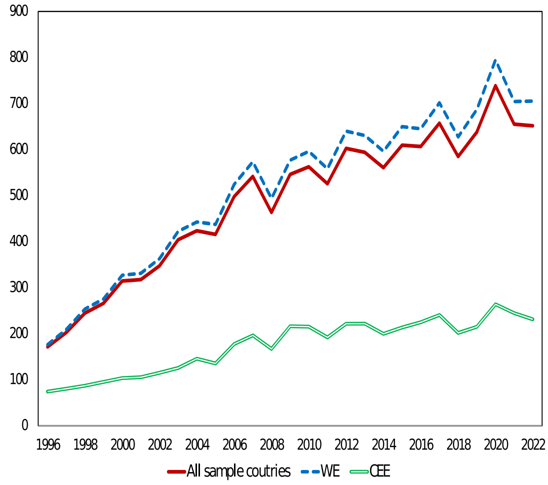
Figure II.2 shows the evolution of financial integration, scaled by a percentage of GDP, in European countries in recent decades. Financial integration has been steadily progressing during this period. While this trend has been driven mainly by Western European (WE) countries, Central and Eastern European (CEE) countries have also seen a steady increase in external assets and liabilities.

***Control variables*** To control for the distributional effects of fluctuations in the business cycle at the national level and the impacts of financial crises, I include the growth rates of real GDP and a dummy of financial crises as control variables. The data on GDP are taken from the World Development Indicator (WDI) database of the World Bank. The data on financial crises are obtained from a database developed by Nguyen et al. (2022), which extends the Systemic Banking Crises Database by Laeven and Valencia (2020). In addition, several control variables at the industry level are also taken from the CompNet dataset. To control the dispersion in firm size within industries, I include the standard deviation of the distribution of firm size by the number of employees as control variables. I also include the industry

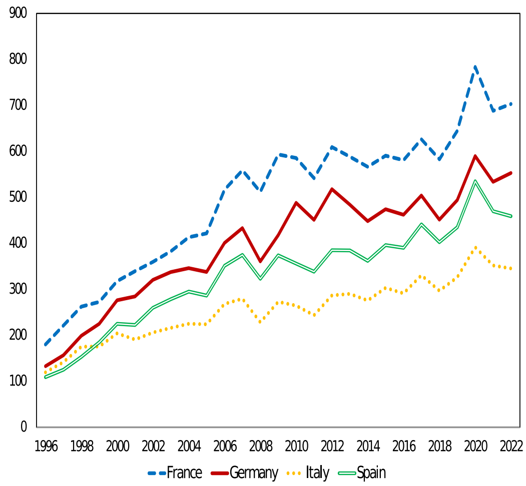
[10] I use the December 2023 version, available at `https://www.brookings.edu/articles/the-external-wealth-of-nations-database/`

Figure II.2: Financial integration

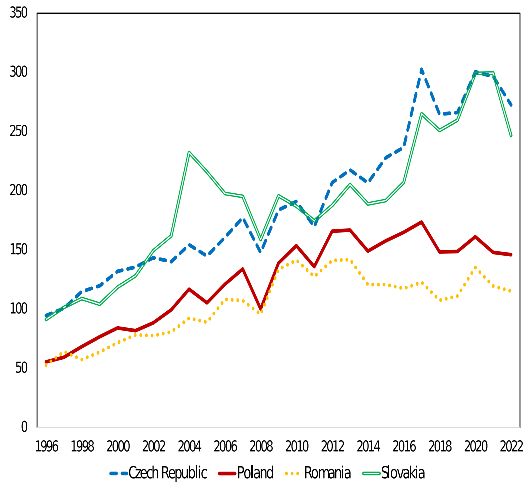
(a) All sample countries, WE, and CEE




(b) Selected countries in WE




(c) Selected countries in CEE




*Note*: The figure shows the unweighted averages of the financial integration indicator by country group. Financial integration is measured by the sum of total external assets and liabilities as a percentage of GDP. *Source*: EWN database and author's calculation

concentration of nominal revenue, as measured by the Hirschman-Herfindahl Index (HHI),[11] and the annual growth rate of nominal revenue.

### II.3.2 Model

To shed light on the contemporaneous as well as the medium to long-term responses of between-firm wage inequality, I estimate impulse response functions using the local projections proposed by Jordà (2005).

[11] Cortes and Tschopp (2024) find a positive correlation between concentration and between-firm wage inequality.

The following equation is estimated

$$\begin{aligned} log(Inequality)_{i,c,t+h} \;=\; & \alpha_h log(Inequality)_{i,c,t-1} \;+\; \beta_h \Delta FIN_{c,t} \\ & + \mathbf{\Gamma}_h \mathbf{X}_{i,c,t} \;+\; \tau_i^h \;+\; \gamma_c^h \;+\; \delta_t^h \;+\; \upsilon_{i,c,t+h} \end{aligned} \tag{II.1}$$

for each $h = 0, 1, .., H$, which denotes the time horizon for the local projections. The indices $i$, $c$, and $t$ represent industry, country, and year, respectively. $Inequality_{i,c,t}$ is the 90-10 ratio (and the 90-50 ratio), a dependent variable that measures wage inequality. To account for the persistence of wage inequality, I include one lag of the dependent variable on the right-hand side of the equation.[12]

$\Delta FIN_{c,t}$ is a financial integration shock that is measured by the first difference in the total external assets and liabilities between $t$ and $t-1$ as a percentage of nominal GDP at $t-1$. I assume that the shock of financial integration is sufficiently exogenous at the individual industry level because it is a country-level variable. The industries at 2-digit classification levels are small enough to support this assumption. Therefore, I do not use any additional restrictions to identify the shock. The parameter of interest is $\beta_h$, which represents the effects of a financial integration shock on wage inequality after $h$ periods. Combining this parameter as a function of time horizon $h$ provides the impulse response function of wage inequality to a financial integration shock. I also demean $\Delta Fin_{c,t}$ by each country and then normalize by its standard deviation to facilitate the interpretation of the magnitude of the impulse responses.[13] The estimated coefficients $\beta_h$, therefore, can be interpreted as a percentage change in between-firm wage inequality to a one-standard deviation increase in financial integration.

$\tau_i$ and $\gamma_c$ are industry- and country-fixed effects, which are included to control for unobserved cross-industry and cross-country heterogeneity in wage inequality, respectively.[14] Country-fixed effects can also mitigate concerns about the cross-country comparability of the CompNet dataset that may arise due to different data sources.[15] $\delta_t$ is time-fixed effects to control for global or regional shocks that affect

[12]The ordinary least squares (OLS) estimator may not provide efficient estimates for a dynamic panel model with individual fixed effects, such as the one in equation (II.1), especially in settings where $N$ is large and $T$ is small (Nickell, 1981). In a robustness check, I use the Generalized Method of Moments (GMM) estimator proposed by Arellano and Bond (1991) to address this possible bias.

[13]To calculate the mean and standard deviation, I use data for the entire period from 1999 to 2021, while the data period used for estimation varies by country.

[14]Failure to adequately control for unobserved heterogeneity can preclude causal inferences and produce inconsistent estimates due to omitted variable bias (Gormley & Matsa, 2014).

[15]Differences in data sampling methodology in each country can cause problems comparing

the entire European region. $\mathbf{X}_{i,c,t}$ is a vector of control variables, and $v_{i,c,t+h}$ is projection residual. I cluster standard errors at the country-year level to control for the potential country-year correlation across industries.[16] Clustered standard errors are also robust to the presence of arbitrary heteroskedasticity.[17]

# II.4 Results

## II.4.1 Main results

In this section, I present the main results based on the impulse response functions estimated using local projections. Figure II.3 shows local projection estimates and confidence intervals for the responses of wage inequality to a financial integration shock. The left panel of the figure represents the impulse response of the log 90-10 ratio of average wages, and the right panel represents the impulse response of the log 90-50 ratio.

The figure shows that financial integration increases between-firm wage inequality, and the effects are statistically significant and persistent. The estimated impulse responses of wage inequality are represented by a slightly sharp hump-shaped curve, which means that the magnitude of the response gradually increases, reaching a peak in the medium term and then declining. More precisely, a one standard deviation shock to financial integration increases the 90-10 ratio to a peak of about 2.91% five years after the shock. Similarly, a financial integration shock leads to about a 1.23% increase in the 90–50 ratio five years after the shock. It is worth pointing out that the similar shape of the responses in the 90-50 and 90–10 ratios implies that an increase in average wages of high-percentile firms mainly drives an increase in overall wage inequality. This could support the argument that financial integration has a favorable impact on big firms rather than SMEs.

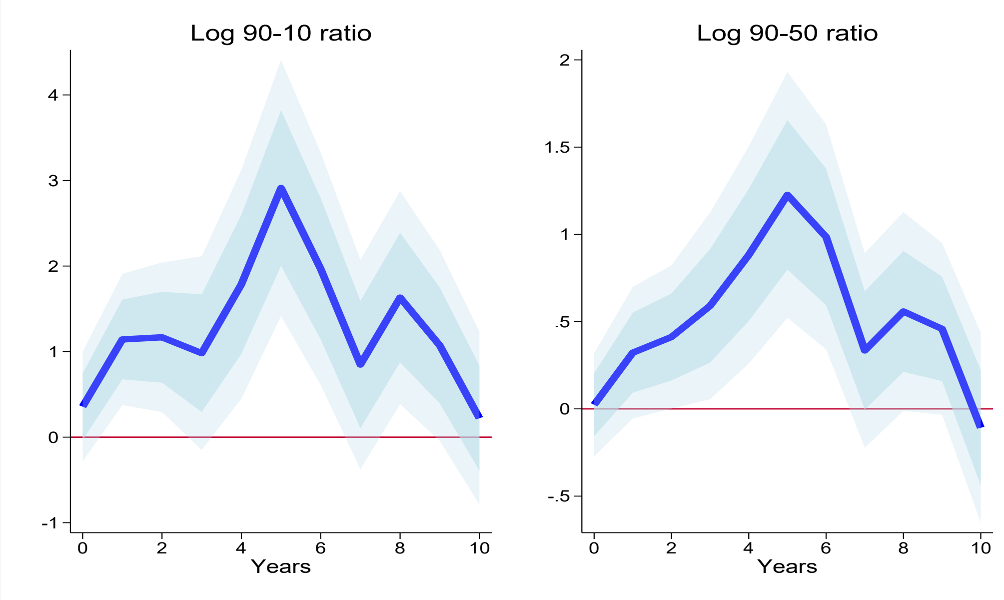
Next, taking advantage of using the *de facto* indicator, I split financial integration by direction, i.e., into inward and outward financial integration. I re-estimate the baseline model, equation (II.1), by dividing $\Delta FIN_{c,t}$ into external assets (outward

data across countries. For more detail, see the User Guide for the 9th Vintage of the CompNet Dataset, available at `https://www.comp-net.org/fileadmin/_compnet/user_upload/9th_Vintage_User_Guide_final_102023.pdf`

[16] Cameron and Miller (2015) argue that the inclusion of fixed effects could not ensure the elimination of the within-group correlation of the error. I assume that standard errors for a given year and country are correlated, while standard errors for different countries are uncorrelated. In this case, the robust standard error could be too small. Note that the cluster standard error can be unnecessarily large in some cases. See Abadie et al. (2023) for detailed discussion.

[17] Montiel Olea and Plagborg-Møller (2021) suggest that researchers should conduct inference based on lag-augmented local projections with heteroskedasticity-robust standard errors.

Figure II.3: Responses of between-firm wage inequality




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

financial integration) and external liabilities (inward financial integration). Inward and outward financial integration shocks can be conceptually interpreted as *gross* capital inflows and *gross* capital outflows, respectively.[18] The following equation is estimated

$$\begin{aligned} log(Inequality)_{i,c,t+h} &= \alpha_h log(Inequality)_{i,c,t-1} + \beta_{1,h}\Delta FIN.in_{c,t} \\ &+ \beta_{2,h}\Delta FIN.out_{c,t} + \mathbf{\Gamma}_h \mathbf{X}_{i,c,t} + \tau_i^h + \gamma_c^h + \delta_t^h + \upsilon_{i,c,t+h} \end{aligned} \tag{II.2}$$

where $FIN.in_{c,t}$ and $FIN.out_{c,t}$ represent inward and outward integration shocks, respectively. As in equation (II.1), $h = 0, 1, .., H$ denotes the time horizon, and the sequence of the parameters $\beta_{1,h}$ and $\beta_{2,h}$ as a function of time horizon $h$ provides the impulse response function to the inward and outward integration shock, respectively.

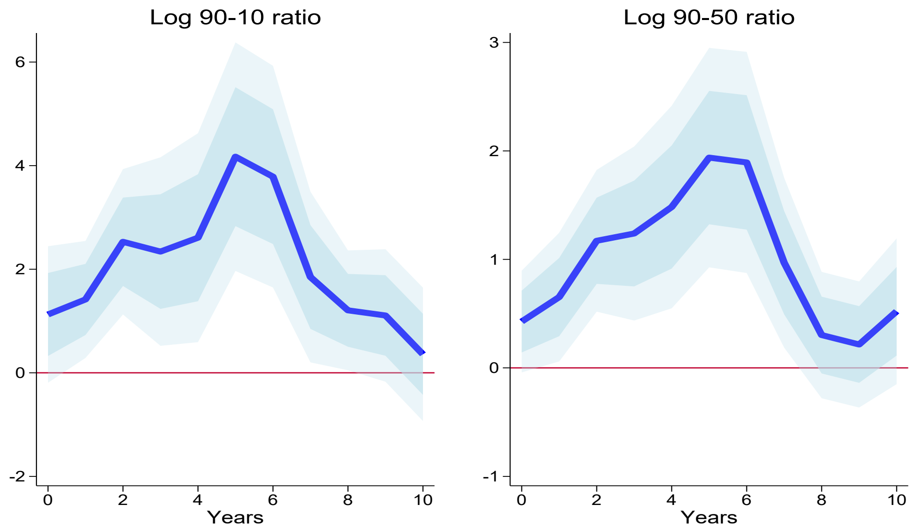
Figure II.4 presents the impulse responses of wage inequality to a shock in financial integration, estimated in equation (II.2). Panel (a) of Figure shows that inward financial integration increases between-firm wage inequality. The shape of the response is broadly similar to the hump in Figure II.3. However, the impacts of an inward financial integration shock are more pronounced than those of aggregate financial integration over the entire forecast horizon. A one standard deviation

[18]Gross capital inflows refer to a net increase in domestic financial assets of non-residents. On the other hand, gross capital outflows refer to a net increase in foreign financial assets of residents.

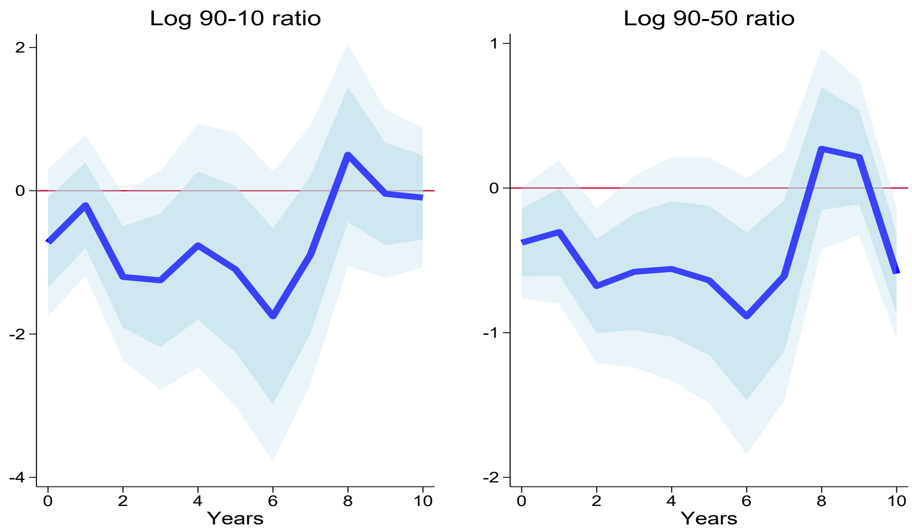
shock to inward financial integration leads to a 4.17% increase in the 90–10 ratio five years after the shock and a 3.79% increase six years after the shock. The 90-50 ratio increases by 1.94% and 1.89% five and six years after the shock, respectively. The estimated coefficients are statistically significant for 1 to 8 years after the shock.

Figure II.4: Responses of wage inequality - direction of financial integration

(a) Inward




(b) Outward




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

On the other hand, as shown in panel (b) of the figure, an outward shock leads to a reduction in between-firm wage inequality. An increase in outward financial integration by one standard deviation reduces the 90-10 ratio by about 0.20-1.75%

and the 90-50 ratio by about 0.30-0.89% until seven years after the shock. However, these results are less significant compared to the results for inward shocks.

In sum, the empirical findings presented in this section suggest that financial integration increases between-firm inequality and that the effect of inward financial integration is more pronounced than that of outward integration. These results are consistent with previous findings (Li & Su, 2021; Liu et al., 2023; Eklou & Foster, 2023), which suggest that increased income inequality (or wage inequality) is mainly associated with inward capital account liberalization rather than outward.

**Robustness**

I also conduct several robustness checks on the main results. First, I re-estimate equation (II.1) (i) without the lagged dependent variable and (ii) with the lagged value of the financial integration shock, respectively. Although impulse response functions estimated using local projections are known to be less sensitive to model misspecification than VARs, this robustness check tests the possibility that misspecification of the lag structure leads to bias in the estimates. Second, I exclude some observations from the full sample: (i) industries with outliers, which are values more than three standard deviations above the average wage inequality for each country, and (ii) observations during the global financial crisis and the COVID-19 pandemic. The second exclusion is because movements in macroeconomic variables, such as cross-border capital flows and wage inequality, were unusual during this period.

Next, to check whether the main results are robust to different sources of variation, I re-estimate the impulse response functions using the various mixtures of fixed effects: (i) industry-year and country fixed effects and (ii) industry-country and industry-year fixed effects. Finally, to address the potential bias in large $N$ and small $T$ settings due to the presence of a lagged dependent variable on the right-hand side of the estimated model and individual fixed effects (Nickell, 1981), I re-estimate impulse response functions using the GMM estimator proposed by Arellano and Bond (1991). I use the deeper lags of the dependent and independent variables as instruments for one lag term of the dependent variable.

I confirm that the main results are robust to a variety of other specifications, as shown in Figures II.A.2–II.A.5 in the Appendix. All results for robustness checks are close to those obtained in the baseline specification. Inward integration is associated with an increase in between-firm wage inequality, and the effect is statistically significant in all results. Outward integration, on the other hand, has an insignificant impact on wage inequality, although it is significant in some cases.

### II.4.2 Additional results

In this subsection, I provide plausible explanations for the following question, focusing on inward financial integration: How does financial integration affect between-firm wage inequality? This question concerns which channels and under what conditions financial integration affects wage inequality.

