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Monetary policy and income inequality: evidence from the Eurozone

By

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Résumé

Ce mémoire explore l'impact de la politique monétaire sur les inégalités de revenus dans la zone Euro. Une revue approfondie de la littérature est réalisée afin d'identifier les approches pertinentes et les résultats de recherches antérieures. Nous estimons les chocs de politique monétaire en utilisant une approche par SVAR avec une décomposition de Cholesky. Les chocs estimés sont ensuite utilisés dans des projections locales (*local projections*) afin d'analyser les effets des chocs monétaires sur les inégalités de revenus pour chaque pays dans notre échantillon. Les résultats indiquent une forte hétérogénéité entre les pays en réponse à un choc restrictif de politique monétaire de 25 points de base. Nous effectuons des tests de robustesse et constatons que la forte hétérogénéité est présente quelle que soit la spécification. Nous observons des augmentations plus faibles et des diminutions plus importantes en réponse à un choc restrictif pour notre échantillon de pays en utilisant les coefficients de Gini nets comme mesure des inégalités de revenus par rapport à l'utilisation des coefficients de Gini bruts, un résultat qui est en contradiction avec la littérature existante. Ces résultats indiquent que les inégalités de revenus sont affectées différemment au sein d'une union monétaire par rapport aux pays disposant de leur propre banque centrale.

Abstract

This thesis explores the impact of monetary policy on income inequality in the Eurozone. A thorough review of the literature is performed to identify the relevant approaches and findings of past research on the matter. We estimate monetary policy shocks using a SVAR approach with a Cholesky decomposition. The estimated shocks are used in local projections to analyze country-specific effects within the Eurozone. The results indicate strong heterogeneity across countries in response to a 25 basis points contractionary monetary policy shock. We perform robustness tests and find consistent strong heterogeneity across specifications. We find lower increases and larger decreases in response to a contractionary shock for our sample of countries using net Gini coefficients as the measure of income inequality relative to using gross Gini coefficients, in contradiction with the existing literature. These results provide evidence that income inequality is affected differently in a monetary union compared to countries with their own central bank.

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1 – Introduction

Members of the central banking community have been increasingly mentioning inequality in recent years, but no central bank currently has a mandate to use policy to influence income or wealth equality. For example, Agustín Carstens, General Manager at the BIS acknowledges that, even though inequality is not a monetary phenomenon, monetary policy does have an impact on inequality when pursuing their mandates.[1] Another example is Tiff Macklem, governor of the Bank of Canada, mentioning in 2021 that "inequality has long been a concern of the Bank of Canada".[2] Macklem also stated that leaving entrenched inequalities unattended can hold back citizens and economies from realizing their full potential.[3] Additionally, while serving as Chairman of the Federal Reserve in 2014, Janet Yellen stated that she was deeply concerned about the prevailing level of inequality in the United States, which has been continuously growing.[4] The increasing references to inequality by central bankers demonstrates the growing worries about rising economic disparity, its economic effects, and the implications for monetary policy.

Following the Great Financial Crisis (GFC), the issue of economic inequality has received greater attention. Piketty (2014) advocates for a new global capital tax to prevent excessive and ever-growing economic inequality. In response, Mankiw (2015) argues that some degree of inequality is essential for economic growth. Figure A.1 in the appendix illustrates the increasing trend in income inequality in many European countries since 1999. While inequality is an inherent part of capitalism and may be beneficial to economic development, excessive disparity may not be desirable for many reasons. In their study, Dabla-Norris *& al.* (2015) explore the repercussions of inequality. These repercussions include slower economic growth, lower levels of education, and an increased likelihood of conflict, which discourages investment. Berg *et al.* (2014) find a strong association between longer periods of strong, healthy, per capita growth and greater equality in the income distribution. They state that over "longer horizons, avoiding excessive inequality and sustaining economic growth may be two sides of the same coin".

[1] Carstens, A. (2021). Central Banks and inequality. Presentation at Markus' Academy, Princeton University's Bendheim Center for Finance, Basel, May 6. Retrieved from https://www.bis.org/speeches/sp210506.pdf
[2] Macklem, T. (2021). The benefits of an inclusive economy. Presentation at Universities of Atlantic Canada, Halifax, Nova Scotia, May 13. Retrieved from https://www.bankofcanada.ca/2021/05/benefits-inclusive-economy/
[3] Kilpatrick, S. (2021, November 9). Bank of Canada governor says central bank's role includes reducing inequality. The Canadian Press. Retrieved from https://www.ctvnews.ca/business/bank-of-canada-governor-says-central-bank-s-role-includes-reducing-inequality-1.5659404
[4] Cassidy, J. (2014, October 17). Rising Inequality: Janet Yellen Tells It Like It Is. The New Yorker. Retrieved from https://www.newyorker.com/news/john-cassidy/janet-yellen-tells

Since monetary policy is an important stabilisation tool in modern economies, it has a significant impact on economic outcomes, including income and wealth inequality (e.g., Coibion *et al.*, 2017 and Furceri *et al.*, 2018). Hafemann & al. (2018) state that any policy intervention focused on addressing important indicators such as inflation and GDP growth will surely have distributive implications. Given the possible repercussions of unchecked inequality and the inevitable impact of monetary policy, the objective of this thesis is to investigate the relationship between monetary policy and income inequality. Some initiatives such as the World Inequality Database (WID.world) and the OECD Wealth Distribution Database (WDD) have aimed at compiling databases of the wealth distribution. Unfortunately, wealth data can be hard to compare between countries, and ownership of assets can be difficult to clearly identify at the top of the wealth distribution. Our study focuses on income disparity using Gini coefficients, since the data is readily available and the findings regarding income inequality could potentially be applicable to wealth inequality.

The Gini coefficient is a widely used statistical measure of income inequality ranging from 0 to 100, where 0 represents perfect income equality and 100 represents perfect income inequality. Other popular measures of income inequality are national income ratios such as the 90/10 ratio, which represents the ratio between the incomes of the top and bottom deciles. We focus on the Gini coefficient as it is the most standard measure of inequality in the literature.

Using a structural vector autoregression (SVAR) with quarterly data, we identify exogenous monetary policy shocks for the Euro Area. We use linearly interpolated Gini coefficients at quarterly frequency along with the estimated monetary policy shocks in local projections to plot the impulse response functions (IRFs) of the gross and net Gini coefficients for individual countries within the Eurozone. We find that the impact is highly heterogeneous across countries. Our results indicate that the net Gini coefficients respond more to monetary policy shocks than the gross Gini, which differs from findings in the literature for countries not in a monetary union. We perform robustness checks to verify that our results are robust to different methodological approaches and find that the choice of interpolation has important implications in terms of the magnitude of the response, but that the ordering of the first 2 variables in the SVAR has no substantial impact. By comparing with other studies, our results indicate that income inequality responds differently to monetary policy shocks in a monetary union than in countries with independent central banks. In the setting of a monetary union, the fundamental process of monetary policy decisions may differ from that of countries with their own central bank. There may be significant disparities in how monetary policy affects income inequality in this environment, or significant differences in how the central bank of a monetary union establishes monetary policy while simultaneously assessing economic indicators from many countries. Considering the significant heterogeneity in the response of income inequality to monetary policy shocks within the Eurozone, more work needs to be done to better understand the interaction between the two, the main mechanisms through which the effects are propagated, and the implications for the ECB's monetary policy.

The rest of this thesis proceeds as follows. Section 2 discusses the recent literature on the subject, focusing on empirical findings regarding monetary policy and income inequality and the channels through which these effects are propagated. Section 3 gives an in-depth analysis of a simple Two-Agent New-Keynesian (TANK) model from Bilbiie *et al.* (2022) to provide guidance on the theoretical effects of interest rates on income inequality. Section 4 discusses the data used in this study and provides a detailed explanation of the methodology used. Section 5 presents the main findings using gross and net Gini IRFs to a monetary policy shock for each country. Section 6 provides robustness checks to establish the reliability of the main results. Finally, section 7 summarizes the findings of this study and the approach used to obtain them.

2 – Literature review

Given the empirical evidence presented by Berg *et al.* (2014), which establishes a robust relationship between extended durations of healthy and robust per capita growth and greater equality in the income distribution, it may be advantageous for welfare purposes if central bankers take these factors into account in their model formulations. Therefore, assessing the distributional impact of monetary policy is an important task in making more informed decisions and improving sustainable prosperity for all. Bilbiie *et al.* (2022) demonstrate, using a TANK model, that an expansionary monetary policy shock reduces income inequality. Their fully specified Heterogenous-Agent New-Keynesian (HANK) model asserts that both consumption and income inequality decrease after an expansionary monetary policy shock, while consumption inequality is more sensitive than income inequality. A baseline form of their model will be revisited in section 3 and is used as a reference to compare our empirical results.

Although the Great Financial Crisis sparked recent interest in the interaction between monetary policy and inequality, an earlier empirical literature on the topic includes works by Brownlee and Conrad (1961), Niggle (1989), Arestis and Howells (1991, 1994), and Thorbecke (1997). In this earlier literature, there is some evidence of increased inequality following contractionary monetary policy.

Recent studies obtain somewhat contradictory results about the distributional impacts of monetary policy. For example, using a panel of 32 developed and emerging countries, Furceri *et al.* (2018) find that income inequality experiences a persistent increase as a result of monetary policy tightening. These results are also in line with Coibion *et al.* (2017) and Mumtaz and Theophilopoulou (2015). This empirical evidence supports the theoretical framework developed by Bilbiie *et al.* (2022). On the other hand, Davtyan (2016) finds that contractionary monetary policy shocks in the U.S. decrease income inequality. Hafemann *et al.* (2018) obtain similar results for six advanced economies, where they find that expansionary monetary policy increases inequality for both gross (pre-tax and pre-transfers) and net (post-tax and post-transfers) Ginis. Hafemann *et al.* (2018) find that this effect on net Gini is not statistically significant for countries with high relative redistribution. Saiki and Frost (2014) find similar results to Hafemann *et al.* (2018) when analyzing the impact of Quantitative Easing in Japan.

As pointed out by Kappes (2023), there have been many attempts to identify the multiple channels through which monetary policy impacts inequality. Ampudia *et al.* (2018) give a concise identification of these channels as direct and indirect channels. Coibion *et al.* (2017) identified five channels through which monetary policy affects inequality: earnings heterogeneity, income composition, portfolio, savings redistribution, and financial segmentation. Finally, Bunn *et al.* (2018) classified these channels into six categories: interest income, labor income, financial wealth, housing wealth, pension wealth and inflation. Most of the literature on the subject uses similar categories for thinking about the transmission of policy.

Of these channels, the income composition is often most emphasized. Both Coibion *et al.* (2017) and Hafemann *et al.* (2018) provide evidence that the impact of monetary policy shocks on income inequality is heavily influenced by income composition. Hafemann *et al.* (2018) determine that an expansionary shock raises capital income more than labor income, which increases income inequality since higher-income households earn a larger share of their income from capital. Additionally, according to Coibion *et al.* (2017), wages, corporate income, and financial income respond differently to a contractionary monetary policy shock. These findings suggest that the response of the income distribution to a monetary policy shock is highly influenced by the income composition of households.

Monetary policy shocks must be identified before assessing the impact on income inequality. This is necessary due to the difficulty in distinguishing exogenous monetary policy shocks from monetary policy variations caused by changing macroeconomic conditions. Multiple approaches have been proposed in the literature to identify these shocks. The approach by Romer and Romer (2004) consists of combining projected federal funds rate changes around meetings of the Federal Open Market Committee (FOMC) with the Federal Reserve's internal inflation and real output predictions. Kappes (2023) states that they regress "the federal funds rate target for each FOMC meeting against key economic indicators such as GDP forecasts, GDP deflator, and unemployment provided by the Greenbook." The monetary policy shock is then identified as the residual of this regression, representing any unexpected change in the federal funds rate target. Coibion *et al.* (2017) as well as Davtyan (2016) use this approach and expand the time series of Romer and Romer's (2004) U.S. monetary policy shocks.

Initially developed to identify fiscal policy shocks, Furceri *et al.* (2018) used the approach from Auerbach and Gorodnichenko (2013) to identify monetary policy shocks. They established a strategy based on short-term interest rate forecasting mistakes, economic growth, and inflation. They compared the identified shocks to the approach by Romer and Romer (2004) and found a series of highly similar policy shocks. Kappes (2023) mentions another approach to defining shocks presented by El Herradi and Leroy (2019), in which the methodology of Stock and Watson (2018) is adapted using external instruments.[5]

[5] El Herradi and Leroy (2019) do not compare their identified monetary policy shocks with other existing series. They mention that these methods are "not tractable in [their] case since it requires (at least) forecasts of short-term interest rates, inflation, and GDP growth, which are not available over the long run". This is because their approach has a longer-term horizon (i.e., 1920-2015).

Once monetary policy shocks have been identified, there are two main approaches in the literature to estimate the linkage between monetary policy and income inequality. The first approach consists of using the local projection method proposed by Jordà (2005), which involves a regression analysis where the response variable (dependent variable) is regressed against a particular variable of interest (e.g., monetary policy shock) combined with lagged values of both variables and additional controls. By performing a series of regressions with different horizons, this method allows for an impulse response function to be estimated. The link between the dependent and independent variables can then be estimated across time. This approach is extensively employed in recent empirical studies in macroeconomics, including research on monetary policy and income inequality (e.g., Coibion *et al.* (2017), Furceri *et al.* (2018), and Hafemann *et al.* (2018)).