**Channel: labor productivity and capital intensity**

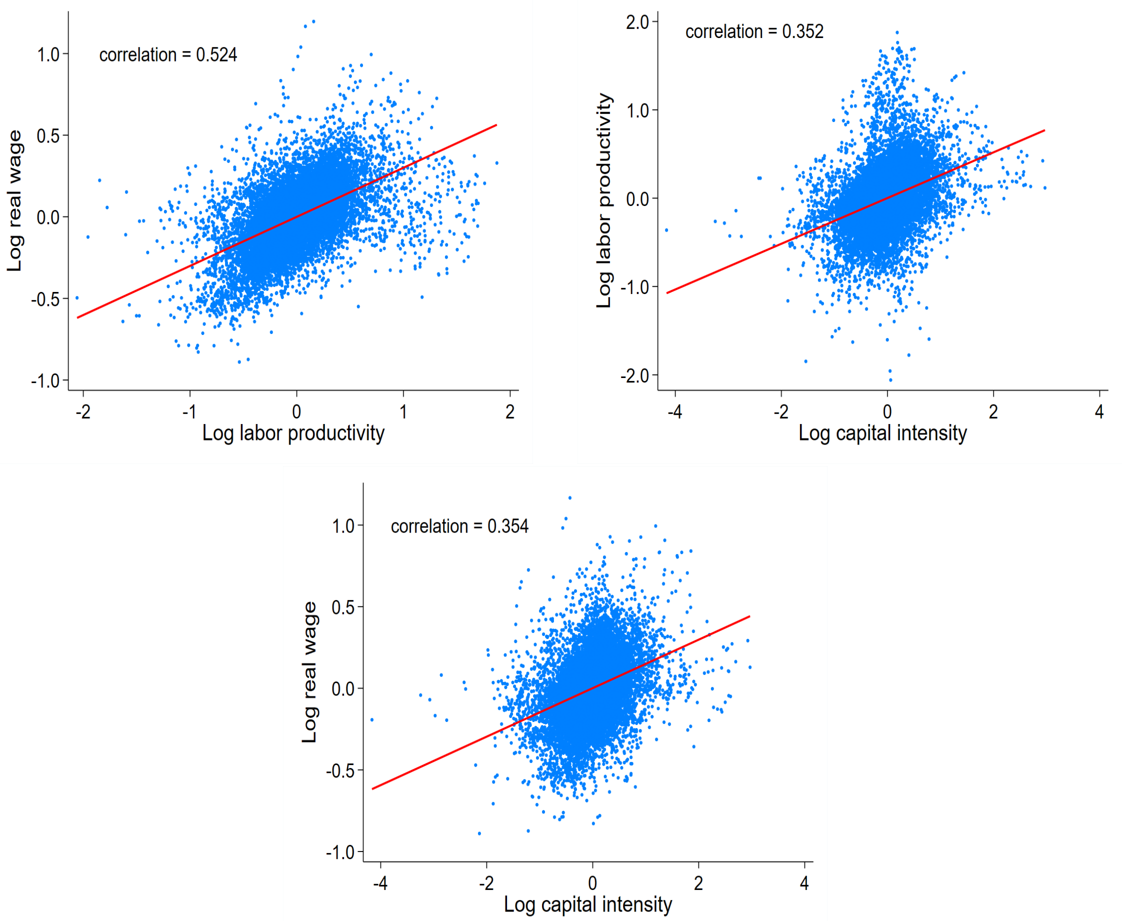
Progress in financial integration allows domestic firms to raise more external financing for physical and intangible capital than before. Additionally, the foreign capital inflows lead to the diffusion of technologies that are superior to domestic technologies. This makes firms more labor-productive and drives up wages. However, if these effects only involve big firms and not SMEs, then the productivity gap between big firms and SMEs widens as financial integration progresses, which in turn increases between-firm wage inequality.

We first look at the relationship between wage, labor productivity, and capital intensity using 2-digit industry-level data. Figure II.5 shows the relationship between three log-transformed variables. Labor productivity is computed as a ratio of real value added to the labor force, and capital intensity is estimated by the real capital-to-labor ratio. I adjust these variables by estimating the residuals of a panel regression using country- and industry-fixed effects. As shown in the upper left panel of the figure, there is a high correlation between labor productivity and wages. The upper right panel of the figure presents a positive correlation between capital intensity and labor productivity, and the bottom panel shows a positive relationship between capital intensity and wages. To put it simply, an increase in capital intensity may increase labor productivity and, consequently, also wages.[19]

Next, I replace the dependent variable $Ineuality_{i,c,t}$ in equation (II.2) with the labor productivity gap and capital intensity gap among firms within industries. These two variables are calculated in the same way as the calculation of between-firm wage inequality. Figure II.6 shows the impulse responses of the labor productivity gap to an inward financial integration shock. The 90-10 and 90–50 ratios of labor productivity positively respond to one standard deviation shock of financial integration, and the shape and magnitude of responses are broadly similar to wage inequality. The impacts are statistically significant over most forecast horizons and peak five years after the shock (5.39% rise in the 90-10 ratio and 2.44% rise in the 90-50 ratio).

[19]Many economists report that the capital-to-labor ratio is positively correlated with wages. Arai (2003) finds that the capital-to-labor ratio systematically positively affects wages using the matched worker-firm data from Sweden. Leonardi (2007) provides empirical evidence of a positive relationship between the dispersion of the capital-to-labor ratio and wage inequality in the US.

Figure II.5: Wage, labor productivity, and capital intensity

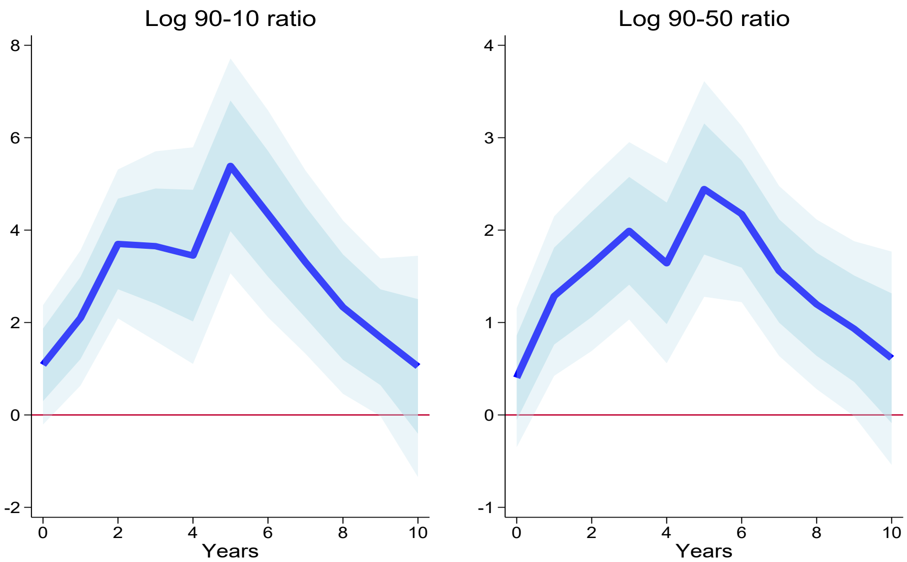



*Source*: CompNet dataset and author's calculation

The results presented in Figure II.7 suggest that the capital intensity gap also rises in response to an inward financial integration shock. It shows that as financial integration progresses, firms that have already accumulated relatively more capital acquire additional capital faster than firms with relatively less capital. The effects are large, statistically significant, and persistent. More precisely, inward financial integration is significantly associated with 5.12-8.99% increases in the 90-10 ratio of capital intensity at each horizon between one and seven years after the shock. Also, in response to a shock, the 90-50 ratio of capital intensity increases by 2.93-3.99% from one year to eight years after the shock. It is not hard to accept that the response of the capital intensity gap is larger and more statistically significant than that of the labor productivity gap and wage inequality because firms can accumulate more capital as a direct result of progress in financial integration.

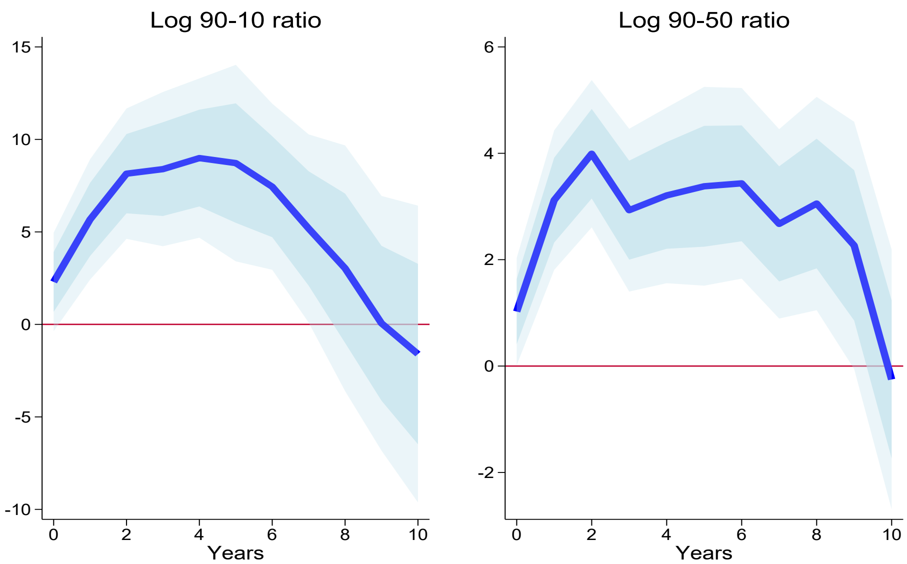
As in the main result in subsection II.4.1, the similar response shapes of the 90-10 and 90–50 ratios imply that big firms primarily drive the widening dispersion in

Figure II.6: Responses of the labor productivity gap




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

Figure II.7: Responses of capital intensity gap




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

labor productivity and capital intensity. Figures II.A.6 and II.A.7 in the Appendix show the difference in labor productivity and capital intensity across quintiles of firm size within industries in a similar way to Figure II.1. These figures suggest that the labor productivity gap and capital intensity gap across firms within industries exist mainly between big firms and SMEs.

To summarize the results in this part, financial integration shocks widen the capital intensity gap within industries, which in turn widens the labor productivity gap. These results suggest that financial integration fosters capital accumulation and productivity gains for big firms, while SMEs benefit less from financial integration. Through this mechanism, financial integration affects between-firm wage inequality.

**External financial dependence**

Several literature has pointed out that the distributional effects of financial integration depend on a sector's inherent necessity for external financing. The previous empirical research exploits the external financial dependence (EFD) index developed by Rajan and Zingales (1998) to identify external financing needs for each sector (Larrain, 2015; Furceri et al., 2019; Eklou & Foster, 2023). EFD is defined as the fraction of capital expenditures not financed by internal cash flow from operations. Rajan and Zingales (1998) constructed the EFD index as a ratio of total capital expenditures minus current cash flow to total capital expenditures. Unfortunately, the original estimation of the EFD index by Rajan and Zingales (1998) cannot be replicated using the variables included in the CompNet dataset. As an alternative, I use the "financing gap" variable from the CompNet dataset. The financing gap is the ratio of nominal investment (fixed investment plus the change in net working capital) minus cash flow to nominal revenue. This variable measures indirectly the industry's demand for external financing over a given period. Although the measuring method of this variable is slightly different from EFD by Rajan and Zingales (1998), it conceptually captures similar characteristics of individual industries.

To "smooth temporal fluctuations and reduce the effects of outliers" (Rajan & Zingales, 1998, p.564), I calculate the mean value of the financing gap for the entire sample period for each industry instead of using values in a given year. Additionally, as in original and other applications, I use the median firm's value in a given industry rather than the industry average. It is also important to note that the absolute value of the financing gap for each industry cannot be compared across countries because the financing gap is endogenous to the development of each country's financial system and credit supply. Therefore, I use the financing gap only as an indicator to rank industries within countries by their inherent demand for external financing and do not use its value directly in the analysis. With the rank of industries for each country, I divide them into industries with relatively higher external financing dependency (above the mean of all industries in a country) and industries with relatively lower external financing dependency (below the mean of all industries in a country).[20,21]

Figure II.8 presents the estimated impulse response functions of wage inequality for both industry groups. The results presented here are based on a panel of 14 countries since six countries[22] are excluded from the whole sample countries due to no data availability on the financing gap. The figure suggests that inward financial integration is more significantly associated with industries with relatively higher EFD than industries with relatively lower EFD.

As shown in panel (a) of the figure, in response to an increase in total external liabilities by one standard deviation, the 90-10 ratio of wage inequality in the industries with high EFD rises by 3.06-3.34% in the medium term (five and six years after the shock), and the 90-50 ratio rises by 1.43-1.52% at the same forecast horizon. The effects of the shock are statistically significant from one year to six years after the shock. On the other hand, the results presented in panel (b) of the figure suggest that wage inequality in industries with low EFD is less responsive to an inward financial integration shock than in industries with high EFD. In industries with low EFD, the 90-10 ratio increases by 0.25-1.50% over the period in response to a shock, but these effects are not statistically significant over most forecast horizons. The 90-50 ratio increases statistically significantly by 1.15% five years after the shock, but the effects at other forecast horizons are small and statistically insignificant.

These results are broadly consistent with previous empirical findings by Larrain (2015) and Furceri et al. (2019), which argued that the distributional effect of financial integration is larger in sectors with high EFD than in sectors with low EFD, although their studies did not address between-firm wage inequality.[23] However, these differ from the results in the ASEAN 5 countries by Eklou and Foster (2023).[24]

[20]Rajan and Zingales (1998) assumed that EFD does not vary across countries, but recent research has shown that this assumption is not supported by empirical evidence.(Villani, 2021; Eppinger & Neugebauer, 2022) I adopt this recent argument and calculate the ranking of industries for each country.

[21]In most countries, construction, wholesale, retail, postal and courier activities, accommodation, and real estate activities are classified as industries with high EFD. It is consistent with existing studies using the EFD (e.g., Larrain (2015), Furceri et al. (2019) among others). The sub-industries of manufacturing are categorized very differently across countries.

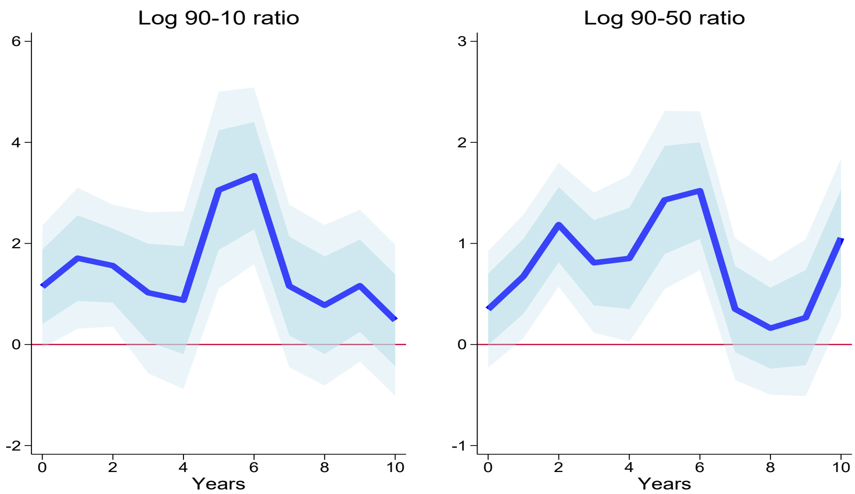
[22]Hungary, Latvia, Lithuania, Netherlands, Romania, and Slovenia

[23]Larrain (2015) has focused on wage inequality between skill- and unskilled labor, and Furceri et al. (2019) has examined the aggregate distributional effects of capital account liberalization using the Gini coefficient and labor share as measures of inequality.

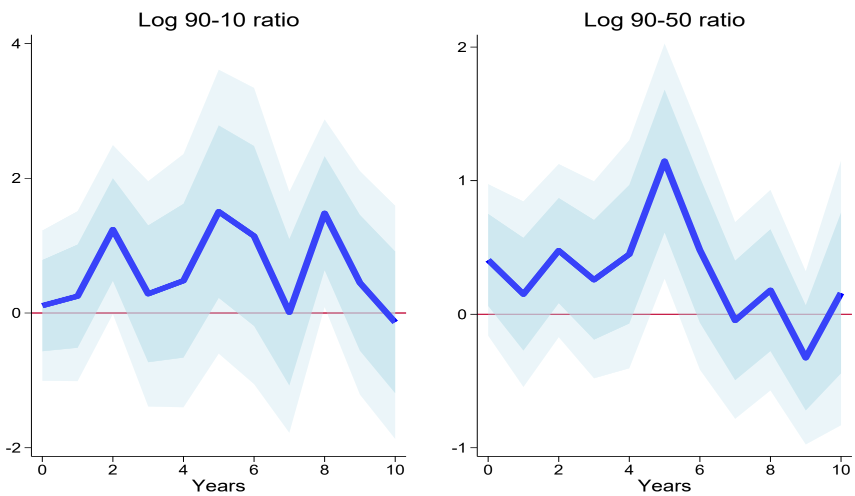
[24]As Eklou and Foster (2023) argued, their results may be because ASEAN5 firms are more financially constrained than firms in Europe or advanced economies.

Figure II.8: Responses of wage inequality - external financing dependence

(a) Industries with high EFD




(b) Industries with low EFD




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

### Financial development

The level of financial development has been discussed in the literature as a moderator of the distributional effects of financial integration. However, there has been conflicting empirical evidence about whether financial development exacerbates or mitigates the effects of financial integration on inequality. For example, Bumann and Lensink (2016) suggest that capital account liberalization can lower income inequality if the financial depth, measured by private credit over GDP, is sufficiently high. On the contrary, de Haan and Sturm (2017) find that financial development enhances the effect of financial liberalization, exacerbating income inequality. Eklou and Foster (2023) also show that financial development dampens the impact of capital account liberalization on wage inequality.

To examine whether the effects of financial integration on the between-firm wage inequality depend on the level of financial development, I employ the Financial Development (FD) index developed by Sahay et al. (2015) and Svirydzenka (2016) and maintained by the IMF. The FD index has a broader coverage and provides nine indices, the multi-dimensional measures for financial market/institution development in terms of depth, accessibility, and efficiency. I rank the sample countries using the aggregated index and divide them into two groups: countries with relatively higher financial development and countries with relatively lower financial development.[25] Table II.A.1 in Appendix reports summary statistics for the FD index by country groups.

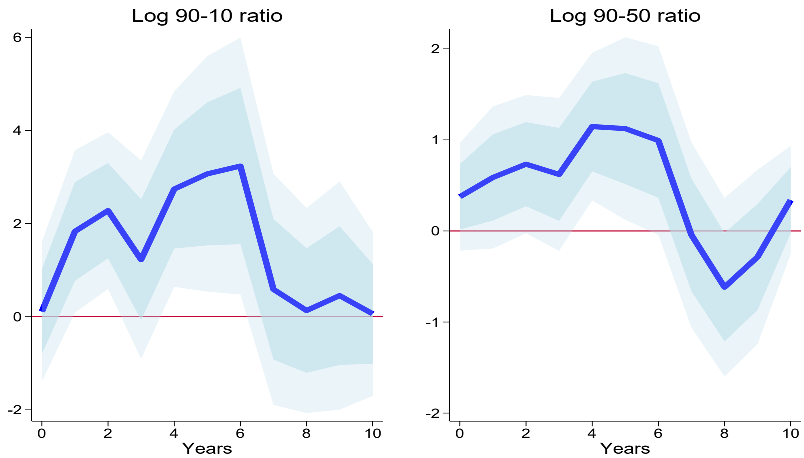
Figure II.9 shows the estimated impulse response functions of wage inequality for both country groups. As before, an inward financial integration shock increases between-firm wage inequality in both country groups, and the effects peak in the medium term. However, the magnitude and shape of estimated impulse response functions differ slightly between the two country groups. In countries with low FD, an inward financial integration shock increases the 90-10 and 90–50 ratios by 4.63% and 1.63% at a peak (5 years after the shock), respectively. In countries with high FD, the 90-10 ratio increases by 3.24% six years after the shock, and the 90-50 ratio increases by 1.15% five years after the shock. These results suggest that the effects are more prominent in countries with low FD than those with high FD. In other words, financial development can mitigate the negative impact of financial integration on wage inequality. However, the difference in the magnitude of effects between the two groups is insufficient to conclude that this alleviation effect is sufficiently significant. Thus, it should not be interpreted as definitive.

In addition, the countries categorized as having relatively low FD (mainly Central and Eastern European countries that were formerly centrally planned economies) also integrated rapidly into international financial markets after joining the EU in the mid-2000s. It would be premature to conclude, based on the results presented here, that financial development does not mitigate the distributional effects of financial integration. Further analysis should be conducted comparing developing countries outside of Europe to obtain more precise implications.

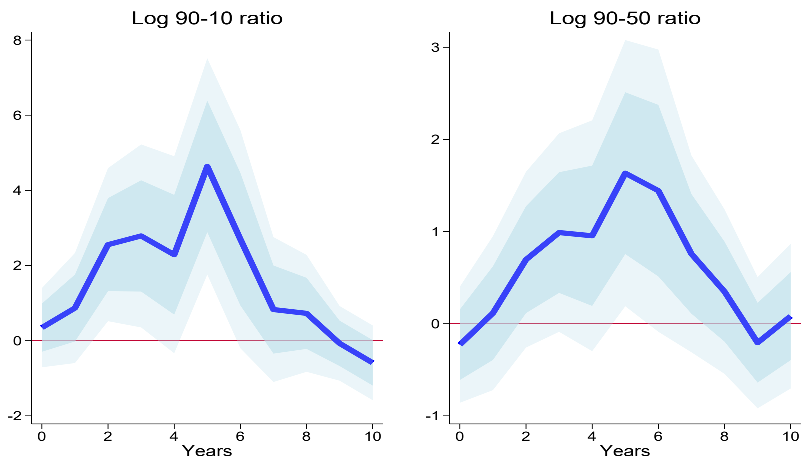
[25] Higher FD countries: Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Portugal, Spain, Sweden, and Switzerland. Lower FD countries: Croatia, Czech Republic, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia, and Slovenia.

Figure II.9: Responses of wage inequality - financial deveploment

(a) Countries with high FD




(b) Countries with low FD




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

### Categories of financial integration

Different types of financial integration may have different distributional implications.[26] EWN database provides data on subtypes of total external liabilities by dividing it into the following three categories: (i) foreign direct investment (FDI), (ii) portfolio investment, including equity and bond investments, and (iii) other investments, including loans to or deposits in non-resident entities, trade credits, etc. It allows us to look more closely at the effects of financial integration on wage inequality, separately by its subtypes.

[26]For a comprehensive discussion on the distributional effects of subtypes of financial integration, see Eichengreen et al. (2021).

To investigate whether the subcomponents of financial integration matter for its impact on between-firm wage inequality, I re-estimate the baseline model, equation (II.1), using data on three subtypes. Each panel in Figure II.10 presents the impulse response functions of wage inequality to shocks in FDI, portfolio investment, and other investment, respectively. As shown in the figure, all three subcomponents contribute to an increase in between-firm wage inequality. However, the magnitude, statistical significance, and time horizons of the effects vary somewhat depending upon the types. Among them, FDI has the most important impact. An increase in FDI by one standard deviation leads to 0.97-1.56% increases in the 90-10 ratio from one year to six years after the shock and 0.68-1.03% increases in the 90-50 ratio over the same horizons. Portfolio investment contemporaneously increases the 90-10 ratio by 1.00% and the 90-50 ratio by 0.41%. Its effects are long-lasting up to five years after the shock. In contrast, other investment has a significant impact over a longer lag than the previous two types. Its impacts on wage inequality peak in the medium term, six and seven years after the shock (1.59-1.73% rise in the 90-10 ratio and 0.89-0.99% rise in the 90-50 ratio).

While the existing literature studying the distributional effects of financial integration by its subtypes provides no homologous empirical evidence[27], the findings in this paper suggest that all three subcomponents are noteworthy for examining the impacts of financial integration on inequality. Additionally, the fact that each type affects wage inequality at different time horizons implies that there may be specific channels of distributional effects for each type. Future research could extend the results presented here to investigate the various channels of financial integration's impact on wage inequality by separately identifying effects by subtype.

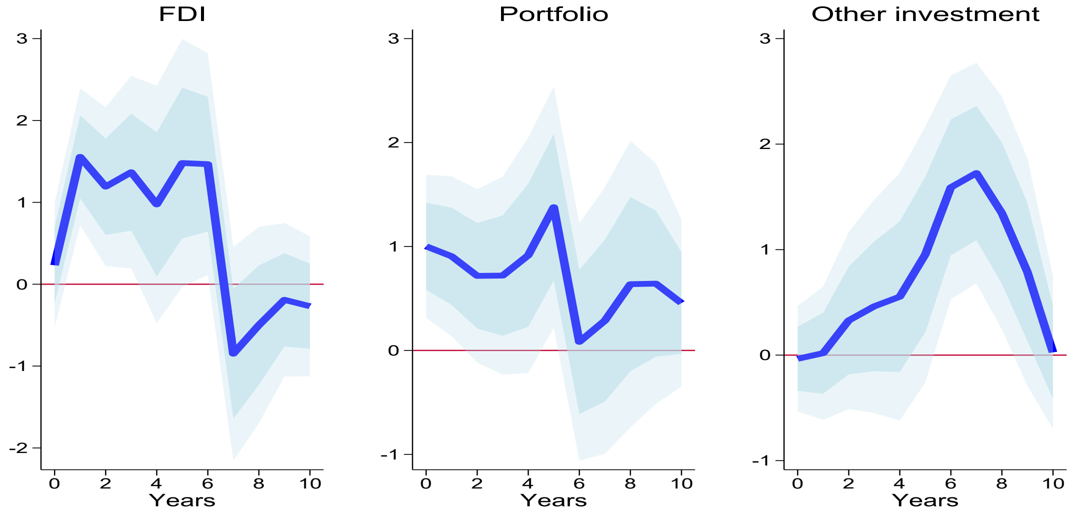
## II.5 Conclusion

In this paper, I empirically examine the effects of financial integration on between-firm wage inequality using an unbalanced panel for 20 European countries from 1999 to 2021. To measure between-firm wage inequality, I exploit the CompNet dataset, which provides micro-aggregated data based on firm-level data. To focus on the mid- and long-term effects, I estimate the impulse response functions using local projections with multi-dimensional (industry-country-year) panel regressions.

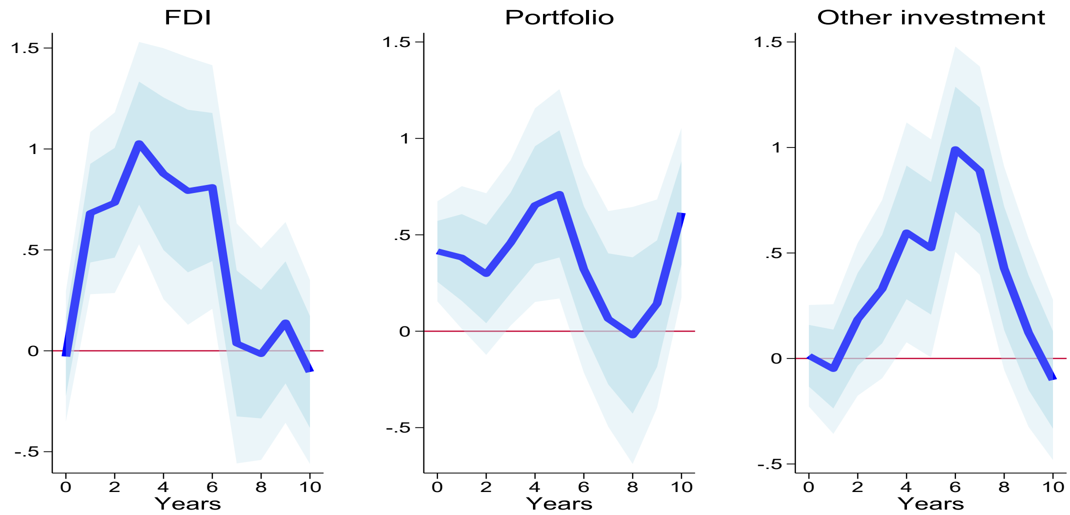
[27]For instance, Jaumotte et al. (2013) and Asteriou et al. (2014) suggest that FDI is positively correlated with increased income inequality. In contrast, Li and Su (2021) report no significant association between FDI and income inequality, while equity market liberalization is associated with increased income inequality. Eklou and Foster (2023) show that inward liberalization of equity market and FDI increases wage inequality in ASEAN 5 countries.

Figure II.10: Responses of wage inequality - subtypes of financial integration

(a) Log 90-10 ratio




(b) Log 90-50 ratio




*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

I find that financial integration at the national level increases between-firm wage inequality within industries. An increase in financing integration by one standard deviation leads to a significant rise in wage inequality, particularly in the medium term. The empirical findings in this paper also suggest that the direction of financial integration matters. Inward financial integration increases between-firm wage inequality, and these effects are statistically significant and persist over time. Outward integration induces a reduction in wage inequality, but the effects are not statistically significant for most forecast horizons. These results are robust to various robustness checks and are broadly consistent with previous research findings.

I also provide empirical evidence on the channel for the distributional effects of financial integration. Similar to the response to wage inequality, the labor productivity gap and capital intensity gap among firms within industries widen in response to a financial integration shock. These results suggest that financial integration accelerates big firms to accumulate more capital and increase productivity, while SMEs benefit less from financial integration. These results provide a milestone to build a theoretical framework for the distributional effects of financial integration. Additionally, financial integration is more significantly associated with a rise in wage inequality in industries with relatively higher external financial dependence than in industries with relatively lower external financial dependence. The financial development of countries may mitigate the adverse impact of financial integration on wage inequality, but it is less definitive. A comparison of the impacts of subcomponents of financial integration (FDI, portfolio investment, and other investment) reveals that all subcomponents contribute to widening wage inequality, although the magnitude and lag structure of the effect vary slightly.

This paper has some limitations. First, as the data is aggregated at the industry level, heterogeneity in workers' characteristics, such as education level, work experience, and gender, is not sufficiently controlled. Accordingly, the results presented in this paper cannot be interpreted as a pure reflection of "between-firm wage inequality" but rather as a mixture of "within-firm wage inequality" to some extent. Controlling for worker characteristics would require using more comprehensive data on employees linked to their employers. Second, the analysis in this paper focuses on European countries, which are primarily advanced economies and already highly integrated into the global financial markets. This study, therefore, should be extended to the case of developing countries in other regions.

The findings in this paper suggest the proper policy actions that could be taken to mitigate the adverse distributional consequences of financial integration. As financial integration progresses, policymakers should design policies that ensure the benefits of financial integration are more evenly distributed across all firms and all workers. We can consider a package of policies to support SMEs in raising more funds from foreign capital, to increase investments in workers' education level and skills, and to manage cross-border capital flows appropriately. The implementation of these policies enables policymakers to distribute the benefits of financial integration more equitably, fostering inclusive and sustainable economic growth for all.

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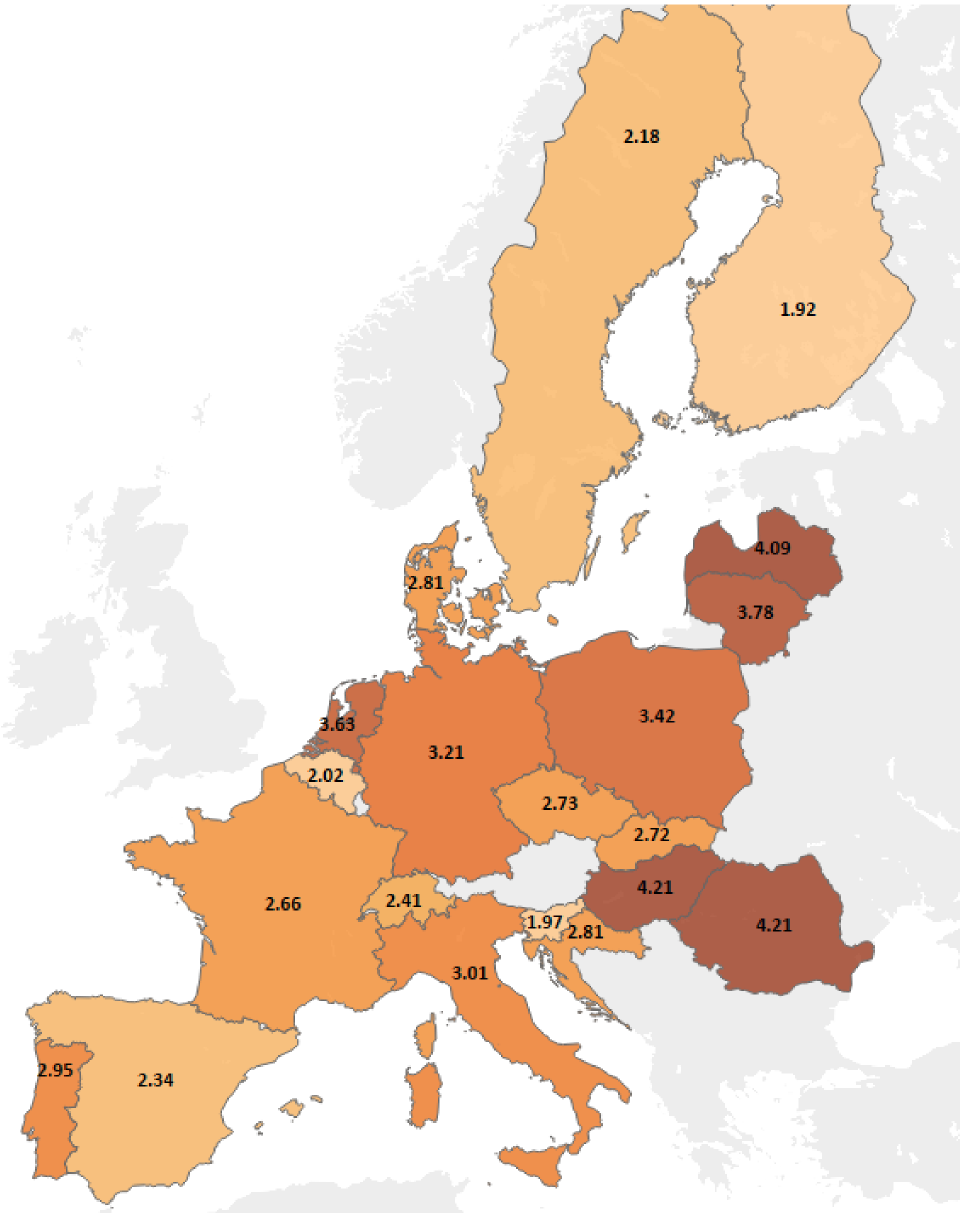
## II.A Appendix

Table II.A.1: Descriptive statistics: FD index by country groups

| | Mean | Median | SD | Min | Max |
|---|---|---|---|---|---|
| Countries with high FD | 0.740 | 0.737 | 0.095 | 0.519 | 0.987 |
| Countries with low FD | 0.364 | 0.361 | 0.111 | 0.144 | 0.571 |
| Total | 0.565 | 0.597 | 0.214 | 0.144 | 0.987 |

*Source*: IMF and author's calculation

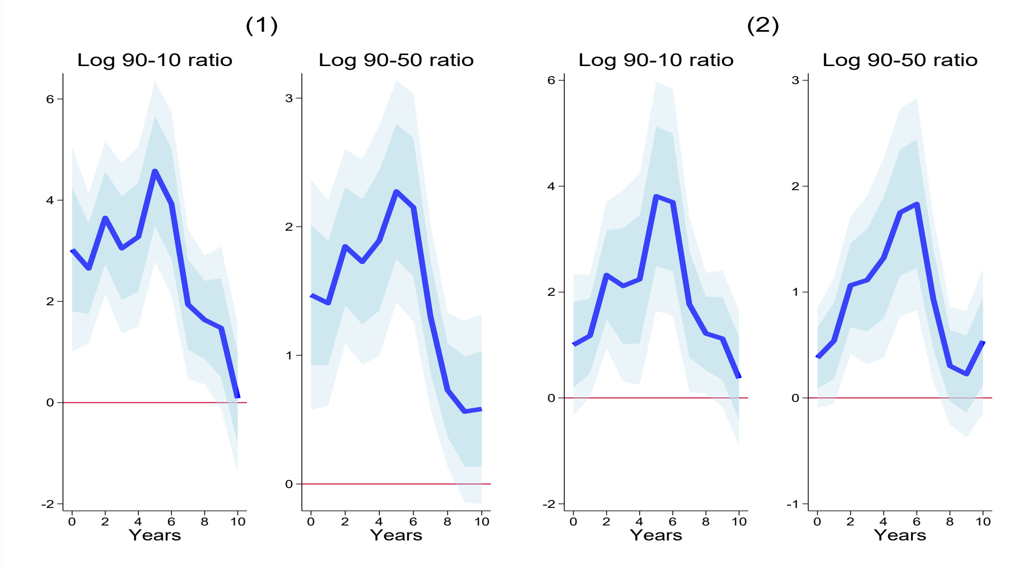
Figure II.A.1: Wage inequality (90-10 ratio) by country