The second method to identify monetary policy shocks is the VAR estimation approach. For example, along with standard macroeconomic variables such as real GDP, consumer prices, short-term interest rates, and the trade-weighted real effective exchange rate (REER), Hafemann *et al.* (2018)'s VAR, building on the approach of Bernanke and Gertler (1995), includes either measures of inequality such as the Gini coefficient or variables to identify transmission mechanisms for each of the six countries in the sample. The VAR approach can be used both to identify monetary policy shocks, as well as the impulse response of the Gini coefficients following a monetary policy shock, after applying restrictions (e.g., sign restrictions used in Hafemann *et al.* (2018) or a recursive ordering using a Cholesky decomposition).

Using local projections and the Romer and Romer (2004) approach to identify monetary policy shocks, Coibion *et al.* (2017) find that contractionary monetary policy shocks increased inequality in job earnings, total income, consumption, and total spending in the United States. These results contrast with Davtyan (2016) who also used the Romer and Romer (2004) approach to identify monetary policy shocks. In contrast to Coibion *et al.* (2017), Davtyan (2016) used shocks estimated with the same approach in a VAR and found that contractionary shocks decreased income inequality. According to Davtyan (2016), these results differ from Coibion *et al.* (2017) because Davtyan (2016) included the top 1%, while Coibion *et al.* (2017) did not.

Using a sample of 32 developed and emerging countries, Furceri *et al.* (2018) report findings of asymmetrical effects of monetary policy shocks on inequality. Their study reveals that income inequality (as measured by Gini mean estimators from the Standardized World Income Inequality

Database (SWIID) developed by Solt (2020)) increases following a contractionary monetary policy shock. The effect is not statistically significant for an expansionary shock. Furceri *et al.* (2018) find that countries with high relative redistribution do not face a statistically significant impact of income inequality following a contractionary monetary policy shock, but countries with low relative redistribution do. They find that nations with higher labor shares are more vulnerable to the effects of monetary policy shocks on inequality.

Using a subsample of six developed countries from the same income data as Furceri *et al.* (2018), Hafemann *et al.* (2018) obtained contradictory results. According to their analysis, expansionary monetary policy increases inequality for both gross (pre-tax and pre-transfers) and net (post-tax and post-transfers) Gini measures. The two studies use different data to measure monetary policy shocks. While Furceri *et al.* (2018) relies on short-term interest rates, Hafemann *et al.* (2018) use shadow interest rates. They also use different approaches to identify monetary policy shocks. Furceri *et al.* (2018) relies on the approach from Auerbach and Gorodnichenko (2013), whereas Hafemann *et al.* (2018) use sign restrictions in a VAR based on Uhlig (2005). Hafemann *et al.* (2018) use the VAR approach but also apply the local projection approach as a robustness check. Similarly to Furceri *et al.* (2018), they find that higher degrees of relative redistribution affect the distributional impact of monetary policy. The effect of an expansionary monetary policy shock on income inequality is not statistically significant for countries with high degrees of redistribution. The difference in the sample of countries considered may also be another reason for divergent results with Furceri *et al.* (2018). Hafemann *et al.* (2018) hypothesize that the difference between their results and Coibion *et al.* (2017) is because Coibion *et al.* (2017)'s data does not cover the top 1% of income earners. This important difference has significant implications considering the importance of the top earners in the income distribution. They also attribute the differences in results to differences in sample periods, as well as the estimation approach used in the respective papers.

Using a panel vector autoregression (PVAR) for Eurozone countries and simple OLS regressions for select countries, Liosi and Spyrou (2022) find that expansionary monetary policy increases income inequality. They use the European shadow interest rate as well as the 80/20 Income Quantile Share Ratio "measured in terms of equivalised disposable income" as their main income inequality measure. At the country level, they find the effect is heterogeneous on impact. They mention that “for Ireland, Germany and the Netherlands monetary policy has no impact on income inequality or a weak impact (France), while for Spain, Portugal, Greece and Italy the impact is more pronounced.” From Panel A1 of Table 2 of their paper, we notice that Greece and Portugal experience an increase in income inequality from contractionary monetary policy, whereas Spain experiences a decrease in income inequality following an increase in the shadow interest rate. They do not plot IRFs to a monetary policy shock for country-specific effects, but plot one average effect across all countries using the PVAR, which illustrates a decrease in income inequality from a contractionary monetary policy shock. The PVAR approach used by Liosi and Spyrou (2022) provides a temporal analysis but does not allow for country-specific analysis over time. Their regressions for country-specific effects only consider the impact of a monetary policy shock on income inequality for the same period. In contrast, this thesis examines the country-specific effects of monetary policy shocks in the Eurozone over time.

In this study, we will also use the shadow interest rate for the Eurozone as our measure of monetary policy. Changes in interest rates, open market operations in purchasing and selling government bonds, changes in reserve requirements, and forward guidance are all examples of monetary policy tools. Shadow interest rates enable us to consider the periods where interest rates hit the zero lower bound (ZLB) and to consider other unconventional monetary policies. Wu and Xia (2016) demonstrate how useful the shadow rate can be to summarize monetary policy information at the ZLB, which is also in line with Bullard (2012) and Krippner (2012), as they point out.

We estimate monetary policy shocks using the SVAR approach with a Cholesky decomposition, which allows for the use of readily accessible data. Historical forecasting data as used in Furceri *et al.* (2018) are not publicly available and would be costly in the context of this study. Another appeal of the VAR approach is that it allows us to compare the results to other studies utilizing the method, e.g., Hafemann *et al.* (2018). Their study covers a similar period to the present study and uses both VAR and local projections (as a robustness check) to estimate the impact of monetary policy shocks on inequality. After estimating the monetary policy shocks using the VAR approach, we use the local projection method and compare the results with those obtained by Hafemann *et al.* (2018), with the difference of focusing on countries within the Eurozone. This study will be able to compare different responses to monetary policy within the Eurozone and compare to the results of Hafemann *et al.* (2018) which focus on countries with independent central banks. The results will then be compared to those obtained in the theoretical model derived by Bilbiie *et al.* (2022) presented in the following section.

3 - TANK model by Bilbiie *et al.* (2022)

This section discusses the simple two-agent model developed by Bilbiie *et al.* (2022). We present the simplest form of their model which will prove useful to interpret our results in section 4. Their model consists of two types of agents: savers (S), corresponding to $(1-\lambda)$ of the population, and hand-to-mouth spenders (H), representing λ of the population. The model is a closed economy with a central bank that exogenously controls interest rates. Savers are the only agents who can access risk-free bonds and capital markets. Bilbiie *et al.* (2022) impose a net zero bond supply. To have a simple model with analytical results, they remain agnostic about the exact functional form of the investment function. Thus, the investment function is exogenously given by:

$$I_t = f(Y_t, r_t, \dots).$$

The budget constraint is

$$C_t^S + \frac{1}{1-\lambda} I_t = \widetilde{Y_t^S} + T^S = Y_t^S,$$

where $\widetilde{Y_t^S}$ represents the gross income of savers, and T^S are steady-state, constant redistributive transfers, which are used to impose that both savers and hand-to-mouth households consume equally at the steady state. This assumption is strong but allows for the model to be solved analytically. Y_t^S is the net income of savers which is equal to gross income plus transfers (or minus redistributive taxes). Since only savers can accumulate capital, and represent (1-λ) of the population, then:

$$I_t = (1-\lambda) I_t^S \qquad \text{so} \qquad I_t^S = \frac{1}{1-\lambda} I_t\,.$$

Assuming the utility function is given by $U(C_t) = lnC_t$, the Euler equation related to bonds is given by:

$$(C_t^S)^{-1} = \beta \mathrm{E_t}[(1+r_t)(C_{t+1}^S)^{-1}].$$

Considering that hand-to-mouth households do not have access to credit or capital markets, their budget constraint is the following:

$$C_t^H = \widetilde{Y_t^H} + T^H = Y_t^H,$$

where T^H are redistributive transfers and Y_t^H represents net income. Hand-to-mouth households consume all their net revenue each period and do not have the option to save through capital investment or bond markets. Imposing equilibrium, we get the following market clearing equation:

$$C_t + I_t = Y_t,$$

where

$$Y_t = \lambda Y_t^H + (1-\lambda)Y_t^S,$$

$$C_t = \lambda C_t^H + (1-\lambda)C_t^S.$$

As mentioned earlier, both households have the same consumption at the steady state $C^H = C^S = \bar{C}$. This is achieved by using tax and transfers $(1-\lambda)T^S = \lambda T^H$ which implies that the government's budget is balanced.

Derived from the investment function, the steady-state investment to output ratio is:

$$I_Y \equiv \frac{\bar{I}}{\bar{Y}}$$

where $\bar{I}$ represents steady-state investment and $\bar{Y}$ represents steady-state real output. From the market clearing equation, the consumption to output ratio is given by:

$$C_Y \equiv \frac{\bar{C}}{\bar{Y}} = 1 - \frac{\bar{I}}{\bar{Y}} \ .$$

The model is log-linearized around the symmetric steady state using the following notation to express the first order log-linear approximation:

$$\widehat{x_t} = \frac{X_t - \bar{X}}{\bar{X}} \cong lnX_t - ln\bar{X}.$$

The Euler equation related to bonds becomes:

$$\widehat{c_t^S} = E_t\widehat{c_{t+1}^S} - \widehat{r_t},$$

where R is the gross interest rate, and $\widehat{r_t} \cong lnR_t - ln\bar{R}$. The budget constraint of savers becomes:

$$C_Y\widehat{c_t^S} + \frac{I_Y}{1-\lambda}\widehat{\iota_t} = Y_Y^S\widehat{y_t^S},$$

where $Y_Y^S \equiv \frac{\overline{Y^S}}{\bar{Y}}$. The hand-to-mouth budget constraint is:

$$\widehat{c_t^H} = \widehat{y_t^H} .$$

The log-linearized market clearing condition is given by:

$$\widehat{y_t} = C_Y\widehat{c_t} + I_Y\widehat{\iota_t}.$$

By imposing that $C^H = C^S = \bar{C}$ at the steady state (using $T^H and\, T^S$), log-linearized aggregate consumption and income are given by:

$$\widehat{c_t} = \lambda\widehat{c_t^H} + (1-\lambda)\widehat{c_t^S},$$

$$\widehat{y_t} = \lambda Y_Y^H\widehat{y_t^H} + (1-\lambda)Y_Y^S\widehat{y_t^S}.$$

Bilbiie *et al.* (2022) assume that the net income of the hand-to-mouth agents responds to aggregate income with elasticity χ:

$$\widehat{y_t^H} = \chi\widehat{y_t}.$$

Thus, by using the log-linearized condition for aggregate income and substituting $\widehat{y_t^H}$ by $\chi\widehat{y_t}$, we get the following solution for the income of saver households:

$$\widehat{y_t^S} = \frac{1-\lambda\chi Y_Y^H}{(1-\lambda)Y_Y^S}\widehat{y_t} .$$

We can derive an equation for income inequality, which is measured as the percentage change differentials in income levels, $\widehat{y_t^S} - \widehat{y_t^H}$:

$$\widehat{y_t^S} - \widehat{y_t^H} = \frac{1-\lambda\chi Y_Y^H}{(1-\lambda)Y_Y^S}\widehat{y_t} - \chi\widehat{y_t} .$$

Using the fact that $Y_t = \lambda Y_t^H + (1-\lambda) Y_t^S$ still holds at the steady state, we can rearrange the previous equation as:

$$\widehat{y_t^S} - \widehat{y_t^H} = \frac{1-\chi}{(1-\lambda)Y_Y^S} \widehat{y}_t \ .$$

The cyclicality of the income distribution is determined by the elasticity χ. As discussed in Bilbiie *et al.* (2022), "income inequality is countercyclical'' in this model if and only if $\chi > 1$ (considering $0 \leq \lambda \ \leq 1$, and that $\lambda = 0$ or $\lambda = 1$ both imply a RANK model). This elasticity (or multiplier) has been described by Bilbiie (2008) and empirically supported by Patterson (2023). Patterson (2023) demonstrates that lower-income individuals tend to have an overall earnings elasticity higher than 1, which provides evidence in favor of the countercyclicality of income inequality derived from the simple two-agent model. In this model, income (real output) is exogenous since there is no supply side. If we assume, based on economic theory, that an increase in real interest rates reduces real economic activity, this simple model predicts that a tightening of monetary policy leads to an increase in income inequality.[6] The more complex HANK model in Bilbiie *et al.* (2022) predicts a similar outcome whereas an expansionary monetary shock decreases income inequality. We use these conclusions to compare to our empirical results obtained in section 5.

4 - Data & Methodology

In this section, we present the methodology used in our baseline empirical analysis. To estimate the impact of monetary policy shocks on income inequality, we consider twelve countries from the Eurozone. The countries were selected based on being part of the Eurozone since its inception in 1999, apart from Greece, which joined in 2001. Our sample consists of Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, and Spain.

[6] According to the model, an increase in nominal interest rates (contractionary policy), which is likely to result in a rise in real interest rates, is expected to increase income inequality.

4.1 - Data

We use the shadow interest rate for the Eurozone computed using the method of Wu and Xia (2017) and Wu and Xia (2020) which starts in September 2004. We analyze the Shadow Interest rate in conjunction with the ECB Deposit Facility Rate for the Euro Area, which was obtained from the Federal Reserve Economic Data (FRED) as our measure of monetary policy before September 2004. We take the average quarterly figures from monthly observations for this monetary policy measure to match the frequency of other variables.

We use the Gini coefficient as our measure of inequality because it is extensively used to measure inequality in the literature, despite not being a comprehensive measure of economic inequality.[7] For example, it does not account for access to education, healthcare services, or social services. Additionally, this research seeks to compare results obtained for the Eurozone with the existing literature, including the results of Hafemann *et al.* (2018). Consequently, Gini coefficients, as computed by Solt (2020) and available in SWIID,[8] are used as our income inequality measures. Mean relative redistribution estimates, also from SWIID, are used to support the research's findings. Relative redistribution is the percentage reduction in income inequality as a consequence of fiscal policies. Finally, the monetary policy shocks are calculated using data from the Area Wide Model (AWM) database in combination with some of the previously mentioned variables.[9] The AWM is a quarterly structural macroeconomic model for the Eurozone. According to Fagan *et al.* (2005), this model aims at assessing economic and monetary conditions, helping with macroeconomic forecasting as well as policy analysis and gaining a better insight of how the Eurozone economy works. Effectively, it allows us to obtain aggregated macroeconomic data for the Eurozone economy, which is used to identify ECB monetary policy shocks. It provides us with market GDP and the GDP deflator for the Eurozone, which together give us a measure of real GDP. We also use the seasonally-and-working-day-adjusted HICP as our price index for inflation. Regarding the span of our dataset, our analysis is bounded by the Eurozone's inception in 1999Q1. The latest available data for our study is 2017Q4, marking the endpoint of the AWM's current availability.

[7] For instance, Coibion et al. (2017), Davtyan (2016), Furceri et al. (2018), and Hafemann et al. (2018) all use Gini coefficients as their primary measure or as one of their measures of income inequality.
[8] See Solt (2020), data downloaded from https://fsolt.org/swiid/
[9] See Fagan *et al.* (2005), data downloaded from https://eabcn.org/page/area-wide-model

4.2 - Income Inequality Trends

Using the mean estimators from SWIID, Figure 1 illustrates the evolution of gross Gini coefficients for the selected countries over our sample period.

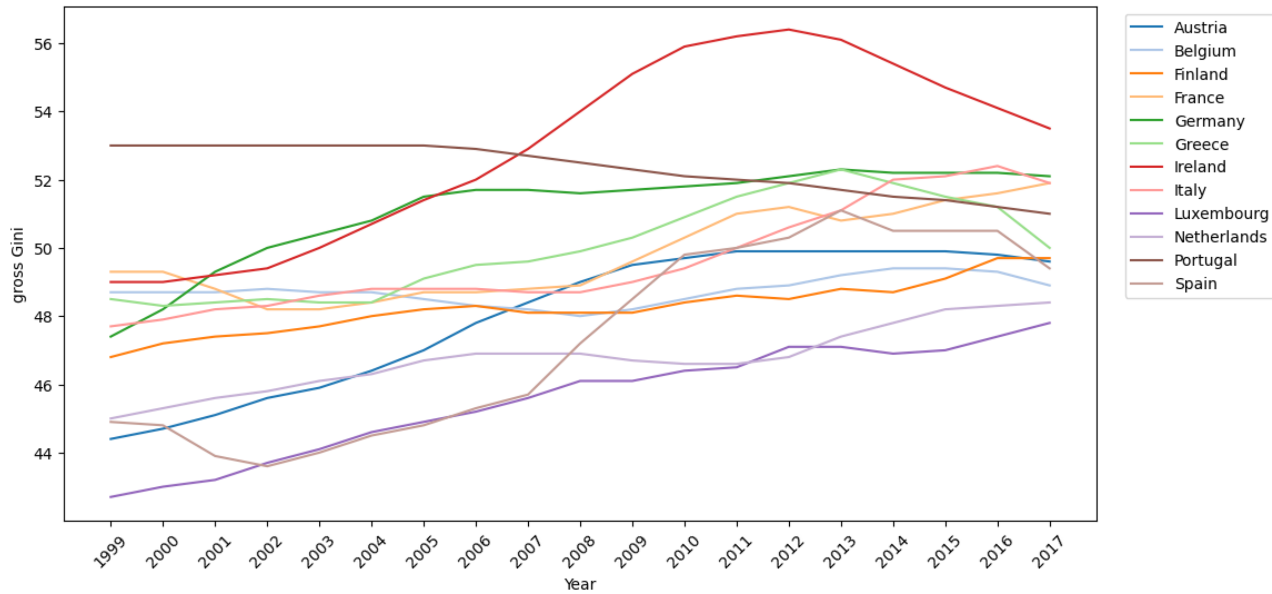

Figure 1 – Gross Gini coefficients from 1999 to 2017

Note: Gross Gini coefficients by country for our sample period (1999 to 2017). Data retrieved from SWIID (https://fsolt.org/swiid/).

We notice an increasing trend in most of our sample countries, except for Portugal which is an outlier in that regard. Portugal did start the sample period with the highest level of income inequality, but it levelled out with other countries over the two decades that followed. Ireland is another anomaly. Starting the sample period with a gross Gini coefficient similar to other countries, Ireland witnessed a drastic increase in income inequality over the following twelve years, followed by a partial reversal in the last five years of our sample. To better understand true income disparity measured by disposable income, Figure 2 presents a similar chart for net Gini coefficients (after tax and transfers).

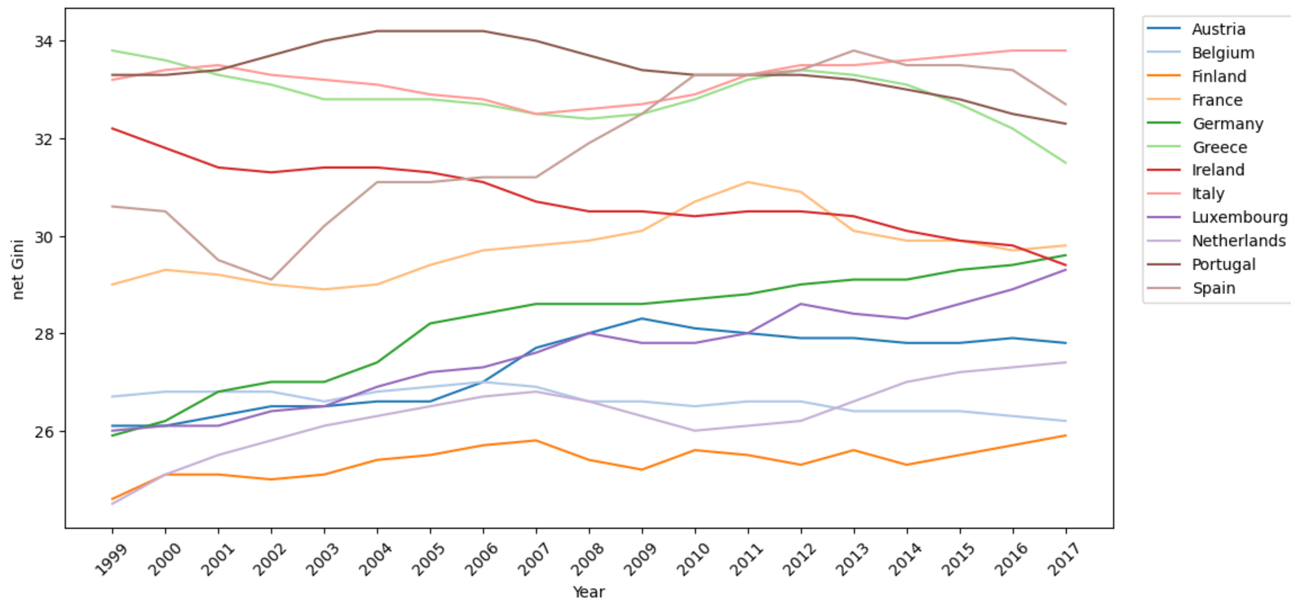

Figure 2 – Net Gini coefficients from 1999 to 2017

Note: Net Gini coefficients by country for our sample period (1999 to 2017). Data retrieved from SWIID (https://fsolt.org/swiid/).

Notice the important difference in the level of income inequality on the Y-axis relative to the gross Gini coefficients (Figure 1). This illustrates sizable impact of redistribution in reducing income inequality. Additionally, there does not appear to be a clear trend as was the case with the gross Gini coefficients. In our sample, Finland has the lowest degree of income inequality after tax and transfers for most of the period under study. Interestingly, most countries experience a noticeable increase in gross income inequality after the GFC, but this effect is greatly reduced when considering net income inequality.

4.3 – Identifying Monetary Policy Shocks

Using an Augmented Dickey-Fuller test both on the level and the first difference, we confirm that (log) HICP, (log) real GDP, and the shadow rate are all integrated of order 1 at a 95% confidence level. Performing a Phillips-Perron unit-root test for robustness yields the same result. By performing an Engle-Granger test of cointegration, we cannot reject the null hypothesis of no cointegration at a 95% confidence level. Thus, estimating the SVAR of the log levels (except for the shadow rate) would give us consistent estimators as pointed out by Hanson (2004). Estimating the SVAR using the first difference would also be adequate here but would not enable us to identify the level of exogenous monetary policy shocks. Instead, this approach would give us an exogenous shock in monetary policy variation, which would not be helpful for our local projection approach to estimate the impact of monetary policy shocks on income inequality. Hanson (2004) also points out that "Bernanke and Mihov (1998) report few differences between estimation in log differences and in log levels". Variables used to estimate our SVAR all have quarterly observations.

The reduced-form VAR is the following:

$$X_t = \Omega + \Phi_p(L)X_{t-p} + V_t. \qquad \text{[Eq. 4.1]}$$

Where X_t is a vector of (log) HICPSYA (p_t), (log) real GDP (y_t), and the shadow rate (S_t). Ω is a vector of constants, V_t is a vector of prediction errors, and $\Phi_p(L)$ is a lag-polynomial matrix in the lag operator L, where p = 1, …, P_{max}. The covariance matrix of V_t is denoted by Σ_V. The process of estimating exogenous shocks is as follows. The first step is to identify monetary policy shocks using the SVAR approach by applying a Cholesky decomposition using recursive ordering. This approach is commonly used to identify impulse response functions (IRFs) and can also be used to identify exogenous shocks. The work of Sims (1980) is one of the earliest influential papers proposing a Cholesky decomposition with a SVAR. With this approach, we impose a causal ordering on the endogenous variables, which enables us to identify the effect of exogenous structural shocks on each variable. We start by estimating the reduced form VAR. The lag P_{max} = 2 is selected by HQIC. The structural VAR related to the reduced-form VAR is the following:

$$X_t = \Omega + \Phi_p(L)X_{t-p} + A^{-1}U_t, \qquad \text{[Eq. 4.2]}$$

which can also be rewritten as

$$AX_t = \Theta + \Gamma_p(L)X_{t-p} + U_t \qquad \text{[Eq. 4.3]}$$

where $\Theta = A^{-1}\Omega$ and $\Gamma_p = A^{-1}\Phi_p$.

When estimating a VAR, Hafemann *et al.* (2018) simply state that they order the variables as real GDP, prices, and interest rates. They do not provide any justification about the order of the variables. Eichenbaum and Evans (1995) impose the same causal ordering, where they start with (log) US industrial production (Y), followed by (log) price level (P), followed by monetary policy (proxied by the ratio of nonborrowed to total reserves [NBRX]). They offer no justification on the ordering of production before price level, but they mention that their ordering "corresponds to the assumption that the U.S. monetary authority looks at the contemporaneous values of P_t and Y_t when setting $NBRX_t$". Thus, we keep the assumption that the monetary authority looks at same-quarter values of real GDP and prices when setting its policy. Our ordering differs for the other two variables, for which we present a credible explanation below as to the reason why our causal ordering makes sense. Alternatively, a robustness test in section 6 presents the results when we use the ordering suggested by these authors which corresponds to placing production before price level in the ordering of the VAR. We compare the findings in section 6 to those obtained in our baseline analysis using the previously mentioned ordering and find no substantial differences.

We impose that aggregated prices are not contemporaneously affected by other endogenous variables. Batini and Nelson (2001) provide evidence that monetary policy has a lagged impact on inflation. Furthermore, we impose that inflation is only affected by real GDP with a lag. This constraint seems reasonable when considering it may take time (at least one quarter) for changes in real GDP to feed through prices. Consumers and businesses may have higher inflation expectations for quarters to come if an increase in real GDP is interpreted as a sign of stronger economic conditions and incoming pricing pressures. In contrast, if real GDP declines, this can be interpreted as a sign of negative economic circumstances and reduced future pricing pressures, which may cause a decrease in inflation expectations, thus resulting in a decrease in actual future inflation.