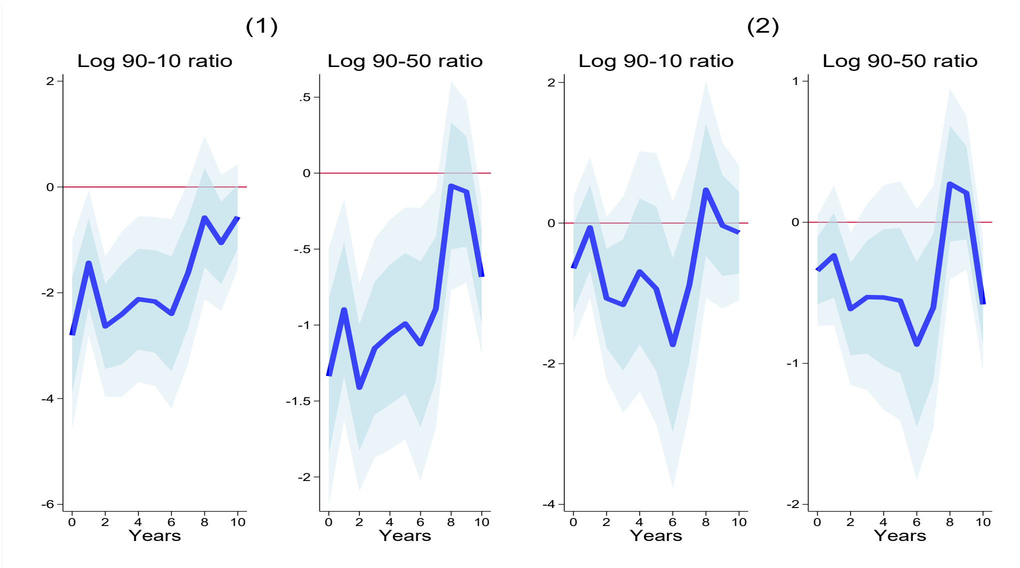

*Source*: CompNet dataset and author's calculation

Figure II.A.2: Robustness check (1) - other specifications

(a) Inward



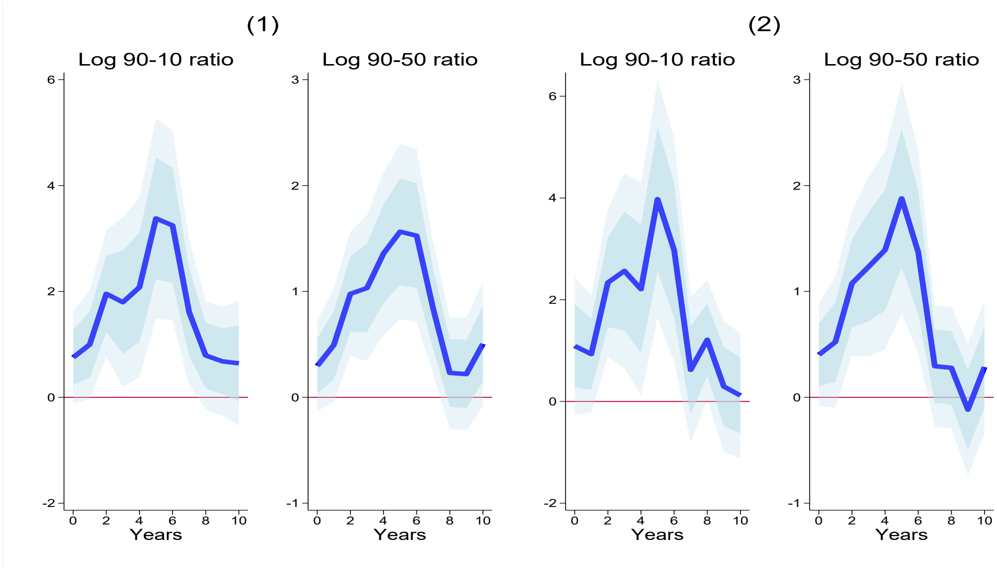

(b) Outward



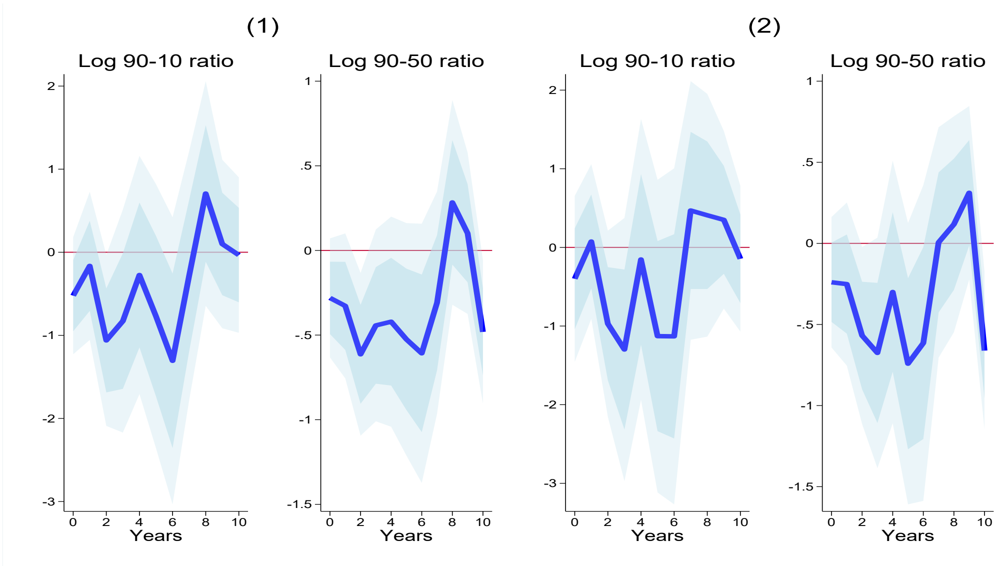

*Notes*: (1) the impulse response function for the specification without the lagged dependent variable, (2) the impulse response function for the specification with the lagged value of the financial integration shock. The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

Figure II.A.3: Robustness check (2) - excluding some observations

(a) Inward



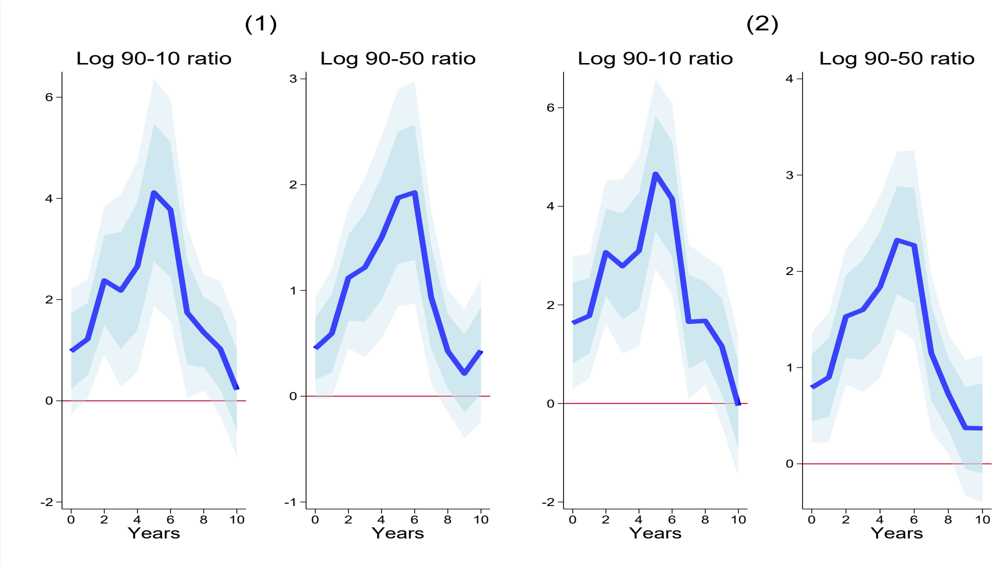

(b) Outward



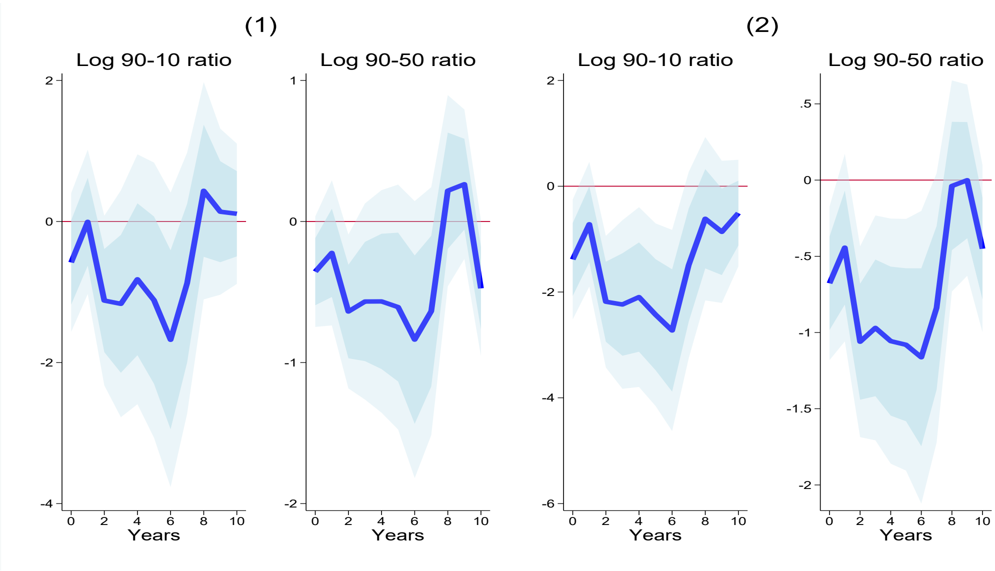

*Notes*: (1) excluding outliers, (2) excluding observations in 2008 and 2020. The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

Figure II.A.4: Robustness check (3) - various mixtures of fixed effects

(a) Inward



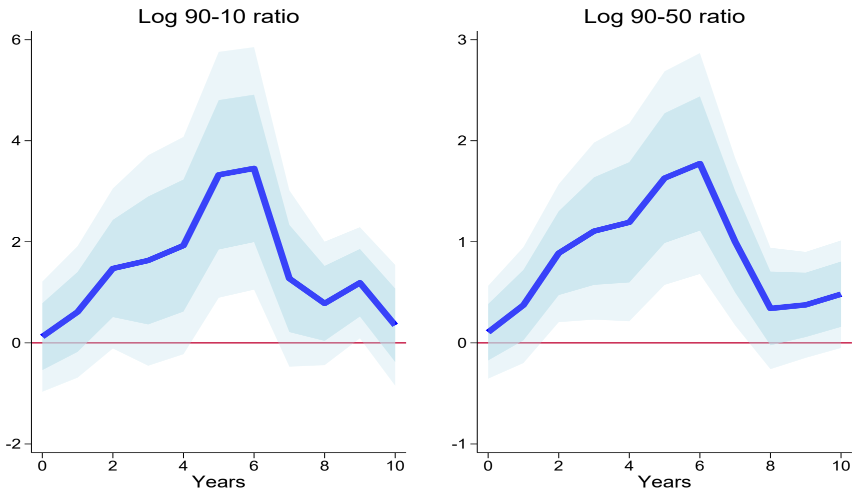

(b) Outward



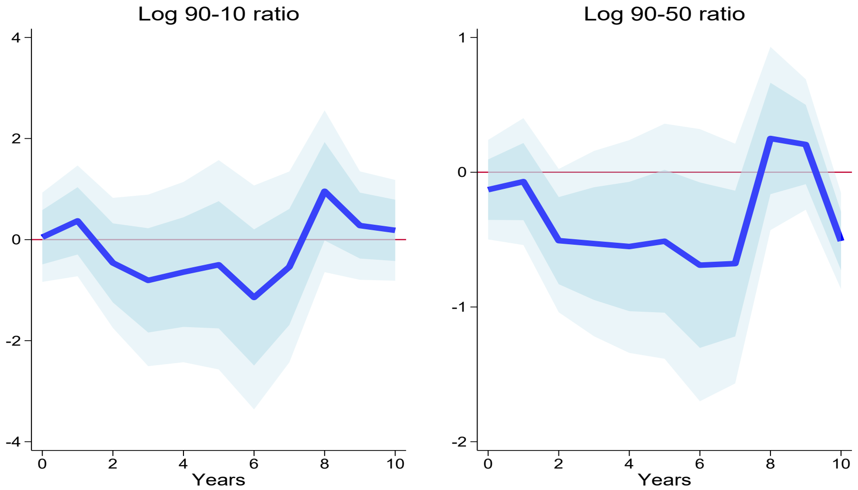

*Notes*: (1) country and industry-year fixed effects, (2) industry-country and industry-year fixed effects. The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

Figure II.A.5: Robustness check (4) - GMM estimator

(a) Inward



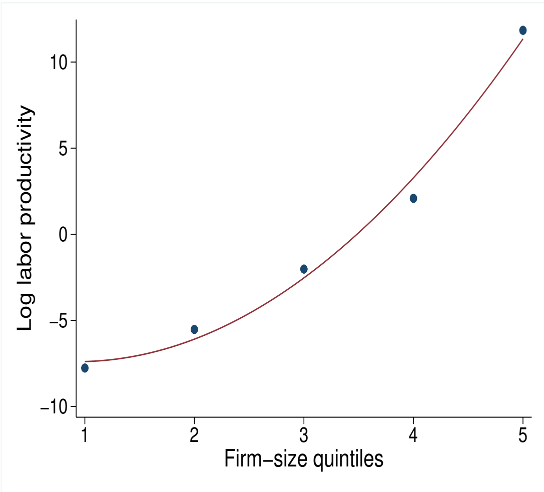

(b) Outward



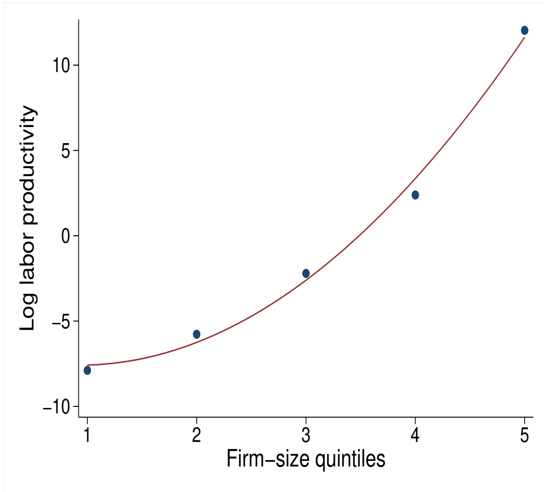

*Notes*: The solid blue line represents the point estimate of the $\beta_h$ coefficient, and the shaded areas are the 68% and 90% confidence bands constructed from clustered standard errors.

Figure II.A.6: Labor productivity by firm size

(a) Mean within each quintile



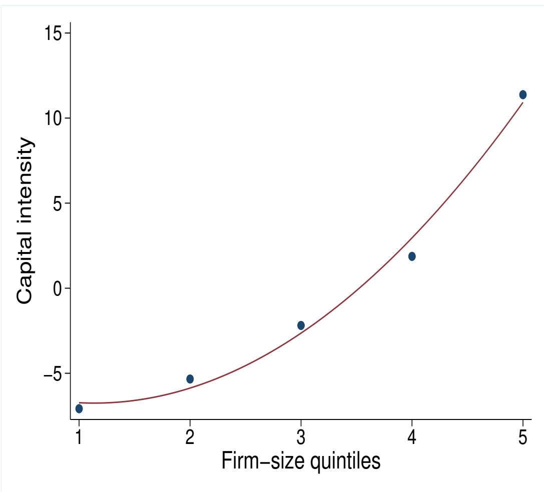

(b) Median within each quintile



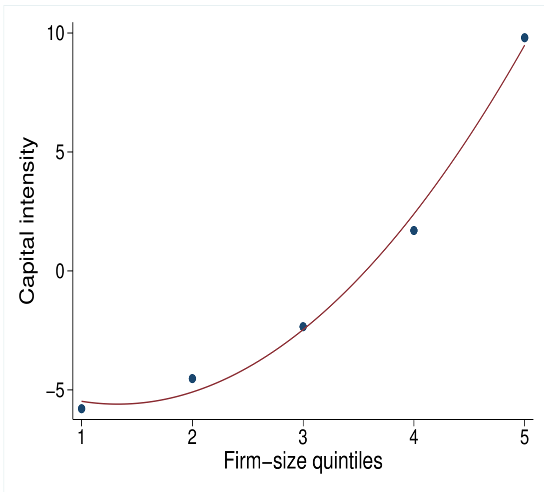

*Notes*: Quintile 1; 0-20% percentiles, Quintile 2; 20-40% percentiles, Quintile 3; 40-60% percentiles, Quintile 4; 60-80% percentiles, and Quintile 5; 80-100% percentiles.
*Source*: CompNet dataset and author's calculation

Figure II.A.7: Capital intensity by firm size

(a) Mean within each quintile




(b) Median within each quintile




*Notes*: Quintile 1; 0-20% percentiles, Quintile 2; 20-40% percentiles, Quintile 3; 40-60% percentiles, Quintile 4; 60-80% percentiles, and Quintile 5; 80-100% percentiles.
*Source*: CompNet dataset and author's calculation

# III. Gains from Gunboats: U.S. Aircraft Carriers and Financial Markets in East Asia

This paper was presented at the following international conference:

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Presenter: Jinyeong Yun

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# Gains from Gunboats: U.S. Aircraft Carriers and Financial Markets in East Asia

Peter Tillmann[†] Jinyeong Yun[‡]

**Abstract**

Asia, in particular East Asia and the South China Sea, is one of the regions that are particularly exposed to geopolitical risk. We study the impact of an increased presence of the U.S. military in East Asia on asset prices and capital inflows in the region. For that purpose, we construct a novel index from local newspapers reporting on the presence of U.S. Carrier Strike Groups (CSG) and estimate the market impact of a surprise change in this index. Our findings reveal that a higher military presence causes an increase in stock prices, an appreciation of the local currency, and an inflow of foreign capital, but also a higher level of geopolitical risk. Interestingly, some expansionary effects disappear or change the sign during the first Trump administration, suggesting that the market benefits are tied to overall U.S. strategic policy frameworks. We also find substantial heterogeneity in the effects. Market responses are positive when a deployment of CSG pertains to North Korean provocations, but muted or even negative when related to tensions with China. Our findings underscore the financial gains from the projection of U.S. military power, and also demonstrate that financial market benefits vary depending on geopolitical contexts and across U.S. administrations.

**Keywords**: Geopolitical risk, U.S. military presence, China, financial markets, Asia
***JEL* codes**: F51, E44, F21

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## III.1 Introduction

Over the past few years, geopolitical risk has re-emerged as an important driving force of economic and financial development. The increasing number of geopolitical crises, such as Russia's full invasion of Ukraine in 2022 and the military escalation in the Middle East in 2023, put geopolitics back on the agenda of economic agents. One important region where risk is permanently high and the danger of further escalations is looming, is East Asia. While tensions between North and South Korea have been in the headlines for years, tensions between China and Taiwan have intensified strongly since 2019-2020. Also, China pushes its interests in the South China Sea with increasingly aggressive means. Often the U.S. responds to increasing tensions in East Asia by deploying additional aircraft carriers to the region. For example, on July 4, 2020, *The New York Times* reported:

> "With Beijing's Military Nearby, U.S. Sends 2 Aircraft Carriers to the South China Sea - The deployment of an American aircraft carrier and its strike force is often used as a signal to deter adversaries. Deploying two at once is recognized as a significant show of force."[1]

In this paper, we study the impact of an increased presence of the U.S. military in the region on asset prices and capital inflows. For that purpose, we use data on local English-language newspaper articles reporting on the presence of the U.S. Carrier Strike Groups (hereafter CSG) in the region. CSG are formations of battleships of the U.S. Navy around a nuclear-powered aircraft carrier. They serve as a main tool for the projection of power overseas. We ask whether the U.S. deploying CSG to the region, for a given level of geopolitical risk, affects financial markets.

We proceed in three steps. The first step is to derive a measure of the surprising presence of CSG in the region. Importantly, measures of the changing military presence as such are not sufficient for our purpose because many of these changes are anticipated. This information should already be incorporated in asset prices, such as stock prices or exchange rates. Financial markets should respond to news about the military presence only, i.e., the unexpected presence of the U.S. Navy. From a range of East Asian English-language newspapers, including the *South China Morning Post* and *Taiwan News*, among others, we construct an index of the number of articles on U.S. CSG in the region. We purge this series from all predictable components by regressing it on lags of itself and the Geopolitical Risk (GPR) index by Caldara and Iacoviello (2022). The residual of this regression serves as our

[1] See `https://www.nytimes.com/2020/07/04/us/politics/south-china-sea-aircraft-carrier.html`.

measure of unexpected changes in the presence of U.S. CSG in the region.

In the second step, we estimate the effect of the shocks to the CSG index on asset prices and capital inflows in the financial markets of seven East Asian countries (South Korea, Taiwan, Japan, the Philippines, Vietnam, Malaysia, and Indonesia) using monthly data from January 2010 to October 2024. Hence, we look at the market responses in the region, not the response of U.S. financial markets. We show that a surprise increase in newspaper articles reporting about CSG boosts asset prices. A one standard deviation increase causes stock prices to increase and the nominal exchange rates to appreciate. In addition, we find that equity inflows increase significantly if the military presence of the Navy increases. Overall, our results suggest that deploying aircraft carriers to East Asia is good news for financial markets. Hence, there are indeed financial gains from gunboats in the region.

In the third step, we extend our model into two directions. We distinguish between newspapers reporting on CSG that relate to tensions in North Korea and tensions involving China, respectively. We find that the market responds differently to either information. Stock prices and the local currencies strongly increase if the U.S. sends the navy due to heightened tensions with North Korea. However, these expansionary effects are less clear if the navy is deployed in response to tensions with China. We also distinguish between the effects before, during, and after the first Trump administration. Under President Trump, our main results are absent or even change sign.