Our ordering also allows for inflation to impact real GDP contemporaneously. Price increases may cause a decrease in real consumer expenditure because customers must spend more money to purchase the same goods and services. Their real budget is reduced since salaries are sticky. As inflation increases, real GDP is negatively impacted in the same period. It also seems reasonable to assume that monetary policy only impacts GDP with a lag. This presumption is because economic agents will assess their investment and spending choices for the quarters or years that follow. It takes time for them to reassess the new economic conditions and adjust their plans accordingly.

Moreover, since central bankers have information regarding same-quarter inflation and GDP fluctuations, we can reasonably assume that monetary policy reacts contemporaneously to these variables.

Thus, we order the variables as follows: $X_t = \begin{bmatrix} p_t \\ y_t \\ S_t \end{bmatrix}$.

By using the recursive approach, imposing the causal ordering corresponds to imposing that the A matrix is a lower triangular matrix:

$$A = \begin{bmatrix} 1 & 0 & 0 \\ \alpha_{yp} & 1 & 0 \\ \alpha_{sp} & \alpha_{sy} & 1 \end{bmatrix}$$

The covariance matrix of the SVAR is left as is (i.e., we do not normalize Σ_U). We use the Cholesky decomposition of the covariance matrix Σ_V to obtain the estimated A matrix. From the relationship between the SVAR and reduced form VAR, we know that $V_t = A^{-1}U_t$. Thus, by multiplying the reduced form VAR innovations by the matrix A, we get the structural shocks. We can then obtain the IRFs related to this SVAR.

With $X_t = \begin{bmatrix} p_t \\ y_t \\ S_t \end{bmatrix}$, we obtain the IRFs by plotting over time the response of each variable to a shock to one of the variables. Without loss of generality we can consider a reduced form VAR(1):[10]

$$X_t = \Omega + \Phi X_{t-1} + V_t \qquad \text{[Eq. 4.4]}$$

which can be written in vector MA(∞) form as:[11]

$$X_t = \mu + \sum_{i=0}^{\infty} \Phi^i V_{t-i}; \qquad \text{[Eq. 4.5]}$$

[10] A VAR(P) can always be expressed as a VAR(1) by using the Companion form.
[11] Inspired by Hamilton, J. D. (1994). Time series analysis (pp. 318-323). Princeton University Press.

where μ = [I – Φ]$^{-1}\Omega$.

Hence, the matrix Φ^i is interpreted as

$$\frac{\partial X_{t+h}}{v_{s,t}} = \Phi^h. \qquad \text{[Eq. 4.6]}$$

Since $V_t \equiv A^{-1}U_t$, the vector MA(∞) in Eq. 4.5 can be rewritten as:

$$X_t = \mu + \sum_{i=0}^{\infty} \Phi^i A^{-1} U_{t-i} \qquad \text{[Eq. 4.7]}$$

Thus, we can represent the response of each variable in vector X at time *t* to a one-standard-deviation shock in S (σ_s) as:

$$\frac{\partial X_t}{u_{s,t}} = \begin{bmatrix} \frac{\partial x_{p,t}}{u_{s,t}} \\ \frac{\partial x_{y,t}}{u_{s,t}} \\ \frac{\partial x_{s,t}}{u_{s,t}} \end{bmatrix} = A^{-1} \begin{bmatrix} 0 \\ 0 \\ \sigma_s \end{bmatrix}, \qquad \text{[Eq. 4.8]}$$

which implies that the response on period ahead is:

$$\frac{\partial X_{t+1}}{u_{s,t}} = \begin{bmatrix} \frac{\partial x_{p,t+1}}{u_{st}} \\ \frac{\partial x_{y,t+1}}{u_{s,t}} \\ \frac{\partial x_{s,t+1}}{u_{s,t}} \end{bmatrix} = \Phi A^{-1} \begin{bmatrix} 0 \\ 0 \\ \sigma_s \end{bmatrix}. \qquad \text{[Eq. 4.9]}$$

We can then generalize to other horizons as:

$$\frac{\partial X_{t+h}}{u_{s,t}} = \begin{bmatrix} \frac{\partial x_{p,t+h}}{u_{s,t}} \\ \frac{\partial x_{y,t+h}}{u_{s,t}} \\ \frac{\partial x_{s,t+h}}{u_{s,t}} \end{bmatrix} = \Phi^h A^{-1} \begin{bmatrix} 0 \\ 0 \\ \sigma_s \end{bmatrix}. \qquad \text{[Eq. 4.10]}$$

Figures A.2 to A.4 in the appendix illustrate these IRFs. All responses that we estimate are in reaction to a one-standard-deviation shock. An increase of one standard deviation in the shadow interest rate has a negative impact on inflation, although not in a statistically significant way. Real GDP is also negatively impacted by a shock to the shadow rate, and the effect is statistically significant after 5 quarters. The trajectory of the shadow rate to its own positive shock is also coherent: the effect is positive and persistent but becomes not statistically significant after 10 quarters. A shock to inflation leads to a statistically significant increase in the shadow interest rate, which fades over time. Real GDP also responds positively to a shock to inflation, but not in a statistically significant way. We notice a statistically significant decrease in real GDP after 7 quarters. A shock to real GDP increases inflation. This effect is persistent, but only significant for quarters 2 to 5 following the shock. Real GDP reacts positively to its own shock, which is not surprising. This effect is significant up to 5 quarters after the shock. Finally, the shadow interest rate reacts positively to a one-standard-deviation shock to real GDP. This effect is statistically significant up to 4 quarters after the shock.

4.4 – Estimated Monetary Policy Shocks

After estimating our SVAR, we recover estimates for the monetary policy shocks. We use the difference between the realization of each variable and its prediction to obtain the non-structural residuals. Using the fact that

$$\mathrm{V}_t \equiv A^{-1}\mathrm{U}_{t}, \qquad \text{[Eq. 4.11]}$$

which implies

$$A\mathrm{V}_t \equiv \mathrm{U}_{t}. \qquad \text{[Eq. 4.12]}$$

Multiplying V_t by the A matrix gives us three series of structural shocks, one for each variable in our SVAR. The estimated monetary policy shocks will be used in the local projection to estimate the impact on income inequality. By computing Portmanteau (Q) tests for white noise, we confirm that all structural shocks are white noise as expected, since we cannot reject the null hypothesis of no serial autocorrelation in the structural shocks at a 95% confidence level. The time series of estimated structural monetary policy shocks is presented in Figure 3. The structural shocks for inflation and GDP are presented in Figure A.5 in the appendix.

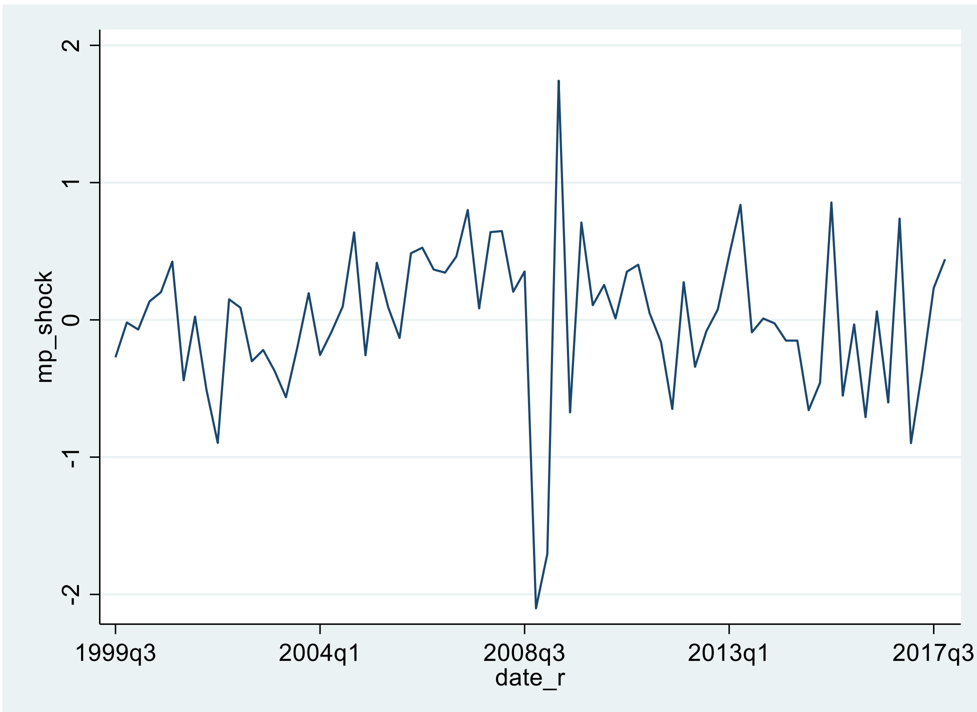

Figure 3 – Estimated Monetary Policy Shocks

Note: Estimated monetary policy shocks using a Cholesky decomposition. The Y axis represents the identified monetary policy shock for each quarter in percentage points.

We confirm that the identified monetary policy shocks are reasonable by looking at the shocks at key dates. First, we see that the expansionary shock of 2001Q4 is coherent with a response of the ECB to financial instability fueled by the bursting of the dot-com bubble. Second, the expansionary shocks we observe in 2008Q4 and 2009Q1 coincide with the GFC and the downfall of Lehman Brothers, which shook the worldwide financial system. Significant losses and liquidity concerns struck major banks and financial institutions. The GFC pushed the ECB to adopt an expansionary monetary policy to prevent further economic deterioration. Finally, the Eurozone was experiencing a sovereign debt crisis in late 2011, which produced significant volatility in financial markets and threatened the Eurozone's economic stability. The ECB had to intervene since several European banks were under stress and were threatening a wider banking crisis. This intervention is consistent with the negative monetary policy shock in late 2011.

4.3 – Impulse-Response Function by Local Projection

To estimate the impact of monetary policy shocks on income inequality, we approximate the change in income inequality by taking the log-difference of Gini coefficients at time $t+h$ and $t-1$. These coefficients are the mean estimators from SWIID. Our baseline analysis uses both market (gross) and disposable (net) Gini coefficients. Inspired by the regression by Furceri *et al.* (2018), we plot the IRF for each country in our sample[12]

$$(y_{t+h} - y_{t-1}) = c^h + \beta^h MP_t + \gamma^h \Delta y_{t-1} + \delta^h MP_{t-1} + \varepsilon_t^h. \qquad \text{[Eq. 4.12]}$$

y is the log of income inequality, c is a constant, MP_t is the previously computed monetary policy shock for a given quarter t (MP is the same monetary policy shock for all countries considering we are working in the context of a monetary union). The index h represents the horizon of the regression in quarters which ranges from 0 to 24. β^h is the parameter of interest in this regression. This coefficient captures the impact of a monetary policy shock at time t on our inequality measure at time $t+h \;\; \forall \;\; h \geq 0$. Using β^h, we can plot the impulse response of our income inequality measure by multiplying β^h for each h up to $h_{max} = 24$ by the intensity of the shock (25-basis-points (bps) in our case). To get an interpretation in percentage points, we further multiply this result by 100. As in Hafemann *et al.* (2018), Newey-West standard error corrections are applied to bypass the serial correlation issue. Using this approach, we will be able to visualize how different countries are impacted by a 25-bps contractionary shock.

Unfortunately, our data sample from SWIID does not provide sufficient observations to get country specific local projection IRFs with yearly observations considering the span of our sample

[12] Furceri *et al.* (2018) use $(y_{t+h} - y_t)$ as their dependent variable, but this specification does not allow for any effect on impact. We choose to follow the specification most used in local projections which corresponds to $(y_{t+h} - y_{t-1})$ and allows for a response on impact. Furceri *et al.* (2018) apply a panel local projection with country and time fixed effects, but we prefer to focus on country-specific effects by evaluating the IRF for each country individually which is in line with the approach of Hafemann et al. (2018). The difference with Hafemann et al. (2018) is that their dependent variable is directly the measure of income inequality.

period. Linear interpolation, as used by Hafemann *et al.* (2018), serves as a simple approach to address this issue. We chose to use linear interpolation as our baseline approach because it is a simple approach, and it is greatly used in empirical research. We were unable to find any other inequality measure available at quarterly frequency. An alternative interpolation approach will be used in section 6 to validate the robustness of our main results. We linearly interpolate Gini coefficients and use them in our local projection after taking their log. This allows us to generate quarterly observations from yearly data. Using the estimated quarterly monetary policy shocks, we now have the necessary data to run individual local projections for each country in our sample.

5 – Results

In this section, we present the findings of our analysis of the effect of monetary policy shocks on income inequality. We specifically investigate how exogenous shocks to the European shadow interest rates impact the income distribution of our sample countries as assessed by the gross (market) and net (after tax and transfers) Gini coefficients from SWIID.

5.1 - Gross Gini

Our baseline case uses linearly interpolated market Gini coefficients as well as our quarterly estimated monetary policy shocks to evaluate the effect of monetary policy on income inequality. We use Jordà's (2005) local projection approach to plot the impulse response of our measure of income inequality to a 25-bps shock in monetary policy. We exclude observations for Greece before 2001Q1 since the country only adopted the Euro at the start of January of the same year. Using individual datasets for each regression, Figure 4 illustrates the IRF for each country.

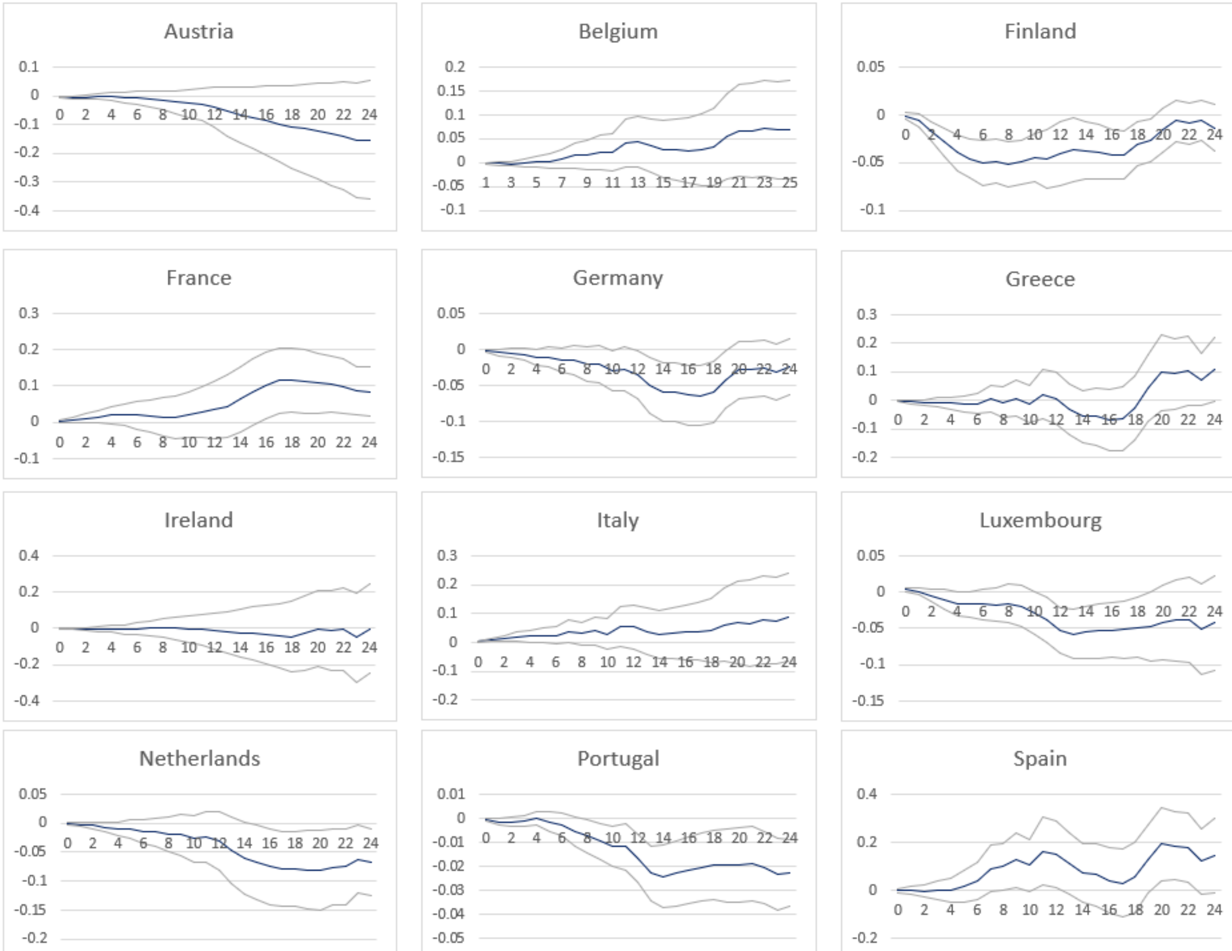

Figure 4 – Market Gini IRFs to 25-bps shock in monetary policy

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

Figure 4 shows important heterogeneity in the IRFs. Six out of twelve countries experience a decrease in income inequality over 24 quarters, five experienced an increase in income inequality, and one (Italy) does not seem to experience any meaningful effect over the period studied. Most countries seem to experience an increase in income inequality in the later quarters (18-24). The effect is negligible for most countries on impact (t=0). Our results regarding the effect on gross Gini are different from Hafemann *et al.* (2018) who found that there was "a clear increase in the Gini gross after an expansionary monetary policy shock". Assuming the relationship between monetary policy and income inequality is linear, this would have translated into a clear decrease in the gross Gini after a contractionary shock, which is not the case in Figure 4.

One reason that could explain the heterogeneity across countries could come from the difference in the income composition of each country. More specifically, the differences in the proportion of household income derived from capital or labor income. Hafemann *et al.* (2018) demonstrate that the transmission of monetary policy shocks on income inequality is largely determined by income composition. They provide evidence that "an expansionary monetary policy shock leads to an increase in both, capital income and labor related income". In our case, this translates to a decrease in both from a contractionary shock (assuming a linear relationship). They also find that "capital income increases more than labor income" from an expansionary shock. According to them, the "income composition channel states that income inequality hikes, if capital income receivers benefit disproportionately, and vice versa." Considering that low-income households tend to hold fewer assets and tend to be more dependent on labor income, if the increase in unemployment (and the loss of labor income related to it) caused by a contractionary monetary policy shock surpasses the impact from a decrease in asset prices, it leads to an increase in income inequality, and vice versa. Additionally, Coibion *et al.* (2017) find that wages, business income and financial income behave differently to a contractionary monetary policy shock. They mention that "these results suggest that heterogeneity in income sources across households may also lead to important distributional consequences to monetary policy actions". Given these findings and those of Hafemann et al. (2018), the difference in impacts for each country may be due to heterogeneity in income sources across countries.

We find that five out of twelve countries in our sample experience an increase in income inequality as predicted by the theoretical model presented in section 3. As stated above, the share of capital and labor income of households may impact the trajectory of income inequality to a monetary policy shock. According to the findings of Hafemann *et al.* (2018), since capital income decreases more than labor income following a contractionary shock, higher-earning households may have a lower share of their income related to capital in countries that experience an increase in income inequality following a contractionary shock relative to countries that experience a decrease. The opposite may also hold: If higher-earning households are more reliant on capital income, the contractionary shock would lead to a decrease in income inequality. Bilbiie *et al.* (2022)'s model could be enhanced by adding capital accumulation as well as this additional weight in income composition. We would then be able to see how the model responds to different weights for the household with access to capital and comment on the validity of this hypothesis related to heterogeneity between countries based on the proportions of labor and capital income for households with access to capital markets.

Hafemann *et al.* (2018) mention that interpolation of income inequality data is a critical element to reflect on. Linear interpolation assumes that the rate of change is constant between yearly observations. Thus, if there was an actual variation within yearly observations generated by a previous or current monetary policy shock, we would not be able to capture this impact in our IRF after linearly interpolating Gini coefficients. In section 6 we use an alternative interpolation approach to establish if the choice of interpolation has considerable implications on the results. Unfortunately, our data sample is too small to use local projections for yearly observations to verify that the results are robust with yearly data, as done by Hafemann *et al.* (2018). As potential future research, a more sophisticated block exogeneity SVAR with Euro Area aggregated data and country-specific variables could be used to address this issue.

Another factor contributing to the ambiguity of the effect could be the short span of our data. This issue could be addressed by extending the AWM to add data for the period from 2018 to 2023, but this goes beyond the scope of this thesis, as the data is not currently available. This presents another opportunity for future research.

5.2 – Net Gini

We turn to the effects on net income inequality (after taxes and transfers) and how they compare to the gross Gini results. To do so, we re-evaluate the baseline approach by replacing the gross Gini with the net Gini (after tax and transfers). The impulse responses from the local projection using net Gini coefficients are given in Figure 5.

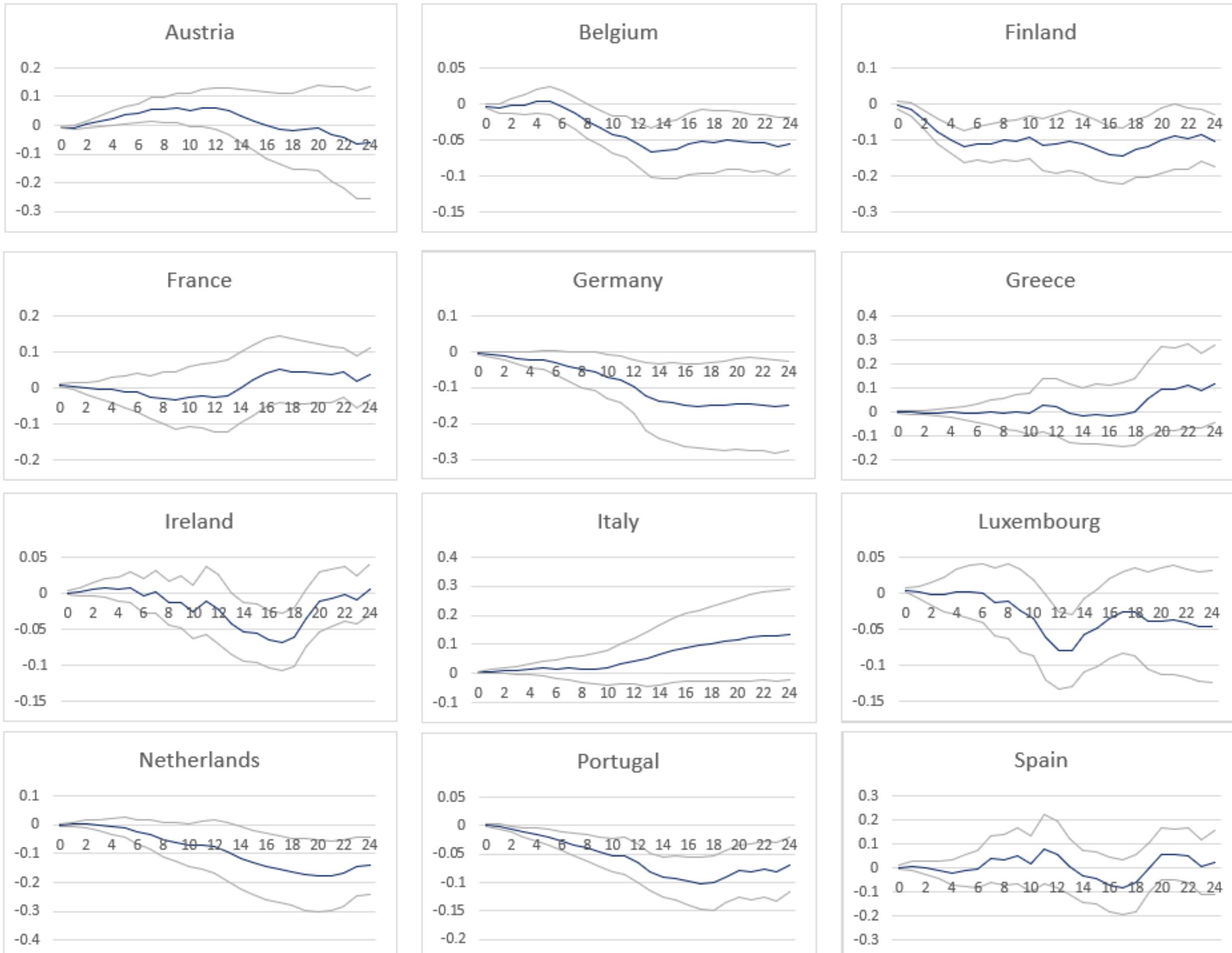

Figure 5 – Net Gini IRFs to 25-bps shock in monetary policy

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

Once again, we notice that there is important heterogeneity in the effects between countries which is in line with the heterogenous impact assessed by Liosi and Spyrou (2022) for the Eurozone.[13] Six out of twelve countries experience a decrease in net income inequality following a 25-bps contractionary monetary policy shock. Greece and Italy experience an increase in disposable income inequality. Austria, France, Spain, and Ireland present no clear effect. Ireland does experience an initial decrease, but this effect reverses after 18 quarters. France has a similar IRF as before (see Figure 4). Austria stands out as an intriguing case. Initially, when considering gross

[13] Unfortunately, they do not provide their estimated monetary policy shocks which would have been helpful to compare with ours.

income inequality, the IRF is stable for the first 8 quarters after impact but subsequently exhibits a downward trend (see Figure 4). With disposable income inequality, we notice an initial increase, which is followed by a decrease starting in the 12th quarter after the shock. Belgium, which initially saw a gradual increase in gross income inequality, now shows signs of a clear decrease in disposable income inequality as a consequence of its redistributive policies.

Since we are interested in how the IRF varies for countries with different degrees of redistribution, the following presents our sample of countries and other OECD countries relative to the cross-country mean in relative redistribution. This approach is derived from how Hafemann *et al.* (2018) classify countries by comparing each country to the cross-country mean of relative redistribution. Relative redistribution is the percentage reduction in income inequality as a consequence of fiscal policies and corresponds to the following equation:

$$\frac{[gross\ Gini - net\ Gini\ ]}{gross\ Gini} \times 100 \qquad \text{[Eq. 5.1]}$$

We take the same sample of countries as Hafemann *et al.* (2018) but include Portugal. We take the average relative redistribution of each country for the span of our study (1999 – 2017). This sample allows us to find a cross-country mean of 35.95 which is illustrated in Figure 6 and acts as the threshold for countries with a high degree of redistribution. This average is consistent with what Hafemann *et al.* (2018) obtain (i.e., slightly above 35).