How can we rationalize our findings? We could think of two main transmission channels that could explain our results. The increased presence of the biggest super power causes a drop in geopolitical uncertainty and an increase in confidence. This should make a country allied with the U.S. more attractive to foreign investors. As a result, asset prices should increase. A second channel operates through international supply-chain and trade linkages. When the U.S. deploys CSG to the region, the safety of shipping routes increases. This should be supportive for exports, imports, and production and should ultimately also be reflected in asset prices. There are also two effects operating into the opposite direction. Deploying military power to the region could contribute to an escalation of tensions, thus increasing rather than decreasing risks. Also, the higher presence of CSG in East Asia or the South China Sea necessarily reduces the presence in other parts of the world, e.g., in the Red Sea or the Persian Gulf. This could increase the risk of conflicts erupting elsewhere. In this paper, we remain agnostic and do not take a stand on the specific transmission mechanisms. We let the data speak and study the overall effect of the CSG presence on financial markets.

Using the presence of CSG to measure the projection of military power has several advantages. First, we can assume that it is exogenous to movements in macroeconomic variables, especially at the domestic level. Second, the presence of CSG requires no infrastructure spending in host countries, as they do not reflect a permanent presence. This allows us to more clearly study the confidence or uncertainty channel, while permanent military bases in host countries come with additional channels such as additional jobs for locals, spending by military personnel, infrastructure investment, etc.

Our paper is related to the strand of literature that employs textual analysis to measure uncertainty and geopolitical risk. For example, Baker et al. (2016) and Caldara and Iacoviello (2022) draw on textual data from English-language newspapers in the U.S. and the United Kingdom to construct indicators of economic policy uncertainty and geopolitical risk, respectively. They show that an increase in political uncertainty or geopolitical risk has contractionary effects on the U.S. economy. A growing body of literature has quickly embraced these indicators, in particular the GPR index of Caldara and Iacoviello (2022), in order to estimate the market responses to changes in geopolitical risk (Smales, 2021; Bussy & Zheng, 2023; Feng et al., 2023; Agoraki et al., 2024; Brignone et al., 2024; Mignon & Saadaoui, 2024; Yilmazkuday, 2024; Wang et al., 2024; Choi & Havel, 2025, among them).[2] Feng et al. (2023) and Agoraki et al. (2024) suggest that rising geopolitical risk reduces international capital flows. Similarly, Bussy and Zheng (2023) find that escalated geopolitical risks deter Foreign Direct Investment (FDI). Using country-specific GPR indices, Choi and Havel (2025) find that U.S. portfolio investment in emerging market economies decreases as the geopolitical risk of those countries rises. Yilmazkuday (2024) documents that a shock to global geopolitical risk leads to significant declines in stock prices for 29 economies. Smales (2021) and Mignon and Saadaoui (2024) show that shocks to the GPR index drive up oil prices, along with fears of oil supply disruption. Wang et al. (2024) find that the GPR index is negatively associated with corporate investment. Brignone et al. (2024) investigate potential non-linearities in the responses of economic variables to geopolitical risk shocks and find that these shocks mainly have a negative impact on equity prices and consumption.

Our study is distinct from these studies in that we employ regional textual sources

[2] Another strand of the literature uses the change in asset prices around geopolitical events, such as wars and terrorist attacks, as an instrument to identify geopolitical risk shocks (Piffer & Podstawski, 2018; Ha et al., 2022, among them). To the extent these events represent exogenous changes in risk, such a procedure allows for a clean identification of geopolitical risk shocks. Their results support the notion that higher geopolitical risk has a contractionary effect on economic activity.

rather than relying on U.S. newspapers. In this respect, this paper is closely related to the works of Jung et al. (2021) and Bondarenko et al. (2024), and Grebe et al. (2024), which use local text sources to construct indexes related to geopolitical risk. For example, Jung et al. (2021) construct a monthly index of geopolitical risk originating in North Korea from South Korean newspapers and find that escalated geopolitical tension on the Korean Peninsula reduces stock returns. Grebe et al. (2024) use data on Twitter posts in German to construct an uncertainty indicator about the war in Ukraine. Bondarenko et al. (2024) show that the geopolitical risk shocks identified from local news sources have more significant impacts on the local economy than those identified from U.S. news sources, using Russia as the case study. Similarly, we concentrate on East Asian newspapers and study the perceptions of the U.S. military presence from a regional perspective. Rather than searching the newspaper reporting for keywords directly related to risk or uncertainty, we search for terms related to the U.S. Navy.

Our study is also related to the existing literature on the economic impact of naval power projection. Deploying gunboats to far-away places in order to project the power of a hegemon is of course not a new phenomenon. The naval powers of the 19th and early 20th century employed "gunboat diplomacy" in order to enforce debt repayment or promote their military or commercial interests. The study of Glaser and Rahman (2016) is particularly relevant for us. The authors use archival data on the deployment of the British, French, German and U.S. navy, including data on ship personnel and tonnage. An estimated gravity model of bilateral trade shows that a country could bolster trade by deploying the navy. At the same time, trade of other nations suffered. Hence, trade globalization did not increase as a results of "gunboat diplomacy".

The remainder of this paper proceeds as follows. Section III.2 briefly sketches the geopolitical landscape in Asia. Section III.3 introduces our data and the empirical strategy. The results and several alternative model specifications are discussed in section III.4. Finally, section III.5 concludes.

## III.2 Geopolitics in East Asia

East Asia has been through severe confrontations since the Cold War, which continues to this day. Although geopolitical tensions in East Asia have not yet resulted in large-scale military action between countries, there is a risk that they could escalate to direct military conflicts and war at any moment. This means that East Asia is one of the regions with the highest geopolitical risk.

Figure III.1 depicts the geography of the region. Since China's rise in the 2000s, the strategic rivalry between the U.S. and China, including trade disputes and military confrontations, has become a central determinant of regional geopolitics in the post-Cold War era. China's assertive foreign policy, exemplified by its expansive territorial claims over Taiwan and in the South China Sea, has significantly increased regional geopolitical tensions. North Korea's continued provocations are also a primary source of conflict in the region. North Korea continues to be a destabilizing force through its persistent nuclear tests and missile launches, posing a direct threat to regional and global security.

While these risks primarily impact Northeast Asian countries (South Korea, Taiwan, and Japan), they also affect Southeast Asian countries such as the Philippines, Vietnam, Malaysia, and Indonesia. The dispute in the South China Sea directly involves some of these countries, creating substantial tensions. Additionally, the U.S.-China rivalry forces these countries to carefully balance security partnerships with the U.S. and economic interdependence with China.

Figure III.1: A map of East Asia


As the world's sole superpower, the U.S. plays a key role in the balance of power in the region with its naval forces. In particular, one CSG has been permanently located in Japan since 2008.[3] The U.S. allies (South Korea, Japan, the Philippines, and Taiwan) have also remained a cornerstone of regional security as it evolves to

[3]Of the eleven U.S. CSG, it is the only forward-deployed CSG with a home port outside the continental U.S.

counter the threats posed by both China's military modernization and North Korea's nuclear development.

The U.S. often responds directly to escalating geopolitical tensions or military activity in the region, such as North Korea's nuclear tests or China's expanding naval power, by deploying an additional CSG to the region. Such a naval operation could be a signal that the U.S. is ready to react to geopolitical tensions in the region. The U.S. may use a CSG's operation to demonstrate its persistent commitment to allies to maintain regional stability.

The existing literature on the economic consequences of geopolitical risk in East Asia has largely focused on North Korea's military provocations and the responses of South Korean economies (Kim & Jung, 2014; Gerlach & Yook, 2016; Kim et al., 2019; Park & Park, 2020; Jung et al., 2021; Ha et al., 2022, among them). These studies suggest that rising geopolitical tensions on the Korean Peninsula have a detrimental impact on the economy, primarily through increased uncertainty. On the other hand, the focus of the existing literature on the conflict between the U.S. and China is on the bilateral trade conflict. A limited body of research addresses the geopolitical conflict between the two countries, except for the work of Rogers et al. (2024). Using the U.S.-China tension (UCT) index they construct, Rogers et al. (2024) find that the heightened tensions between the two largest economies negatively affect the economy. This impact is attributed to the disruption of supply chains, reductions in firm investments, and a decline in the volume of bilateral trade.

## III.3 Empirical strategy

In this section, we first construct a measure of surprise change in the presence of CSG in the region. We then use this measure and estimate its impact on Asian financial markets.

### III.3.1 Measuring the presence of Carrier Strike Groups

To measure the presence of CSG in the region, we rely on East Asian English-language newspaper articles reporting. We extend and cross-check the analysis using data on the approximate locations of CSG provided by open-source data from the U.S. Navy.

We exploit the monthly number of news articles as a proxy for the market's attention to the presence of CSG. The newspaper articles are retrieved from `Nexis Uni`, an online research platform provided by Lexis Nexis. It offers access to a vast collection of news, business, legal, and academic information worldwide. `Nexis`

`Uni` gathers content from thousands of newspapers, magazines, news websites, and broadcast transcripts. We select the nine East Asian English-language newspapers that represent the local media: *China Daily*, *Global Times* (China), *Japan News*, *Korea Times*, *Korea Herald*, *Philippine Daily Inquirer*, *Straits Times* (Singapore), *South China Morning Post* (Hong Kong, China), and *Taiwan News.* We search for articles that contain U.S. CSG-related keywords and references to names of places in East Asia in the headlines or content of the news articles.

As previously stated, we are agnostic about the effects of an increased U.S. military presence. Therefore, we do not use other keywords such as "tension," "threat," or "conflict" when searching for newspaper articles. Appendix III.A.1 shows the search query we use to find relevant news articles. For each newspaper, we normalize the number of articles with a unit standard deviation from January 2010 to October 2024,

$$\tilde{x}_{i,t} = \frac{x_{i,t}}{\sigma_i}, \tag{III.1}$$

where $x_{i,t}$ is the monthly count of articles on CSG of newspaper $i$ at time $t$ and $\sigma_i$ is the standard deviation of $x_{i,t}$ over the entire sample period. To mitigate the potential distorting effect of outliers from a single newspaper at a given point in time, we winsorize the number of articles for each newspaper, $x_{i,t}$, at the 99th percentile. We then average the standardized series across newspapers

$$News_t = \frac{1}{N}\sum_{i=1}^{N} \tilde{x}_{i,t}, \tag{III.2}$$

where $N$ is the number of newspapers.

We collect a total of 1,677 articles that mention the presence of the U.S. CSG during the sample period in the manner described above. Compared to other studies that have employed textual analysis to identify policy uncertainty (Baker et al., 2016) or geopolitical risks (Caldara & Iacoviello, 2022), the number of collected articles in our paper is quite limited. This is because our analysis focuses on specific issues, namely the presence of the U.S. CSG, rather than the overall geopolitical landscape of the region. Therefore, it is necessary to assess whether the constructed variable adequately captures the actual presence of CSG and the market's attention to them.

To this end, we also use information on the actual presence of CSG from the USNI (United States Naval Institute) News in order to cross-check the newspaper-based data.[4] The USNI News has reported the approximate locations of deployed CSG on a weekly basis since July 2017. To distinguish the locations of CSG that have a meaningful regional impact from other locations further out in the Pacific, we adopt the concept of *island chain*, which is a key geographical security concept used by the U.S. and China. The origin of this concept is known as the U.S. strategic concept during the early Cold War. For the U.S., the island chain is a line of defense against Communist forces at sea (now the rise of China). For China, it is a line that needs to be broken by military force.[5] Officials distinguish between the first, second, and third island chain, though the definition of the first island chain is most appropriate for our purpose. The exact definition of the first island chain varies in the literature. However, it is generally accepted that it refers to a "long band of islands centered primarily on the Japanese home islands, the Ryukyus, Taiwan, and the Philippines" (Yoshihara, 2012, p294).

Figure (III.A.1) in the Appendix provides geographic information about the first and second island chains. The inner part of the first island chain includes the waters around the Korean Peninsula, the East China Sea, and the South China Sea. We assume that the presence of a CSG is only significant to the region when it is located in this area. If a CSG is located inside the first island chain, we assign a value of one to the variable. Also, if a CSG is located inside a second island chain (usually in the Philippine Sea), we give it half the value. Since one CSG is permanently located in East Asia, we only consider the deployments of additional CSG.[6]

Figure (III.2) shows the relationship between two CSG indicators on a monthly basis from July 2017 to October 2024. The monthly location of CSG is calculated as a simple average of the weekly series. The two indicators demonstrate comparable movements over the sample period, thereby confirming the validity of the measure constructed using the count of articles. Nevertheless, we also spot periods of divergence from each other at certain points in time. The correlation between the two measures is 0.26. We also compare the two indicators at the weekly frequency. This is helpful below when we differentiate between before, during and after the first Trump administration. Based on monthly data, the three subsamples would not have enough observations. As shown in Figure (III.A.2) in the Appendix, the movements of the two weekly series appear similar, as observed in the monthly series. However, they also diverge from each other at specific points in time. The

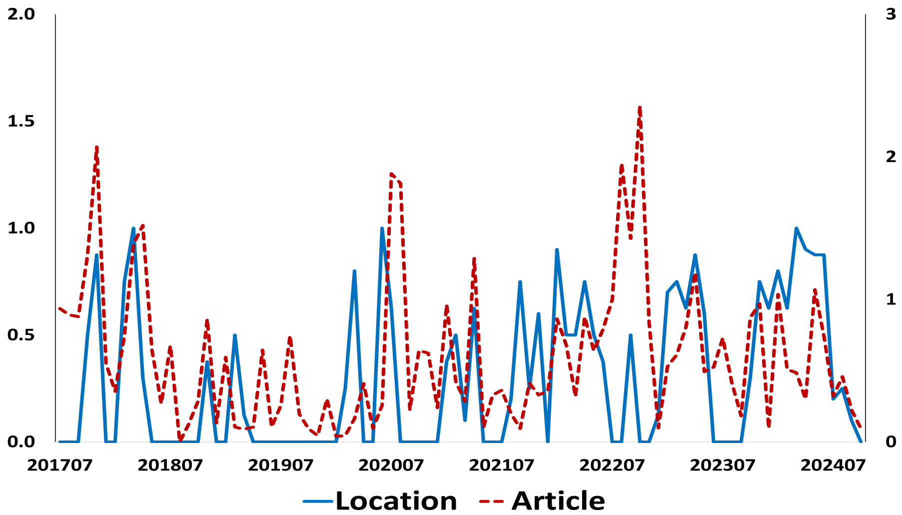
[4]The USNI is a private non-profit military association that offers independent, nonpartisan forums for debate on national security issues, see `https://news.usni.org/category/fleet-tracker`.

[5]For a comprehensive discussion of the island chain, see Yoshihara (2012).

[6]More specifically, the Ronald Reagan CSG (CVN 76) has been based in the Yokosuka Port (Japan) since 2015.

correlation between the two weekly series is 0.30.

Figure III.2: Comparison of CSG indicators




*Notes*: A unit in the location of CSG (left-hand scale) is one additional deployment of CSG in the region, and a unit in the number of articles on CSG (right-hand scale) is one standard deviation, as discussed in the text.

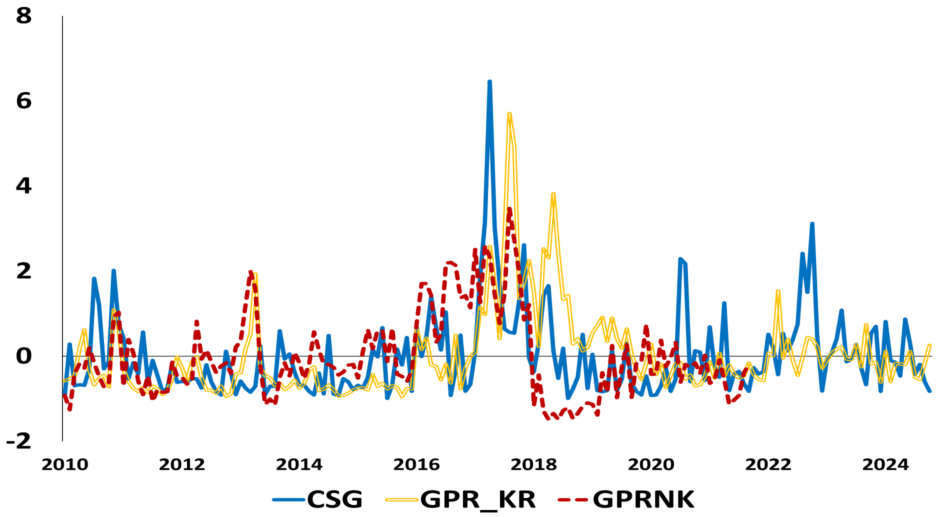
The observed lower-than-expected correlation between the two series may be attributed to the following explanations. First, the information on the approximate location of CSG contains noise, such as simple transit of CSG to other waters or their routine operations. Such information is of limited value to the market. Second, newspapers also report when CSG begin to move toward the region. While we cannot precisely determine the final destination or objective of the movement until CSG arrive, such information is sometimes publicly announced by the U.S. Navy or can be reasonably inferred. These cases should be reported by newspapers, which are reflected in our index. We acknowledge that the expectation of their arrival also affects the market, not only the actual arrival of CSG. Consequently, the index constructed using newspapers is more informative than location data. In addition, while the number of articles reflects the level of the market's attention to the presence of the U.S. CSG, location data does not. For these reasons, we employ the count of articles as a primary indicator rather than the location data to examine the impacts of the increased presence of U.S. CSG.

Figure (III.3) displays our CSG indicator and other indices of geopolitical risk related to Asia, i.e. the GPR index for South Korea, Taiwan and China (Caldara & Iacoviello, 2022) and the GPRNK index developed by Jung et al. (2021). As shown in the figure, the presence of CSG is mainly associated with North Korean provocations before 2019. Since 2019, it appears to be related to the conflicts and military tensions between China and Taiwan. Specifically, the South Korea GPR

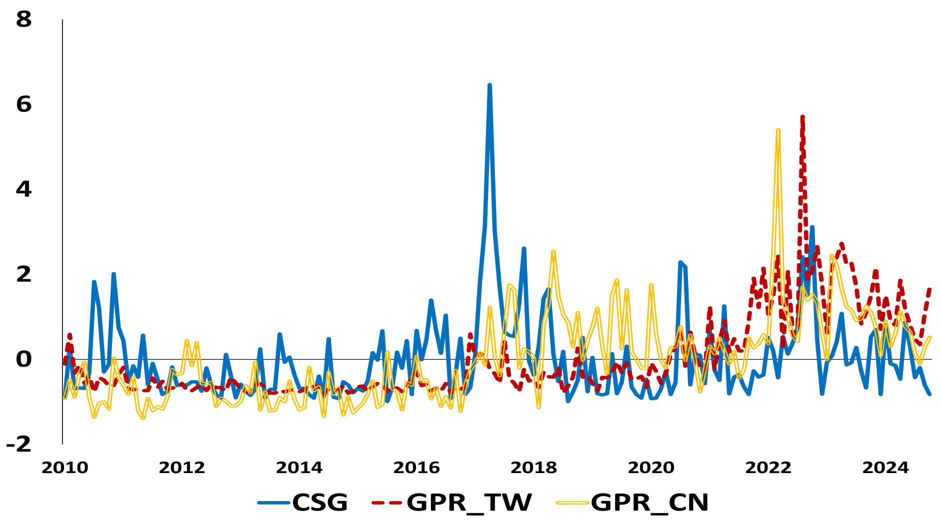
Index has a high correlation of 0.43 with the CSG Index over the period from 2010 to 2018. However, this correlation declines to 0.06 in the period after 2019. Conversely, the Taiwan GPR Index exhibits a high correlation of 0.47 with the CSG index in the period after 2019, but a relatively low correlation of 0.28 before that.

Figure III.3: Comparing the CSG Index to related indices

(a) CSG and Korea-related indices




(b) CSG and Taiwan-related indices

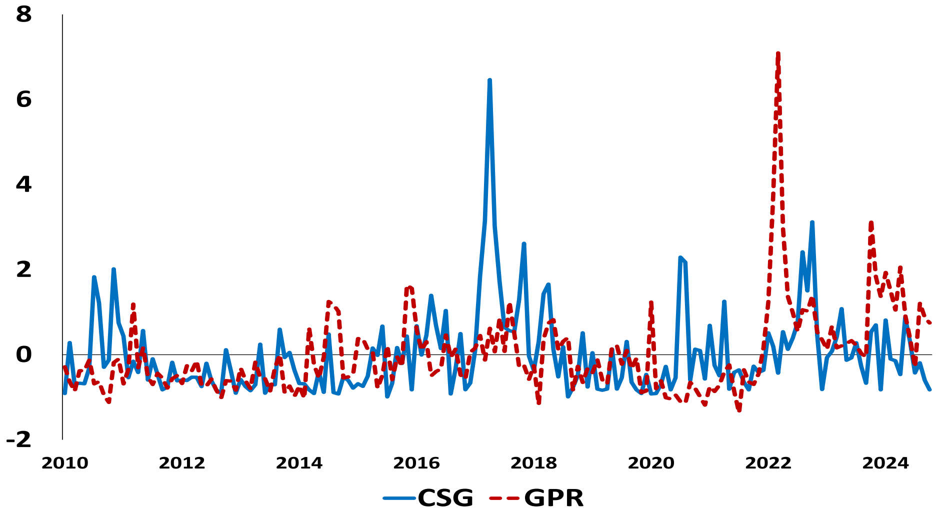



*Note*: For comparability, all series are normalized so that the mean is equal to zero and the standard deviation is equal to one. The construction of the CSG index is explained in the text. GPR indices are from Caldara and Iacoviello (2022) and the GPRNK index is from Jung et al. (2021).

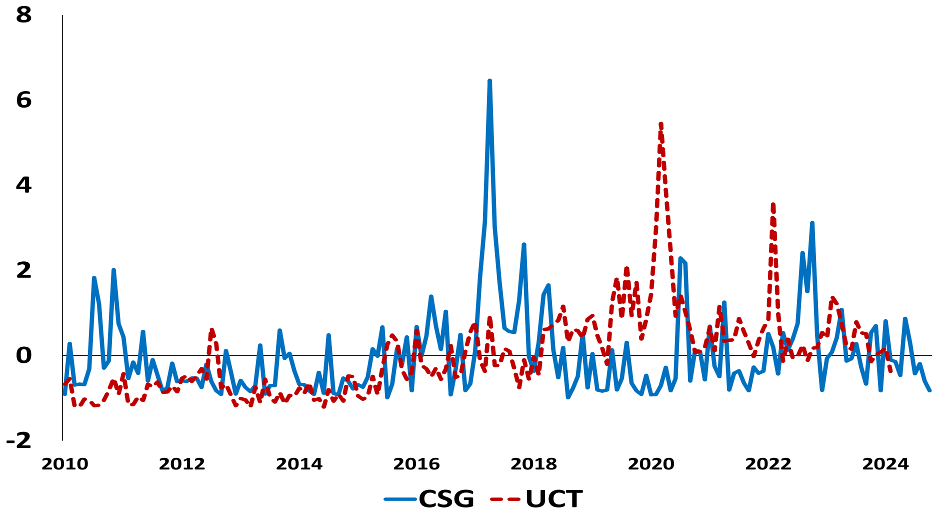
Figure (III.4) depicts the relationship between the CSG index and the global GPR index, as well as with the UCT index constructed by Rogers et al. (2024). The CSG index is significantly associated with the global GPR index, with a correlation coefficient of 0.17. However, this correlation is less pronounced than the association with country-specific GPR indexes. On the other hand, there is no significant correlation between the CSG and the UCT index (correlation coefficient = 0.05), as shown in panel (b). While the UCT index directly addresses the tensions between the U.S. and China, it appears to be less correlated with military actions because it comprehensively captures multiple categories of tensions, such as trade and human rights. This indirectly confirms that our index is not influenced by other geopolitical issues unrelated to the actual presence of CSG.

Figure III.4: The CSG index and the other indices

(a) CSG and global GPR




(b) CSG and UCT




*Note*: For comparability, all series are normalized so that the mean is equal to zero and the standard deviation is equal to one. The CSG index is explained in the text. The global GPR index is from Caldara and Iacoviello (2022) and the UCT index is from Rogers et al. (2024).

### III.3.2 Construction of shock series

For our estimation below, we need a measure of the unexpected change in the military presence. This is because the expected presence should already be fully incorporated into asset prices. We define the unexpected presence of CSG as shocks that occur i) when CSG begin unexpectedly moving into the region, ii) when CSG actually enter and operate in the region (which was already expected, but remained uncertain), or iii) when CSG conduct an irregular operation after entering, such as a large-scale military exercise.[7] Changes in the CSG index reflect both expected and unexpected news. Therefore, we need to purify this series from expected changes before we use it in our regression model.

We regress the CSG index at time $t$, $csg_t$, on one lag of itself and the log of the GPR index via simple ordinary least squares (OLS),

$$CSG_t = \alpha + \beta CSG_{t-1} + \gamma \log(GPR)_t + \varepsilon_t \tag{III.3}$$

where the residual $\varepsilon_t$ should reflect the unexpected component of the CSG index. To check the properties of the residual $\varepsilon_t$, we conduct two tests for serial correlation: the Durbin–Watson test and the Breusch-Godfrey LM test. Neither test rejects the null hypothesis of no serial correlation.

Figure (III.5) shows the resulting shock series. We can clearly spot historical episodes of heightened tensions such as North Korean provocations, China's military exercises and the "We are sending an armada"-statement of President Trump in April 2017.

### III.3.3 Empirical model

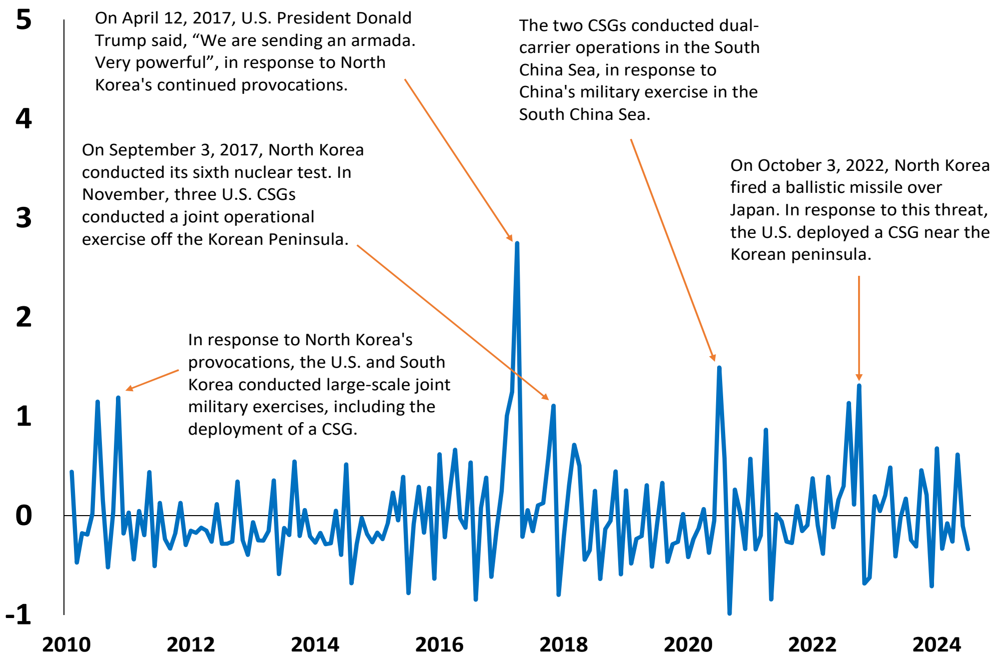
Based on the constructed shocks, we estimate the effect of the surprise change in the presence of CSG on East Asian financial markets using a panel for seven East Asian countries (South Korea, Taiwan, Japan, the Philippines, Vietnam, Malaysia, and Indonesia). The sample period is from January 2010 to October 2024. We estimate impulse response functions using local projections proposed by Jordà (2005) with panel regressions. The regression equation is

$$y_{i,t+h} = \beta_h \varepsilon_t + \mathbf{\Gamma}_h \mathbf{X}_{i,t} + \tau_i^h + \upsilon_{i,t+h} \tag{III.4}$$

for a time horizon $h = 0, 1, ..., H$. The indices $i$ and $t$ denote country and year, respectively. The dependent variable, $y_{i,t}$, is driven by $\varepsilon_t$, which is our shock series,

[7] Unfortunately, our index cannot sufficiently distinguish between these cases.

Figure III.5: Estimated shocks of the CSG index




*Notes*: The estimated shock series is shown in units of standard deviations around its mean.

i.e., the residual from the auxiliary model estimated in Section III.3.2. It reflects a surprise increase in the presence of CSG in the region. The parameter of interest is $\beta_h$, which represents the effect of a shock on the dependent variables after $h$ periods. The sequence of this estimated parameter as a function of $h$ provides the impulse response function. $\tau_i$ is the country-fixed effect, which we include to control for unobserved cross-country heterogeneity, and $\upsilon_{i,t+h}$ is the projection residual.

We employ the following five dependent variables: the stock market index,[8] the nominal exchange rate against the U.S. dollar, the 10-year government bond yield, the size of portfolio equity inflows, and the country-specific GPR index. The stock market index and the exchange rate are included in logs (multiplied by 100). An increase in the exchange rate is a depreciation of the domestic currency. We use fund flow data from EPFR as a high-frequency proxy for portfolio equity inflows.[9] We also include the log of the country-specific GPR index as a dependent variable

[8]The following stock market indexes are employed for the respective countries: KOSPI for South Korea, TAIEX for Taiwan, Nikkei 225 for Japan, PSEi for the Philippines, Ho Chi Minh Stock Index for Vietnam, KLCI for Malaysia, and JSX Composite Index for Indonesia.

[9]The EPFR dataset is the most widely used fund flow dataset in the literature. The series normalized as a percentage of total net assets at the beginning of the period. The dataset provides monthly, weekly, and daily estimates of fund flows and allocations to specific countries and sectors. For the monthly data, reporting equity funds cover approximately 93% of the assets under management of all global mutual funds. Furthermore, the fund flows in the EPFR dataset are free of valuation effects as they remove portfolio changes due to asset returns and exchange rate changes.

to examine the changes in the perceptions of geopolitical risks at the country level in response to an increased U.S. military presence. It is worth noting that the country-specific GPR indices constructed by Caldara and Iacoviello (2022) may reflect the U.S.-centered perspective on the geopolitical risks posed by a given country, as discussed in Bondarenko et al. (2024) and Choi and Havel (2025). To the extend the presence of CSG calms tensions, we expect a surprise increase in the presence of CSG to cause an increase in stock prices, an appreciation of the domestic currencies, a drop in bond yields, and an increase in capital inflows. Consequently, the GPR index should fall.

$\mathbf{X}_{i,t}$ is a vector of control variables including the lagged values of dependent variables and the following exogenous variables: the log of the CBOE Volatility index (VIX), the U.S. 10-year Treasury bond yield, the log of the S&P 500 stock market index, the log of the GPR index and a dummy variable that takes the value one from February 2020 to June 2020 and 0 in the rest of the sample periods. This dummy captures the extraordinary movements of the variables during the Covid-19 pandemic. We also include two domestic variables, the year-on-year percentage changes in the Industrial Production and the Consumer Price Index. Definitions and sources of data for all variables are presented in Table (III.A.1) in the Appendix.

We estimate Equation (III.4) using the OLS estimator. The presence of the lagged dependent variables and individual-fixed effects in the panel model can generate bias in the coefficients estimated by OLS (Nickell, 1981). However, the small $N$ and large $T$ setting ($N$=7, $T$=178) in our dataset mitigates this concern.

# III.4 Results

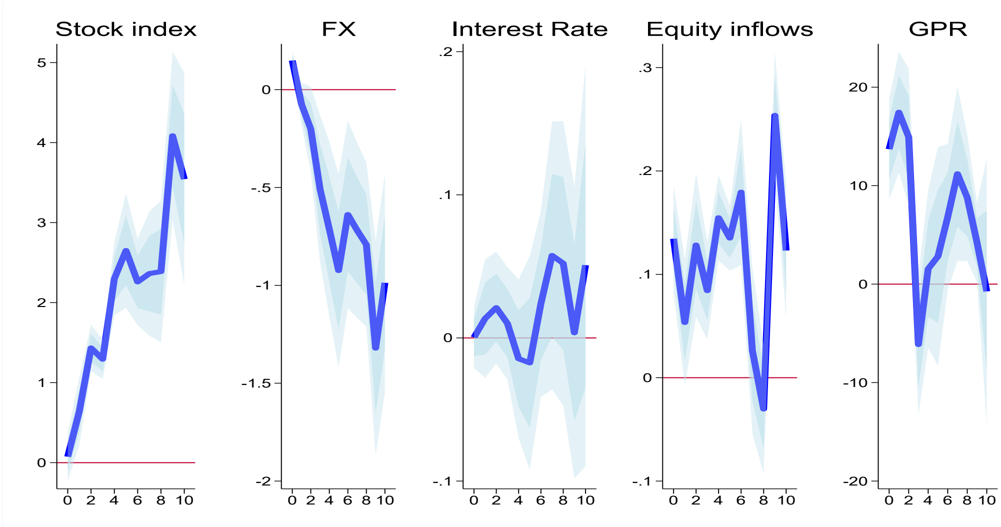
## III.4.1 Main results

In this section, we present our main findings based on the monthly panel regression. The figures show the estimated $\beta_h$ parameters as well as 68% and 90% confidence bands. We cluster standard errors at the country level to control the within-country correlation across time.

An unexpected increase in the presence of CSG has strong and significant effects on East Asian financial markets, see Figure (III.6). Following the shock, the valuation of the stock market increases significantly and steadily, with the increase reaching a peak of about four percent. Furthermore, the local currencies appreciate by more than one percent against the dollar. While equity inflows show high volatility, we also find a significant increase in the volume of equity inflows. On the other hand, the long-term government bond rates do not respond significantly.

Overall, the deployment of the Navy seems to support confidence in the region and has an expansionary effect on financial markets. The country-specific GPR index also tends to increase. Our interpretation is that the deployment of the navy causes a more intense coverage of the region in U.S. newspapers using words listed in the dictionary of Caldara and Iacoviello (2022). As a result, the GPR index increases. Hence, we interpret this in terms of additional newspaper coverage rather than a material increase in objective risk.[10]

Figure III.6: Responses to military presence




*Notes*: The solid blue line is the point estimate of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

In order to more plausibly capture the heterogeneity in the effects of CSG shocks on the financial markets between Northeast Asia and Southeast Asia, we divide the sample countries into two groups: i) Northeast Asian countries, namely South Korea, Taiwan, and Japan, and ii) Southeast Asian countries, namely the Philippines, Vietnam, Malaysia, and Indonesia. We estimate the impulse response functions of each country group using the same panel regression model as in Equation (III.4).

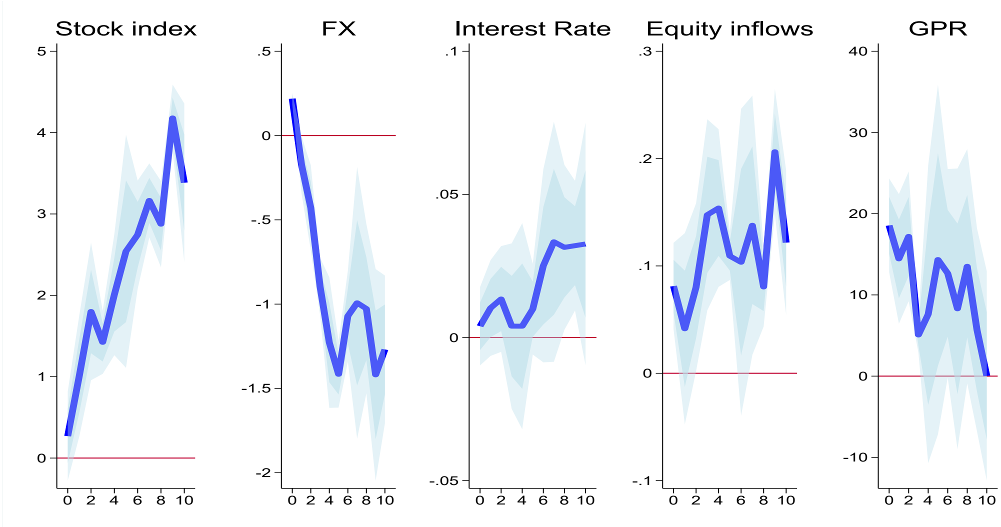
For Northeast Asian countries (group 1), Figure (III.7) shows very similar results to market responses observed for all sample countries. Again, stock prices increase, the national currency appreciates, and equity inflows increase. Long-term bond yields tend to rise, though this response is not statistically significant.[11] Figure (III.8) presents the results for Southeast Asian countries (group 2). While stock

[10]As noted by Bondarenko et al. (2024), the GPR index tends to reflect the U.S.- (or Western) centered perspective.

[11]The rise in interest rates may be attributed to a decline in demand for safe assets as uncertainty diminishes. Alternatively, current interest rates may face upward pressure due to expectations of future inflation and growth driven by the expansionary effects of CSG.

prices also respond strongly and increase by about four percent, the responses of other variables are not as clear as before. The local currency appreciates, but strong statistical significance does not emerge until eight months after the shock. Equity inflows exhibit a high degree of volatility. Overall, the market response to the CSG shock is more pronounced in Group 1. We consider these results highly plausible since Northeast Asian countries are more directly related to the geopolitical tensions associated with the deployment of CSG in the region, as stated in Section III.2.