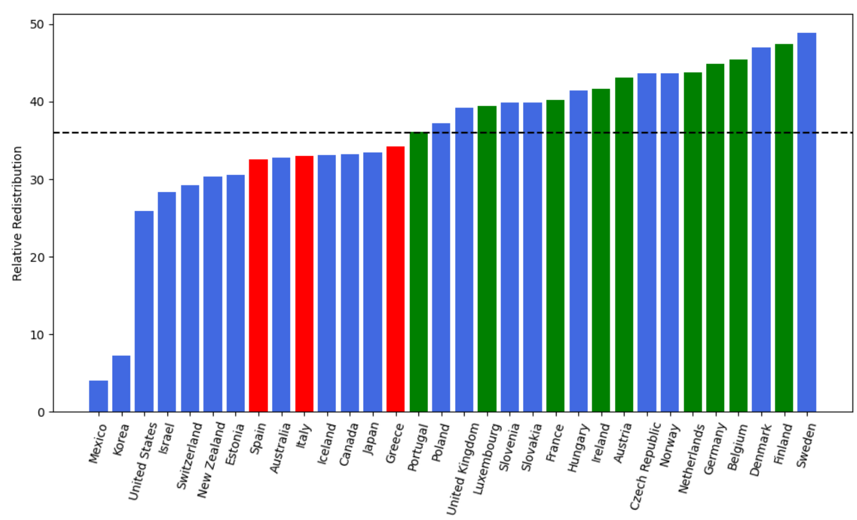

Figure 6 – OECD countries - average relative redistribution 1999-2017

Note: Average relative redistribution by country for our sample period using a similar sample of OECD countries as Hafeman et al. (2018).

Based on this classification, all countries in our sample except Spain, Italy and Greece have a high degree of redistribution. This is different from Hafemann *et al.* (2018) where they had half (3 out of 6 countries) of their sample that did not have a high degree of redistribution. One interesting observation is that 2 out of our 3 countries with low degrees of redistribution (Greece and Italy) experience a clear increase. These are the only countries that experience an increase in disposable income inequality.

While both Furceri *et al.* (2018) and Hafemann *et al.* (2018) found that countries with important redistributive policies experience no significant effect from monetary policy shocks, our results indicate otherwise. In our case, we find that most countries with high degrees of redistribution show a stronger reduction in income inequality following the same 25-bps contractionary monetary policy. These results are puzzling. Assuming that the relationship between monetary policy and income inequality is linear, this would mean that income inequality increases more after tax and transfers from an expansionary shock. These results contradict the existing literature on the subject, as well as economic theory (base on a closed-economy analysis). However, results could differ in the case of a monetary union due to the disconnect between fiscal and monetary policies. That is, the central bank could be decreasing the interest rate to fight inflation in the overall union, while at the same time specific countries with more transfer programs could be decreasing these programs given their own country-specific circumstances. Exploring these potential explanations in a theoretical model would be an interesting avenue for future research.

Figure A.6 in the appendix illustrates countries based on their average net and gross Gini coefficients over the period 1999-2017. Four out of five countries with the lowest average net income inequality are the ones for which the monetary policy shock induces the biggest reduction in income inequality, namely Finland, Netherlands, Belgium, and Luxembourg. Overall, our results are partly in line with those of Hafemann *et al.* (2018). Their results suggest that countries "with a high degree of redistribution show no clear pattern" when using net Gini coefficients, which is also the case here. The difference between their results and ours is that for us, countries with a high degree of redistribution experience either a clear decrease or no clear pattern. The only countries which experience a clear increase are countries with a low degree of redistribution. Once again, these results are not in line with the theoretical model derived from Bilbiie *et al.* (2022), which predicted an increase in income inequality following a contractionary monetary policy shock, except for Italy and Greece (two countries with a low degree of redistribution). This may be due to the ad-hoc nature of the model. To test the importance of these features, Bibliie *et al.* (2022)'s model could be expanded by adding endogenous redistribution policies as well as capital accumulation and the previously mentioned weights regarding the income source of higher-earning households. Finally, the model could be extended to a two-country framework to explicitly study the effects in a monetary union. This presents an opportunity for future theoretical research on the subject.

6 – Robustness checks

In this section, we perform robustness checks to assess the reliability of the results obtained in section 5. We investigate two key elements to ensure the validity of our findings: a different interpolation approach and a different ordering in our SVAR to identify monetary policy shocks. By adjusting these methodological choices, we aim to determine whether our baseline results are consistent and robust across alternative approaches.

6.1 - Proportional Denton's method of interpolation

We apply an alternative interpolation approach which will help determine if our results using linearly interpolated Gini coefficients are robust to the interpolation approach. Linear interpolation of Gini coefficients imposes by construction that the rate of change is constant between yearly observations, whereas there may be essential variations in the monetary policy shocks.

Denton's method was first introduced by Denton (1971). This method can be used to interpolate yearly data to quarterly or monthly observations using an indicator variable that matches the desired frequency. This allows us to capture variations between yearly observations based on the variation of the indicator variable. According to Chapter 6 of Bloem *et al.* (2001), the method is "relatively simple, robust, and well-suited for large-scale applications." We interpolate our Gini coefficients using Denton's method with real GDP as our high-frequency indicator variable (real GDP is available quarterly for each country from the FRED). This allows us to determine if using a different interpolation method yields different conclusions from our baseline approach. We use real GDP as our high-frequency indicator because it is readily available and is positively correlated with Gini coefficients.

We compare the original Gini coefficients and the interpolated series using Denton's method in Figures A.7 and A.8 in the appendix. We notice that the series correspond for most observations, while there is some smoothing of the curve between yearly data points. Interpolated values at Q4 of each year match the values of the original annual data points, confirming that the interpolation was done correctly.

To investigate if the slight difference between the two interpolation methods has considerable implications for the results of the IRF, we re-estimate our local projections. Figures 7 and 8 illustrate the new IRFs computed by local projections using the gross Gini coefficients interpolated using Denton's method.

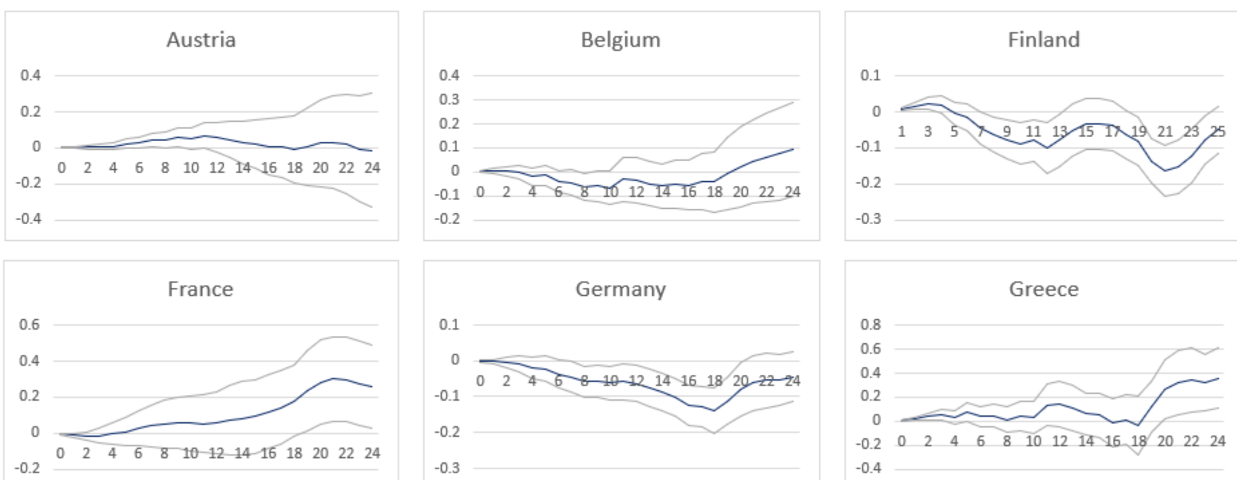

Figure 7 – Gross Gini IRFs by Local Projection using Denton's method

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

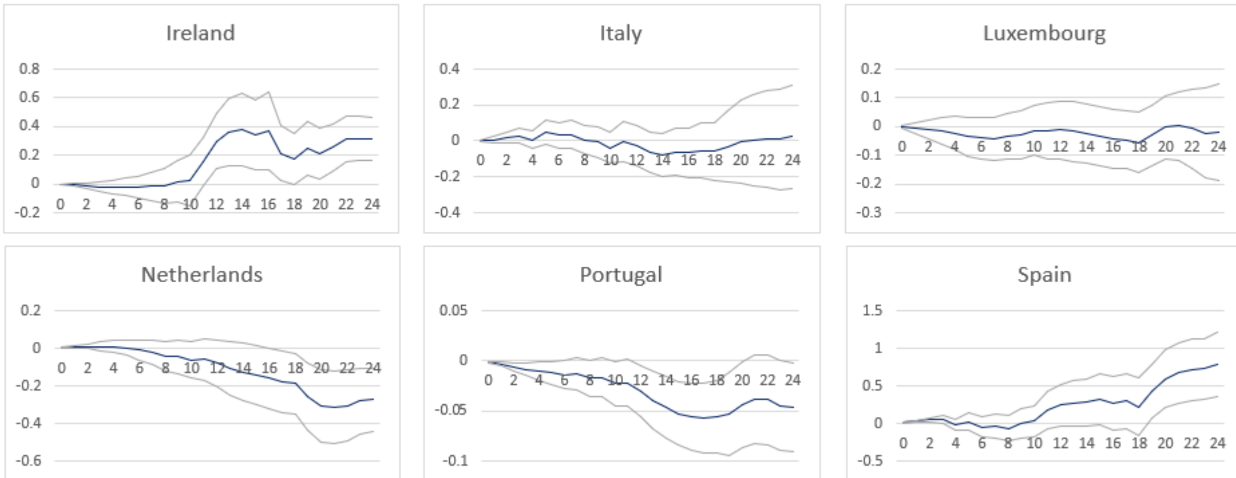

Figure 8 – Gross Gini IRFs by Local Projection using Denton's method

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

Most responses are similar to our baseline scenario (see Figure 4). Their patterns are very close to the approach using linear interpolation. However, some countries experience different IRFs than what we obtained in section 4. Austria, which initially saw a clear decrease over time from a contractionary shock, now experiences no evident effect. Belgium also experiences a different impact than when using linear interpolation. Previously, the shock had a strictly positive impact on gross income inequality, which is no longer the case. Under Denton's method, Ireland witnesses significant shifts in its IRF compared to the approach using linear interpolation, revealing that the initial analysis using linear interpolation failed to capture some considerable effect of a contractionary monetary policy shock. We notice an apparent increase in income inequality that starts ten quarters after the initial shock. There are differences for a few countries, but most IRFs exhibit similar patterns as in Figure 4. We notice differences in the magnitude of the IRFs, which can be explained by the smoothing of the Gini coefficients between yearly observations with Denton's method which increases the variation in income inequality following a shock.

Figure 9 presents the new IRFs computed by local projections using the net Gini coefficients interpolated using Denton's method.