Figure III.7: Responses to military presence (Group 1)



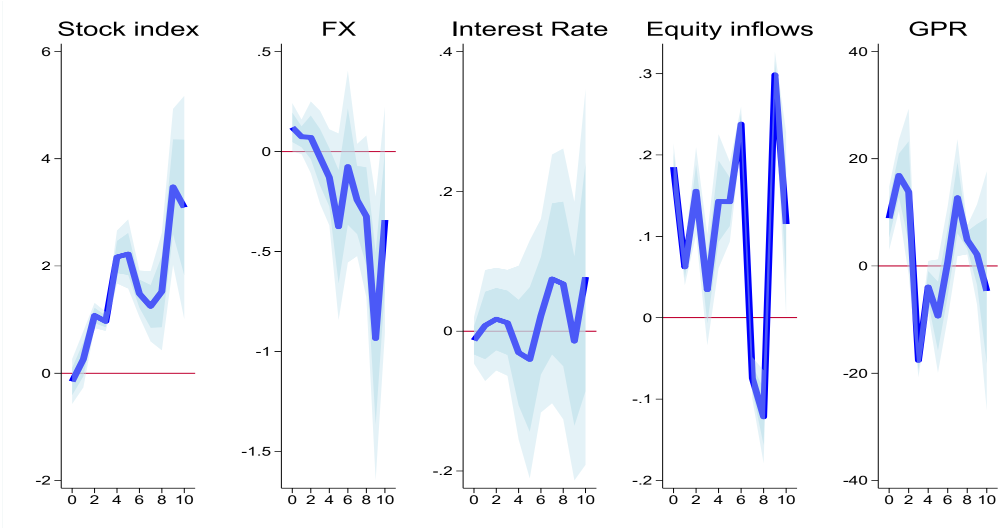

*Notes*: The solid blue line is the point estimate of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

### III.4.2 Decomposition of shocks related to North Korea and China

As discussed in Section III.2, the geopolitical tensions in East Asia can primarily be attributed to two sources: the first is North Korea's military provocations, which include nuclear tests and ballistic missile launches. The second is China's expansive territorial claims over Taiwan and in the South China Sea. The aforementioned search query and resulting variable may capture the full range of U.S. military responses, represented by the deployment of CSG, to both sources of escalating regional geopolitical risks.

We now differentiate between these two sources of risk and construct two components of the CSG index, one pertaining to North Korea (hereafter CSG-NK) and the other to China (hereafter CSG-China). In order to construct the CSG-NK, we retrieve articles using only the names of places associated with North Korea (e.g., the Korean Peninsula, Seoul, and Pyongyang) among the names of regions in our

Figure III.8: Responses to military presence (Group 2)




*Notes*: The solid blue line is the point estimate of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

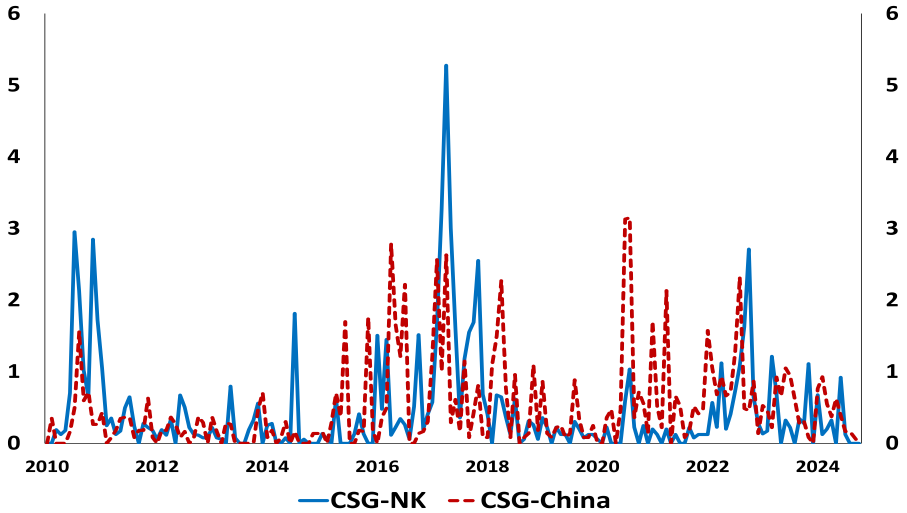
search query presented in the Appendix. Similarly, to construct the CSG-China, we use only the names of places associated with China (e.g., Beijing, Taipei, and the South China Sea). It is worth noting that the deployment of CSG is sometimes associated with tensions related to North Korea and China simultaneously due to the proximity of the two countries' geographical locations and the intertwined nature of their foreign and securities policies. For this reason, it is not easy to divide the CSG index into two completely distinct and mutually exclusive subcategories. Figure (III.9) plots the CSG-NK and the CSG-China over the entire sample period. The two series have a correlation of 0.29.

Figures (III.10) to (III.12) show the estimated impulse responses of financial variables to the two shocks.[12] The shock to the CSG-NK is strongly associated with a rise in the stock price, an appreciation of the local currency, and increased equity flows. These results are broadly similar to the baseline results but more significant. In contrast, the shock to the CSG-China appears to result in a depreciation of the local currency, particularly for Southeast Asian countries. East Asian stock markets experience an initial decline in value, and the response of capital inflows seems to be highly volatile.

These results suggest that the deployment of U.S. CSG in response to North Korean provocations can deter further provocations and mitigate investors' concerns about heightened tensions. Conversely, the operations of CSG in the context of

[12]The shocks to the CSG-NK and CSG-China are identified in a manner analogous to equation (III.3), but are regressed on the relevant country-specific GPR rather than the global GPR.

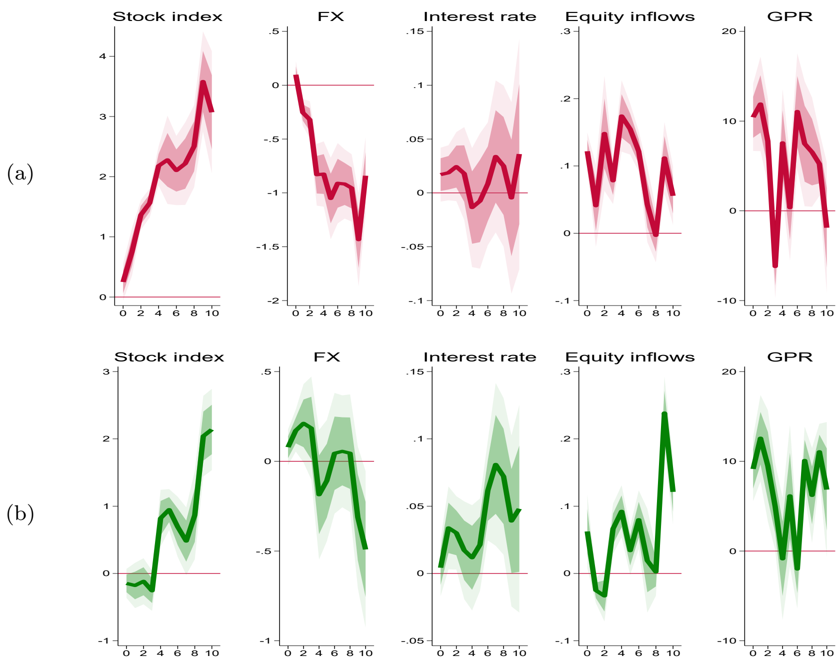
Figure III.9: The CSG-NK and the CSG-China indices




*Notes*: Both series are shown in standard deviations.

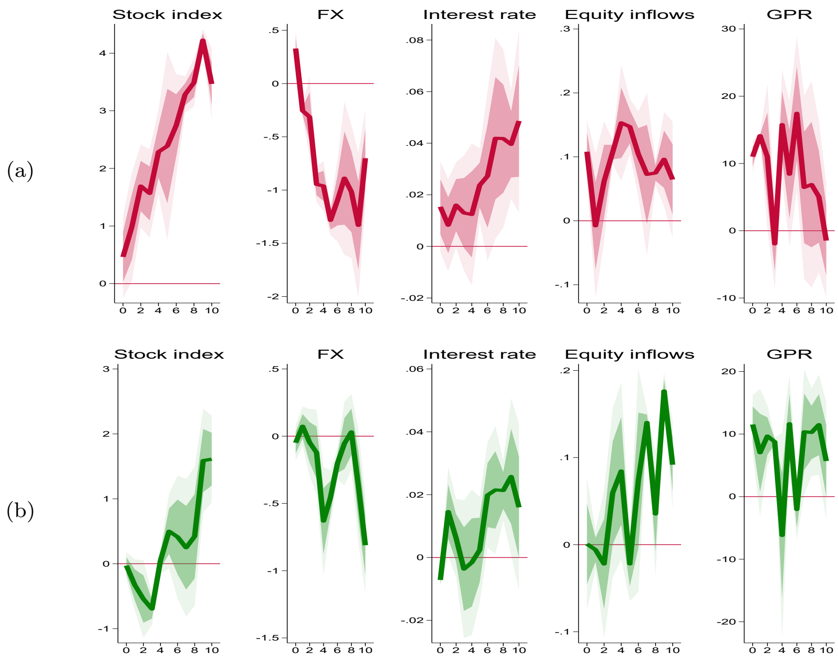
tensions with China may not have these effects or may even escalate tensions. As mentioned above, however, the two shocks are neither orthogonal nor additive. That is, the aggregate effect of two sub-indices does not equal the overall impact of the CSG index. For this reason, we should interpret the results with caution.

Figure III.10: Responses to military presence (all sample countries)



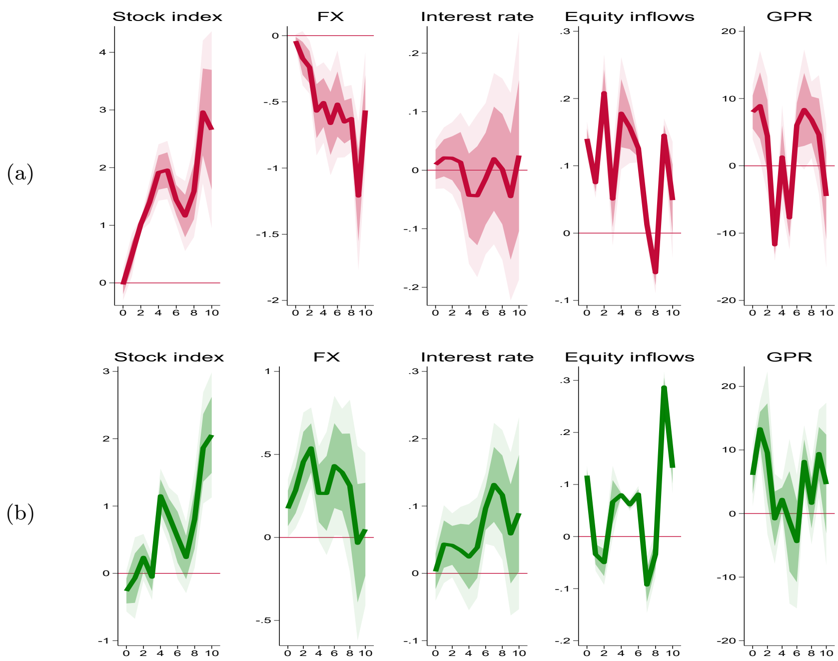

*Notes*: Panel (a) shows the effects of shocks to CSG-NK, and panel (b) shows the effects of shocks to CSG-China. The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

Figure III.11: Responses to military presence (Group 1)




*Notes*: Panel (a) shows the effects of shocks to CSG-NK, and panel (b) shows the effects of shocks to CSG-China. The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

Figure III.12: Responses to military presence (Group 2)




*Notes*: Panel (a) shows the effects of shocks to CSG-NK, and panel (b) shows the effects of shocks to CSG-China. The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

### III.4.3 The role of the Trump administration

We now analyze the responses at a weekly frequency. We simply aggregate the count of newspaper articles reporting at the weekly frequency and use weekly financial data series to run the regressions. The use of a higher frequency allows us to better understand short-term responses, which are washed-out at a lower frequency. Furthermore, the use of weekly data with a larger number of observations helps us to study whether the responses to the CSG index on financial markets differ across U.S. administrations. We divide the overall sample periods into three distinct periods: before the first Trump administration, during the first Trump administration, and after the first Trump administration.

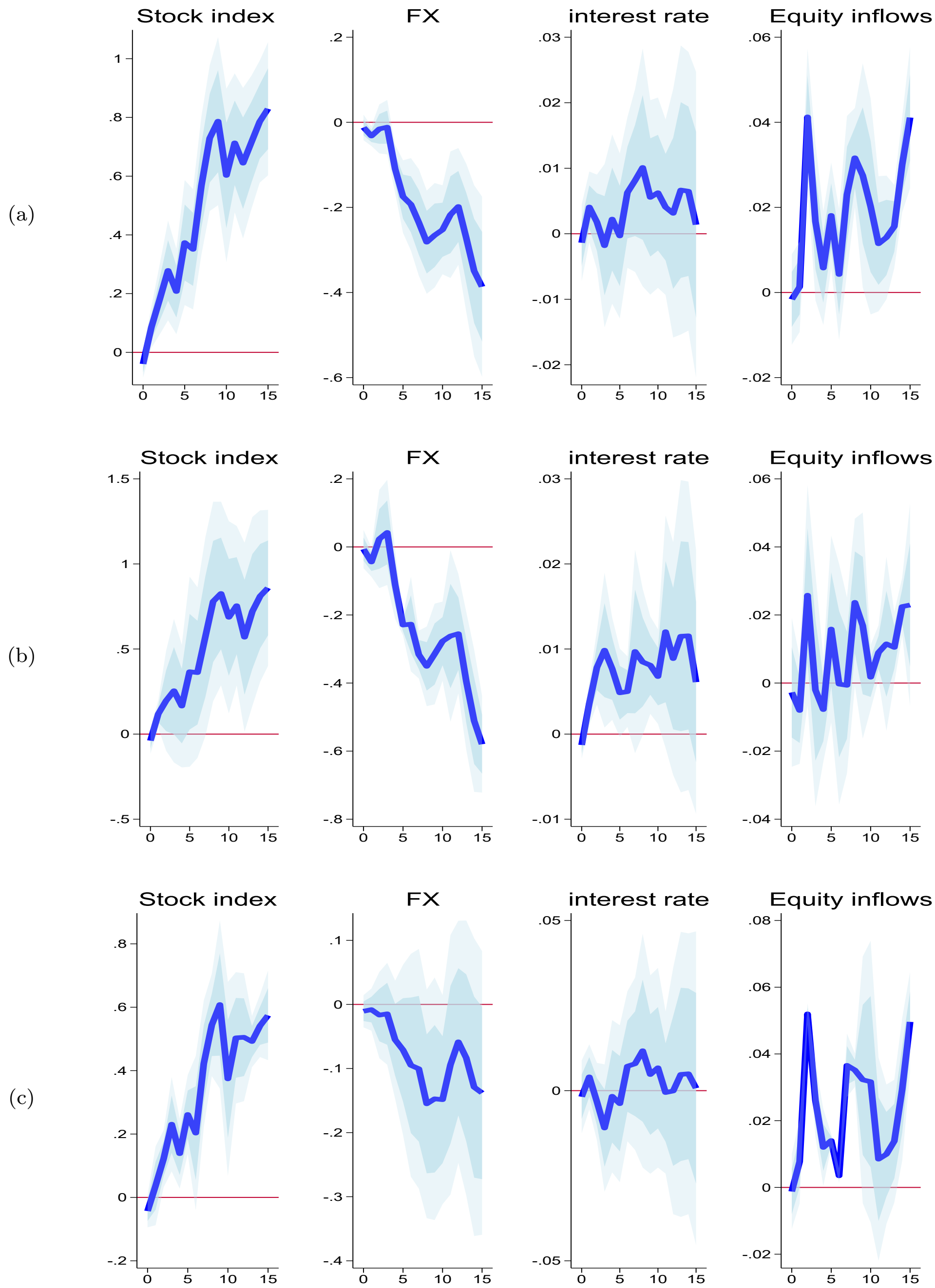
First, we present the results of the weekly panel regression over the entire sample period. The shocks to the weekly CSG index are identified in the same manner as in the monthly regression. In the absence of weekly country-specific GPR indices, we include only the four financial variables in the regression as dependent variables. Figure (III.13) presents the results. The results are fully in line with our baseline findings in Section III.4.1: a positive shock causes higher stock prices and a nominal appreciation. The increase in interest rates is less significant, and the responses of inflows remain inconclusive.[13]

Second, we show the responses before, during, and after the first Trump administration. The subsample index is constructed by standardizing the number of articles in each sample period, respectively, to facilitate comparing the responses of financial variables across time periods. Figures (III.14)-(III.16) present the results of weekly regression by subsample period for all sample countries, Northeast Asian countries, and Southeast Asian countries, respectively. We find that the impact of increased CSG presence on East Asian financial markets varies across the U.S. administrations. Before the first Trump administration, namely during the Obama Administration, the unexpected increase in the presence of CSG leads to an increase in stock prices and an appreciation of local currency, consistent with the results for the entire sample period. In contrast, during the first Trump administration, the projection of U.S. naval power lowers stock prices and capital inflows. The effect on exchange rates is somewhat ambiguous, but unlike in previous periods, it also leads to currency depreciation at certain forecast horizons. In the periods after the first Trump administration, the impact of shocks to the CSG index on financial markets is less clear. Given that developments of CSG during this period are highly corre-

[13]The weekly flow data from EPFR has less coverage than the monthly data. Not all funds that report on a monthly basis provide information on fund allocations and flows on a weekly basis. Consequently, the response of equity flows in weekly regressions may not be consistent with the response in monthly regressions.

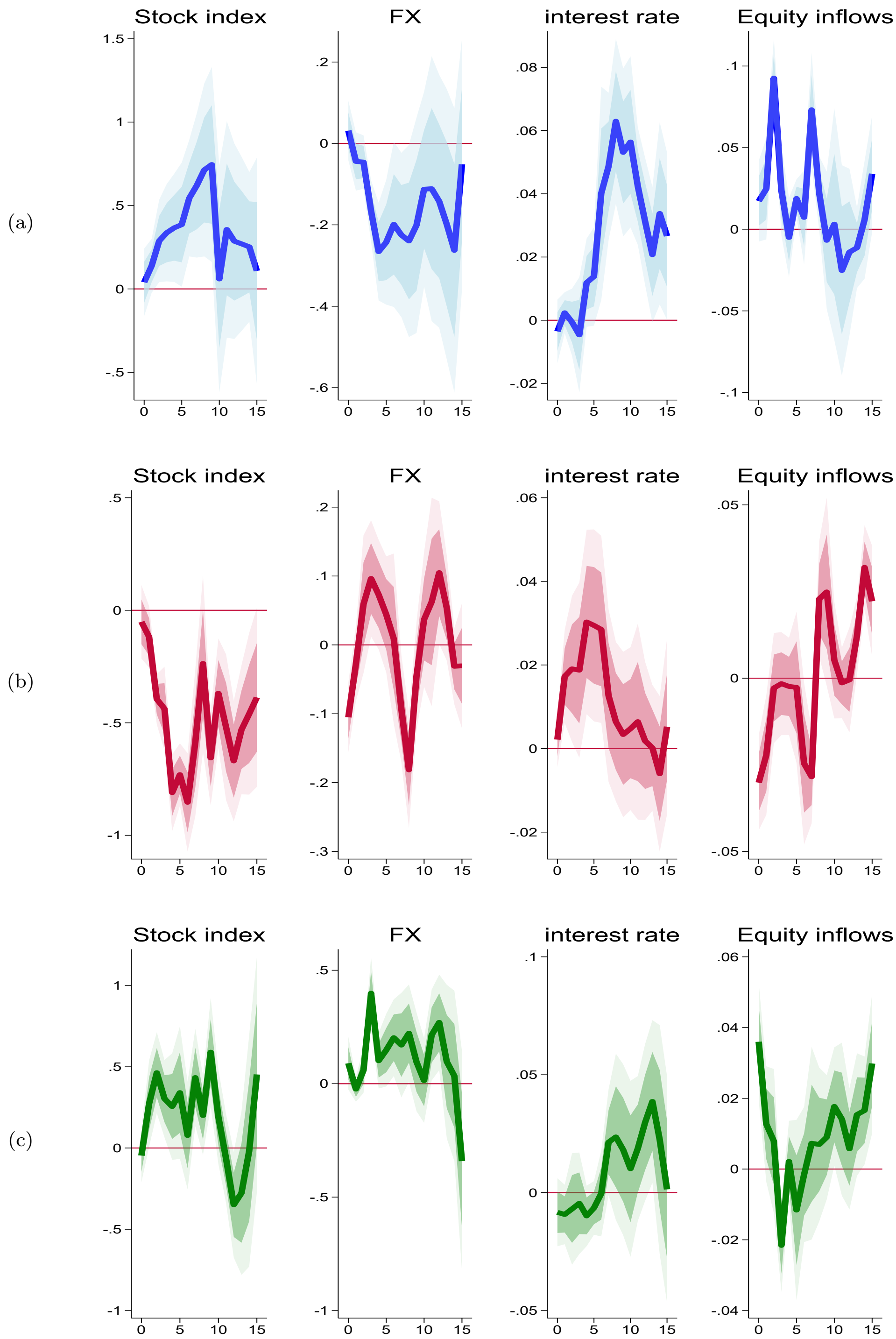
lated with the conflict over Taiwan, this result is largely consistent with the results obtained by decomposing the CSG index into two components in Section III.4.2.

Figure III.13: Responses to military presence (weekly regression)



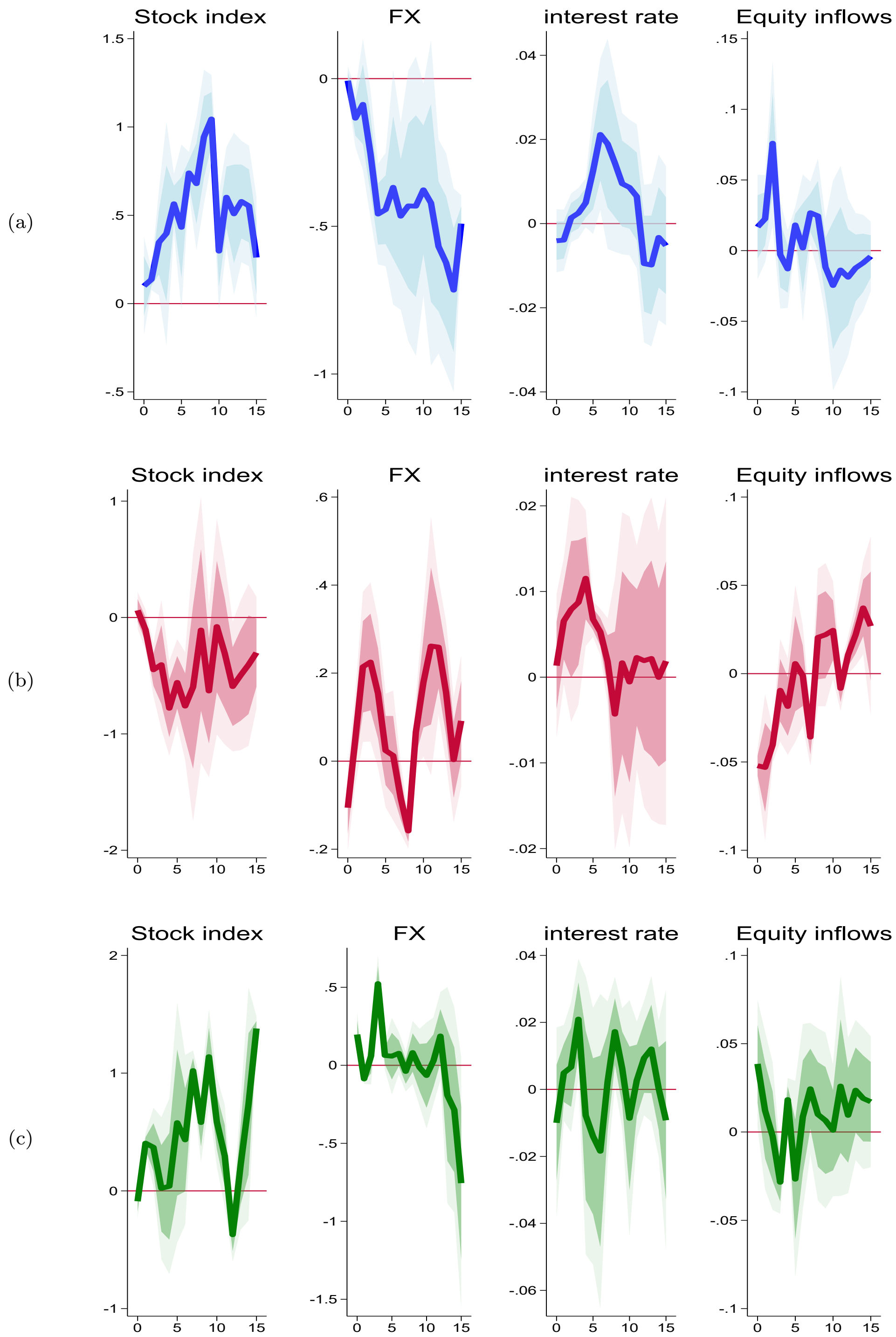

*Notes*: The three panels represent the following country groups: (a) all sample countries, (b) group 1 (South Korea, Taiwan, and Japan), and (c) group 2 (the Philippines, Vietnam, Malaysia, and Indonesia). The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

Figure III.14: Responses to military presence (all sample countries, divided periods)



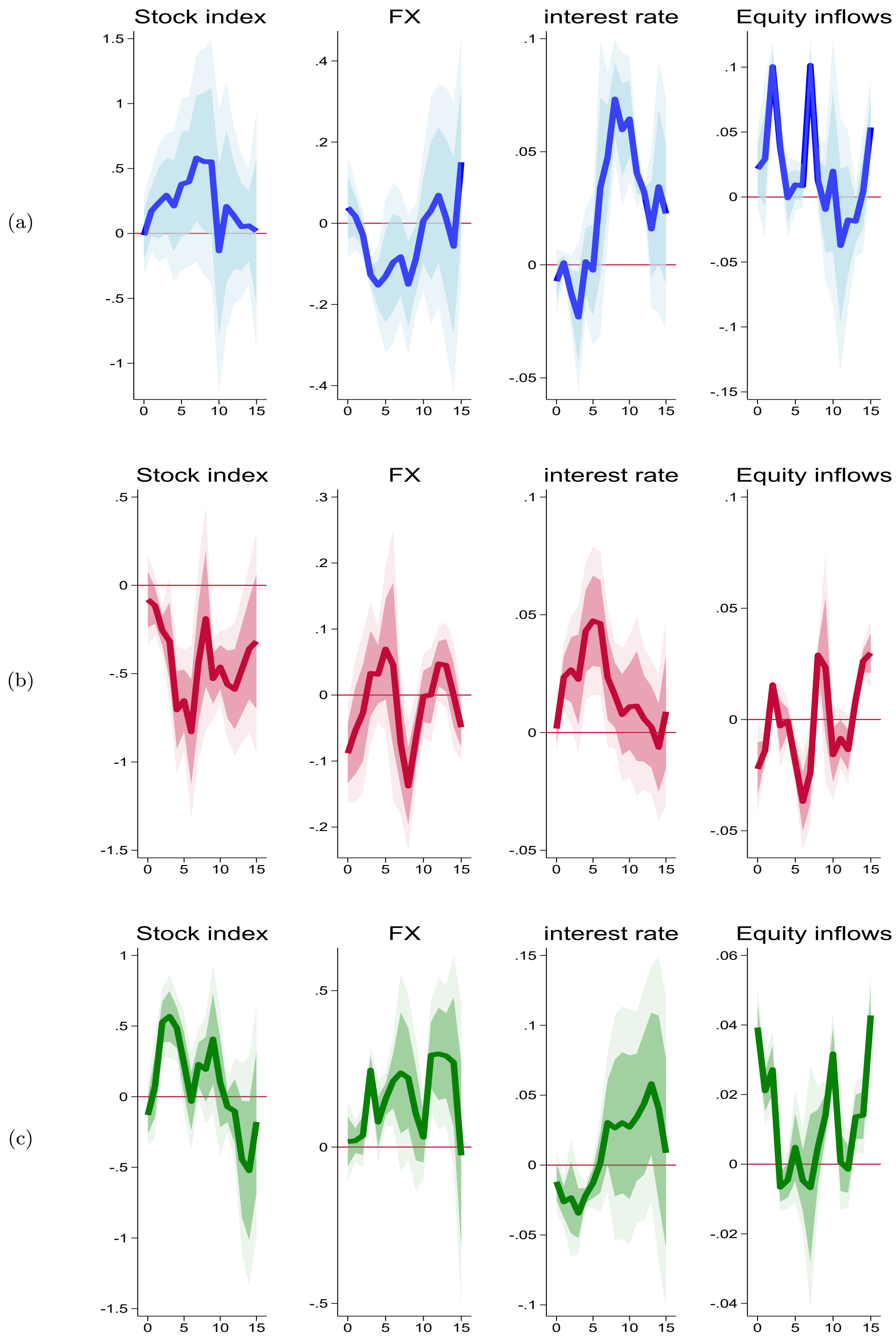

*Notes*: The three panels represent the following time periods: (a) before the first Trump administration, (b) during the first Trump administration, and (c) after the first Trump administration. The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

Figure III.15: Responses to military presence (Group 1, divided periods)




*Notes*: The three panels represent the following time periods: (a) before the first Trump administration, (b) during the first Trump administration, and (c) after the first Trump administration. The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

Figure III.16: Responses to military presence (Group 2, divided periods)




*Notes*: The three panels represent the following time periods: (a) before the first Trump administration, (b) during the first Trump administration, and (c) after the first Trump administration. The solid lines are the point estimates of the $\beta_h$ coefficient reflecting the impact of a surprise increase in the presence of CSG in the region. The shaded areas are 68% and 90% confidence bands constructed from clustered standard errors.

## III.5 Conclusion

For years, geopolitical tensions in East Asia have been high. These tensions stem from North Korean provocations and China's expansive territorial claims on Taiwan and in the South China Sea. In the background lies the unfolding rivalry between China and the U.S. about economic, military, and political dominance in the region. Often, the U.S. responds to heightened disputes by deploying aircraft carriers, or Carrier Strike Groups, to the region. In this paper, we studied the effects of news about a change in the presence of CSG on East Asian financial markets.

We have constructed a novel index derived from reporting on the U.S. Navy's CSG in East Asia in regional English-language newspapers. We show that an unexpectedly higher CSG presence leads to a significant appreciation of the stock market valuation, an appreciation of the nominal exchange rate to the U.S. dollar, and higher capital inflows. Overall, we find that an increased presence of the U.S. Navy is positive news for financial markets. In this sense, there are financial gains from gunboats, and news about the U.S. deploying additional naval resources generates expansionary effects. This finding suggests that an additional CSG deployment contains tensions in the region, contributes to stability, and reduces the uncertainty of investors. In contrast to the 19th century "gunboat diplomacy", which served the geopolitical ambitions of the hegemonic powers, these gains accrue not on U.S. financial markets, but on financial markets abroad.

One potential policy implication pertains to the financial burden-sharing between the U.S. and its allies regarding the deployment of CSG. If news about an increased military presence boosts financial markets, it could be argued that the allied countries should shoulder a larger part of the associated financial burden. However, this argument oversimplifies the complicated market impact of the deployment of CSG. Our findings reveal that the financial market benefits vary significantly depending on geopolitical contexts and across U.S. administrations.

The deployment of the U.S. CSG in response to North Korean provocations has a positive effect on financial markets by deterring further provocations and alleviating investors' concerns about heightened tensions. Conversely, the operations of CSG in the context of tensions with China may not have these effects or might even escalate existing tensions. These results suggest that the effects of an increase in CSG presence on financial markets vary depending on the geopolitical contexts surrounding the deployment of CSG. Furthermore, the CSG index may encompass information about the changes in U.S. strategic responses following the deployment of CSG. Such changes may involve various military and/or diplomatic pressures or negotiations. Our finding that the financial market responses vary across different

U.S. administrations implies that the market effects can be closely tied to overall U.S. strategic policy frameworks rather than merely reflecting the physical presence of CSG. These findings underscore the complexity of the interrelationship between military deployments, geopolitical tensions, and financial market responses within the global system.

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# III.A Appendix

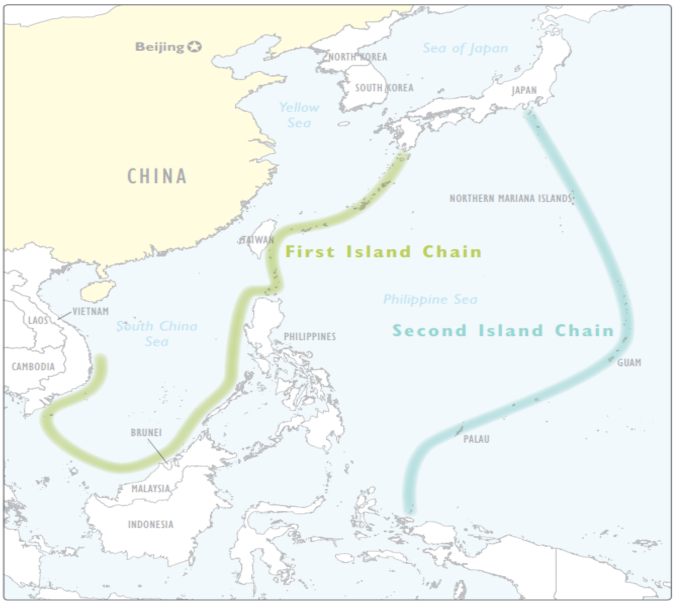
## III.A.1 The search query

We use the following search query on `Nexis Uni` to find articles mentioning the presence of U.S. CSG. To eliminate the effects of attention resulting from events that are unrelated to regional geopolitics, such as the outbreak of COVID-19 inside a U.S. aircraft carrier or supporting operation for disaster relief efforts, we exclude articles containing the related terms (e.g., “Covid”, “Thypoon”):

```
("U.S. aircraft carriers" OR "U.S. aircraft carrier" OR "US aircraft
carriers" OR "US aircraft carrier" OR "American aircraft carriers"
OR "American aircraft carrier" OR "U.S. Navy carriers" OR "U.S.
Navy carrier" OR "US Navy carriers" OR "US Navy carrier" OR "U.S.
carrier strike groups" OR "U.S. carrier strike group" OR "US
carrier strike groups" OR "US carrier strike group" OR "U.S.
Navy fleet" OR "US Navy fleet" OR "U.S. CVN" OR "US CVN" OR "U.S.
CGS" OR "US CGS" OR "nuclear-powered aircraft carrier" OR "nuclear-powered
aircraft carriers" OR "USS Ronald Reagan" OR "USS Nimitz" OR
"USS Gerald R. Ford" OR "USS Carl Vinson" OR "USS Theodore Roosevelt"
OR "USS Abraham Lincoln" OR "USS John Stennis" OR "USS George
Washington") AND ("South China Sea" OR "East China Sea" OR "Sea
of Japan" OR "East Sea" OR "Yellow Sea" OR "China" OR "Japan"
OR "Korea" OR "South Korea" OR "North Korea" OR "Korean Peninsula"
OR "Taiwan" OR "Beijing" OR "Seoul" OR "Pyongyang" OR "Taipei"
OR "Taiwan Strait" OR "Guam" OR "Okinawa" OR "Philippines" OR
"Philippine Sea" OR "Ryukyu Islands" OR "Paracel Islands" OR
"East Asia" OR "West Pacific") AND NOT ("Covid" OR "Virus" OR
"Corona" OR "Typhoon" OR "Haiyan" OR "Yolanda" OR "Haiti")
```

## III.A.2 The concept of island chains

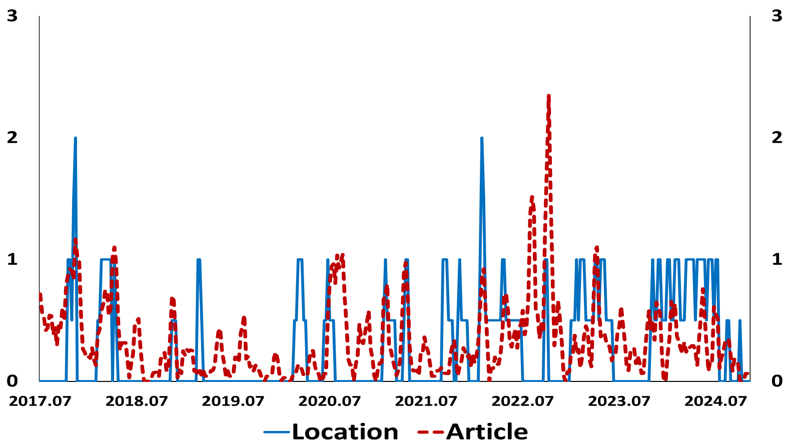
Figure III.A.1: The first and second island chains




*Note*: The figure is taken from U.S. Department of Defense (2012).

## III.A.3 Comparison of CSG indicators

Figure III.A.2: Comparison of CSG indicators on a weekly basis




*Notes*: A unit in the location of CSG (left-hand scale) is one additional deployment of CSG in the region, and a unit in the number of articles on CSG (right-hand scale) is one standard deviation, as discussed in the text. Both are weekly series, and the number of articles is a three-month moving average. *Source*: The location of CSG is from the USNI News. The number of articles on CSG is retrieved from Nexis Uni.

### III.A.4 Data sources

Table III.A.1: Data Definition and Sources

| Variable | Description | Sources |
|---|---|---|
| **Variables for individual countries** | | |
| Stock index | Major stock market indexes for each country | Bloomberg |
| FX | Nominal exchange rates against USD | Bloomberg |
| Interest rate | 10-year government bond yields | S&P Capital IQ |
| Equity inflows | Fund flows (%, divided by assets in the start of the period) | EPFR |
| CPI | Consumer Price Index (YoY change rate) | National statistics office |
| IND | Industrial Production (YoY change rate) | National statistics office |
| **Control variables** | | |
| GPR | The Geopolitical Risk index | Caldara and Iacoviello (2022) |
| VIX | CBOE Market Volatility Index | Bloomberg |
| SNP | Standard and Poor's 500 index | Bloomberg |
| US10Y | U.S. 10-year Treasury yield | FRED |

## Affidavit

Ich erkläre hiermit, dass ich die vorgelegten und nachfolgend aufgelisteten Aufsätze selbstständig und nur mit den Hilfen angefertigt habe, die im jeweiligen Aufsatz angegeben oder zusätzlich in der nachfolgenden Liste aufgeführt sind. In der Zusammenarbeit mit den angeführten Koautoren war ich mindestens anteilig beteiligt. Bei den von mir durchgeführten und in den Aufsätzen erwähnten Untersuchungen habe ich die Grundsätze guter wissenschaftlicher Praxis, wie sie in der Satzung der Justus-Liebig-Universität Gießen zur Sicherung guter wissenschaftlicher Praxis niedergelegt sind, eingehalten.

Jinyeong Yun
10. November 2025

I hereby declare that I completed the papers submitted and listed hereafter independently and only with those forms of support mentioned in the relevant paper. When working with the authors listed, I contributed no less than a proportionate share of the work. In the analyses that I have conducted and to which I refer in the papers, I have followed the principles of good academic practice, as stated in the Statute of Justus Liebig University Giessen for ensuring good scientific practice.

Jinyeong Yun
November 10, 2025

## Submitted papers

I. **Yun, J.** (2025), 'The Distributional Effects of Capital Inflows', *Open Economies Review* 36(5), 1527–1566.

II. **Yun, J.** (2025), 'Financial Integration and Wage Inequality: Evidence from European Countries', *Available at SSRN 5102501*

III. **Tillmann, P. & Yun, J.** (2025), 'Gains from Gunboats: U.S. Aircraft Carriers and Financial Markets in East Asia', *Available at SSRN 5085402*