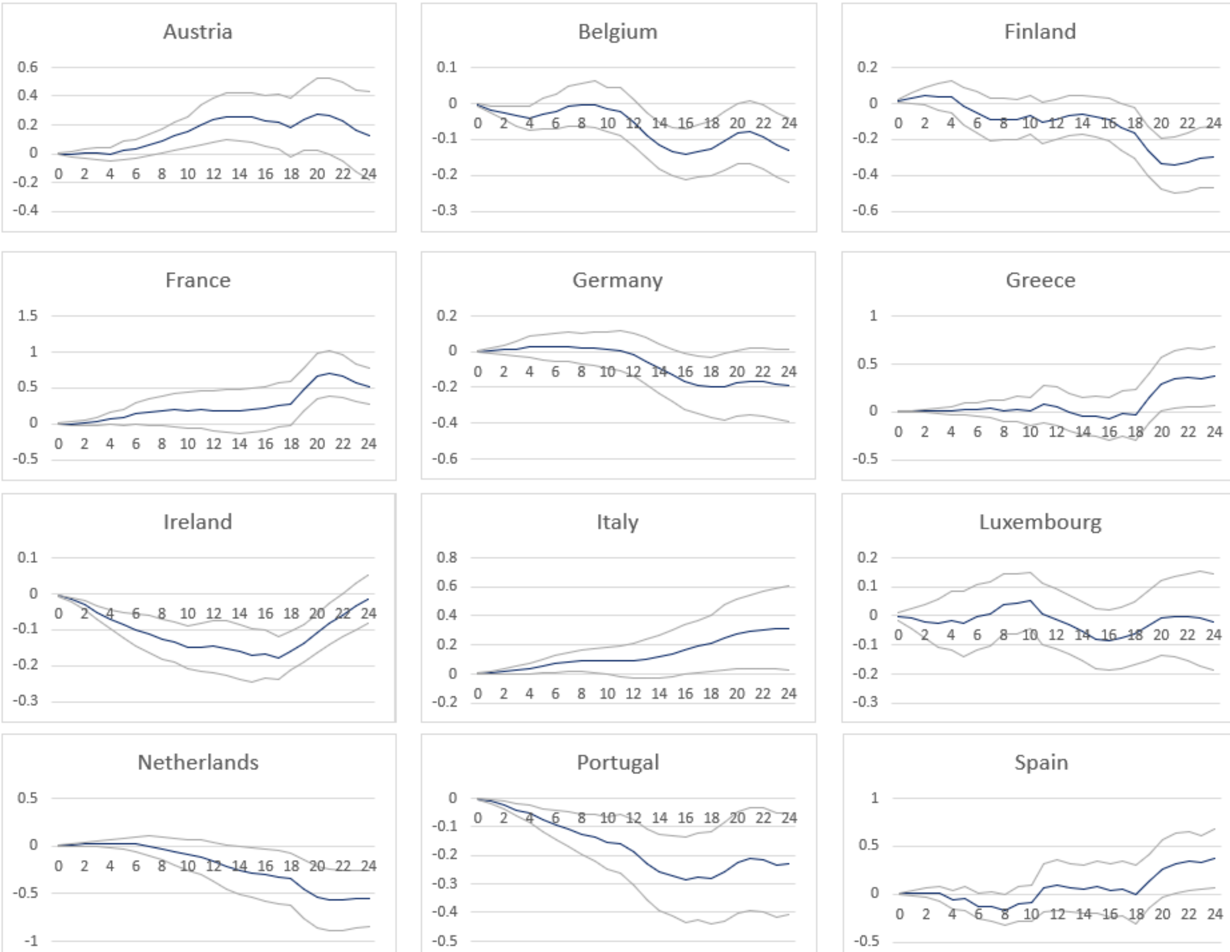

Figure 9 – Net Gini IRFs by Local Projection using Denton's method

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

Once again, the IRFs look similar to the ones we identified in Figure 5 using linear interpolation. Most patterns are almost identical to their counterpart with linearly interpolated Gini coefficients. Some cases worth noting are Austria, France, Luxembourg, and Spain. Austria, which initially experienced an increase followed by a slight decrease in income inequality, now shows a clear increase in income inequality following a contractionary shock. This goes against our original findings, where only countries with low redistribution (Greece and Italy) experienced increased income inequality following a 25-bps contractionary monetary policy shock. This is also the case for France, which initially saw a very slight increase in Figure 5 but now exhibits a clear increase.

Spain experiences something similar. It initially showed no significant impact in Figure 5 but now experiences a clear rise in income inequality. As Spain is considered a country with a low degree of redistribution, this is not surprising considering the findings in section 4, but the fact that Austria and France now show an increase in net income inequality while also being countries with a high degree of redistribution does contradict our original findings.

The results could be more robust to the interpolation choice. While most countries experience a similar effect from a 25-bps contractionary shock, there are some considerable differences in the magnitude of the impact and some considerable changes in the direction of the effect on some countries when considering net income inequality. The patterns of most countries are not affected other than in magnitude and usually stay within one standard deviation from the original impulse response. We conclude that the interpolation choice matters for countries which show important changes in their IRF. The methodological choice of interpolation has considerable implications regarding the impact of monetary policy shocks on income inequality. Using Denton's method, we obtain that even some countries with a high degree of redistribution show an increase in net income inequality, which is not the case using linear interpolation. This is more in line with the theoretical model in section 3 where an increase in the real interest rate increased income inequality, but only five out of twelve countries in our sample react this way using Denton's method. Six out of twelve still experience a decrease in income inequality from a contractionary monetary policy shock. The effect using Denton's method is more ambiguous than the effect using linear interpolation for net income inequality, highlighting once again the importance of the interpolation approach. We prefer Denton's method of interpolation because it allows for non-linear movement between yearly observations. This allows us to capture more movement as a response to a monetary policy shock between yearly observations, which is not the case with linear interpolation, which imposes the constant rate of change between yearly data points.

6.2 - Alternative SVAR ordering

Next, we explore the implications of using a different ordering in the SVAR. Since Hafemann *et al.* (2018) and Eichenbaum and Evans (1995) impose a different ordering for the first and second variables in the SVAR, we test if our findings differ when using this approach to estimate exogenous monetary policy shocks. Our baseline approach used the following ordering:

$$X_t = \begin{bmatrix} \mathrm{p}_t \\ \mathrm{y}_t \\ \mathrm{S}_t \end{bmatrix}$$

The alternative ordering will be used to test if our results are robust:

$$X_t = \begin{bmatrix} \mathrm{y}_t \\ \mathrm{p}_t \\ \mathrm{S}_t \end{bmatrix}$$

Using this ordering, we impose that real GDP is not impacted contemporaneously by prices, but prices are contemporaneously affected by real GDP. This ordering is consistent with the idea that changes in economic activity can influence price levels contemporaneously by influencing supply and demand dynamics. Rapid growth in economic activity leads to increased demand for products and services, potentially leading to price increases in the same quarter. We still impose that the shadow interest rate is impacted contemporaneously by the other two variables because central bankers have information on same-quarter inflation and GDP fluctuations and act accordingly. We use the same approach described in section 4 but change the ordering previously mentioned. Figure 10 compares both series of identified exogenous monetary policy shocks using the two different orderings.

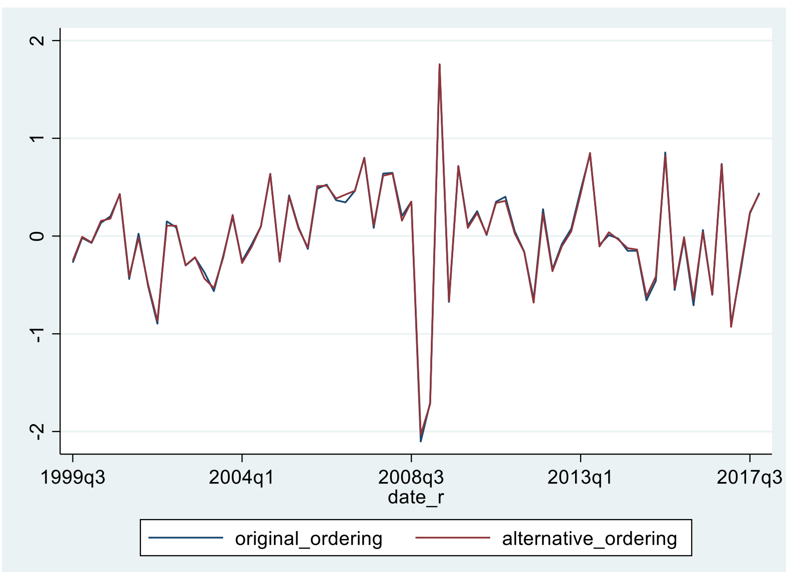

Figure 10 – Original vs alternative ordering MP shocks

Note: Comparison between the originally estimated monetary policy shocks and the monetary policy shocks estimated using the alternative ordering. The original ordering corresponds to placing inflation before production in the VAR. The Alternative ordering corresponds to placing production before inflation. Both series have quarterly data points and are identified using a Cholesky decomposition.

As we would expect, the two series are very similar. Next, we compute the IRFs by local projections for each country using our baseline model with gross Gini and net Gini coefficients to compare if these negligible differences in estimated monetary policy shocks have some noticeable impact on the IRFs. As for our measure of income inequality, we use the same linearly interpolated Gini coefficients used in our baseline model to compare the results with our baseline approach. Figure A.9 in the appendix presents the new IRFs computed by local projections using the gross Gini coefficients with the monetary policy shocks estimated using the alternative SVAR ordering. The IRFs all experience similar patterns as in Figure 4. There is no specific case worth mentioning. We notice slight differences in the magnitude of the IRFs, but nothing drastic.

Figure A.10 in the appendix presents the new IRF computed by local projections using the net Gini coefficients with the estimated monetary policy shocks using the alternative recursive ordering in the SVAR. Once again, we notice no important differences when comparing the IRFs estimated with those in Figure 5. The magnitude does change slightly, but most IRFs present matching patterns. Ireland is an exception, where we now notice a late increase, but this increase stays within one standard deviation from zero. We conclude that, although there are some minor variations in the magnitude of the IRFs, our baseline results are robust to the ordering of the variables in the SVAR for both gross and net Gini IRFs.

7 – Conclusion

This thesis explored the impact of monetary policy shocks on income inequality in the Eurozone. A thorough review of the current literature was performed to identify the relevant approaches and findings. A detailed explanation of the data sources and methodology used in this thesis was provided to ensure transparency regarding the methods employed to obtain the main results. We estimated monetary policy shocks using a SVAR approach with a Cholesky decomposition. The estimated shocks were used in local projections for each country to analyze country-specific effects.

The results indicate strong heterogeneity between countries in response to a 25-bps contractionary monetary policy shock. Our results partially align with the simple two-agent model from Bilbiie *et al.* (2022) presented in section 3, which predicted an increase in net income inequality from a contractionary shock. This model could be expanded by incorporating capital accumulation, endogenous redistribution, and different weights on the income sources of savers.

We performed two distinct robustness tests to verify the validity of our findings. The significant heterogeneity in the IRFs of the original results remained consistent regardless of the interpolation approach or the recursive ordering in the SVAR. While linear interpolation is more widely used, we prefer the results obtained by interpolating Gini coefficients using Denton's method. This approach allows for non-linear variations between yearly data points using real GDP as an indicator variable. The results demonstrate that using net Gini coefficients leads to lower increases and larger decreases in IRFs, contradicting the results of Furceri *et al.* (2018) and Hafemann *et al.* (2018). This was the case for both linear interpolation and interpolation using Denton's method. We believe that this difference may arise from the context of a monetary union given that this is the essential difference between the approach in this thesis and these two studies. This difference may also come from the estimation of monetary policy shocks for a monetary union which may differ from countries with their own central bank.

Regarding future research avenues, updating the AWM would allow for a longer sample period to be considered. Additionally, estimating ECB monetary policy shocks using an approach based on forecasting similar to that used by Furceri *et al.* (2018) or Coibion *et al.* (2017) would allow to compare with the results obtained in this thesis and determine how important is the choice of the method for shock identification in the context of a monetary union. Adding another measure of income inequality to the analysis in this thesis could be considered for future research on the subject. Finally, investigating the different mechanisms through which monetary policy impacts income inequality could enable the ECB to incorporate income inequality in their models in order to promote long term growth.

8 - Appendix

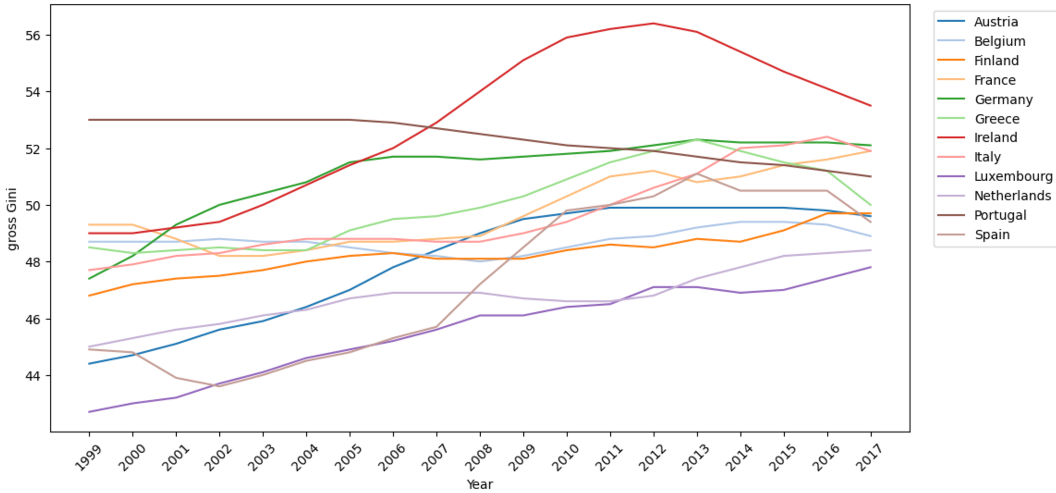

Figure A.1 – Gross Gini coefficients 1999-2017

Note: Gross Gini coefficients by country for our sample period (1999 to 2017). Data retrieved from SWIID (https://fsolt.org/swiid/).

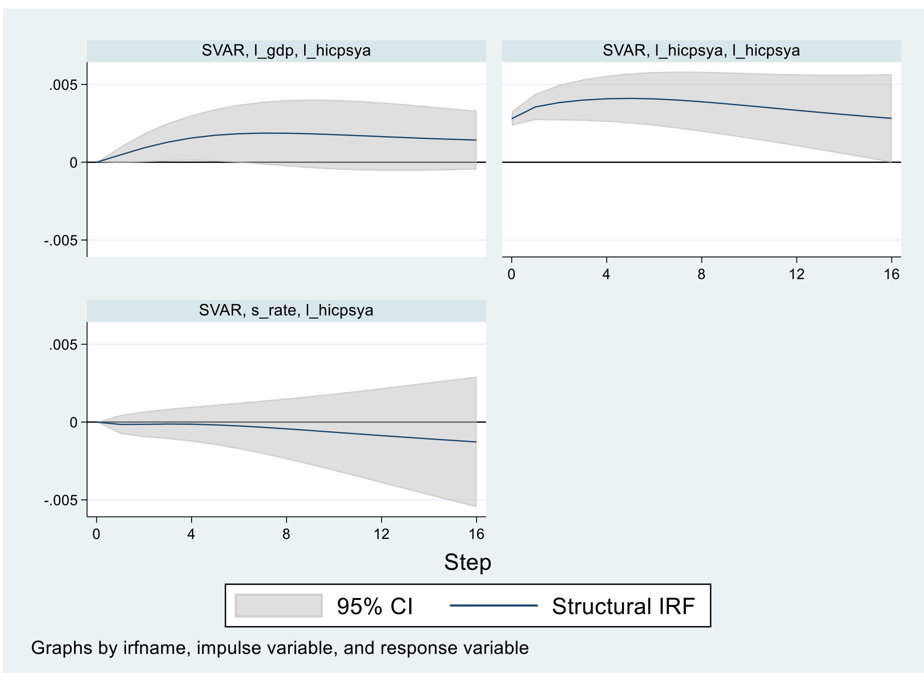

Figure A.2 – Inflation impulse-response (quarterly)

Note: Impulse response functions of (log) HICPSYA to a one-standard-deviation shock in (log) real GDP, (log) HICPSYA and shadow interest rate. The X-axis represents the number of quarters after the shock. The Y-axis represents the impact on (log) HICPSYA.

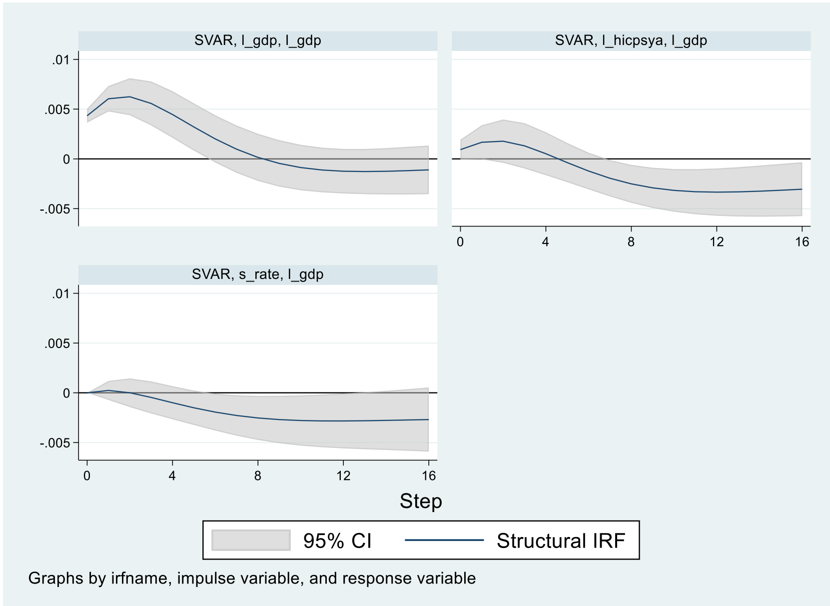

Figure A.3 –real GDP impulse-response (quarterly)

Note: Impulse response functions of (log) real GDP to a one-standard-deviation shock in (log) real GDP, (log) HICPSYA and shadow interest rate. The X-axis represents the number of quarters after the shock. The Y-axis represents the impact on (log) real GDP.

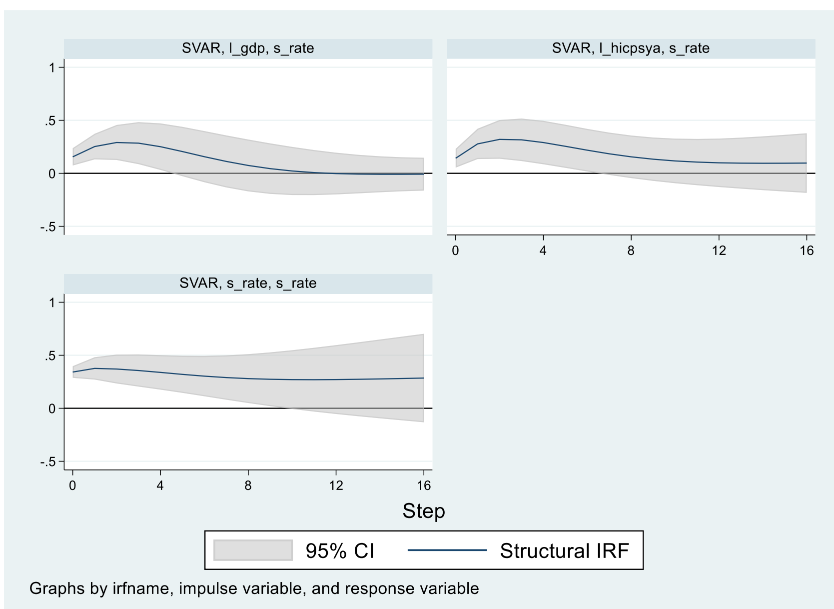

Figure A.4 – Shadow Interest Rate impulse-response (quarterly)

Note: Impulse response functions of the shadow interest rate to a one-standard-deviation shock in (log) real GDP, (log) HICPSYA and shadow interest rate. The X-axis represents the number of quarters after the shock. The Y-axis represents the impact on the shadow interest rate.

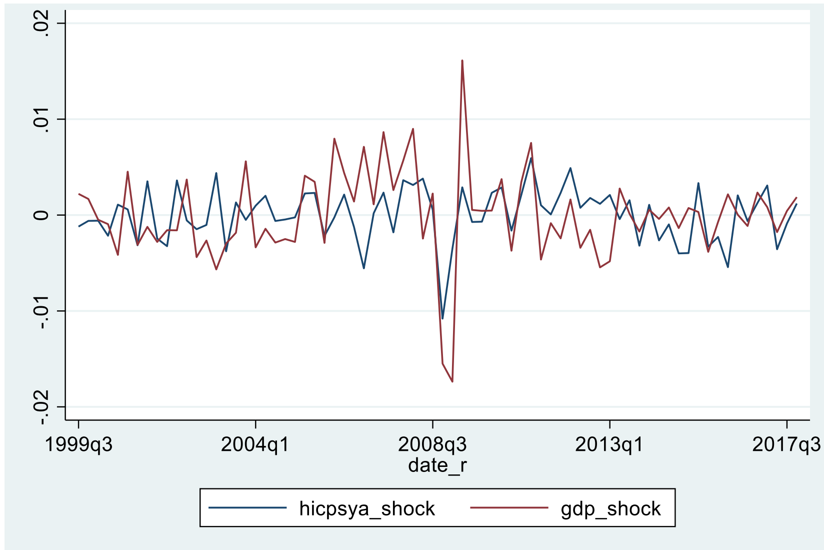

Figure A.5 – Inflation and real GDP shocks

Note: Identified shocks of (log) real GDP and (log) HICPSYA using a Cholesky decomposition over our sample period.

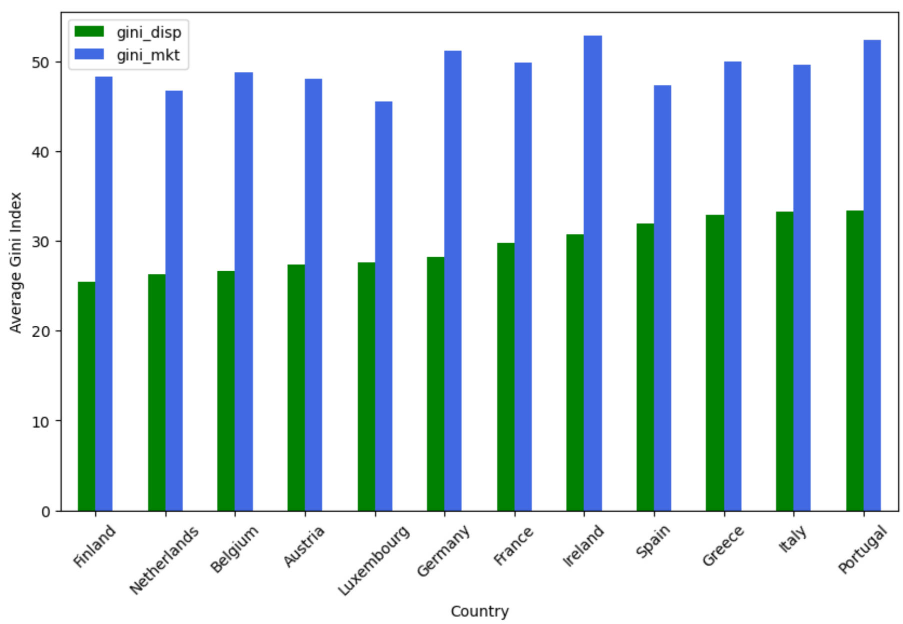

Figure A.6 - Average Gini for sample countries (1999-2017)

Note: Average gross (gini_mkt) and net (gini_disp) Gini coefficients over our sample period. Data retrieved from SWIID (https://fsolt.org/swiid/).

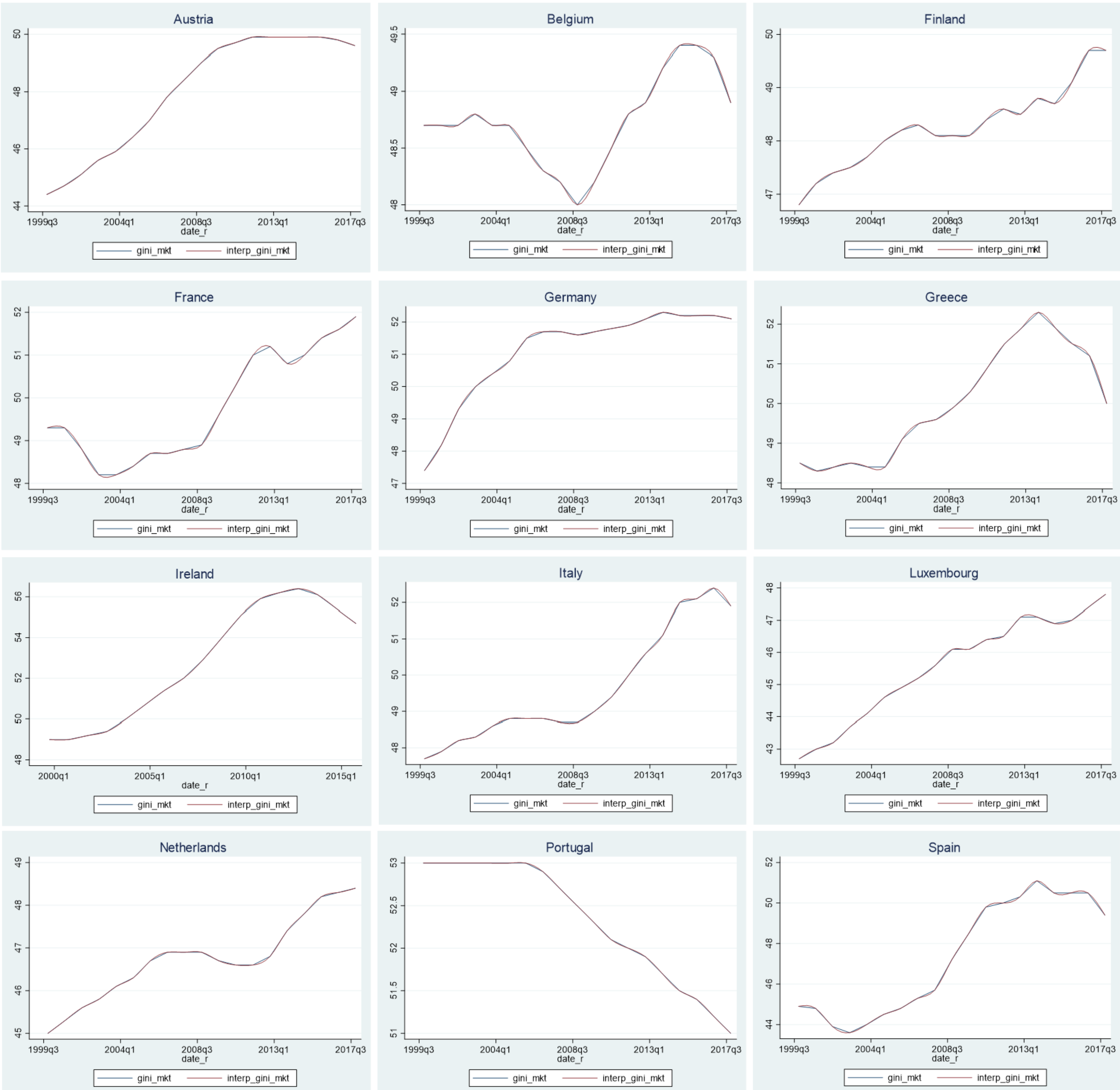

Figure A.7 – Gross Gini Interpolated using Denton's method

Note: Gross Gini coefficients for each country interpolated using Denton's method of interpolation with real GDP as our indicator variable. All series are similar to their non-interpolated counterpart, with some smoothing of the curves from quarterly interpolation.

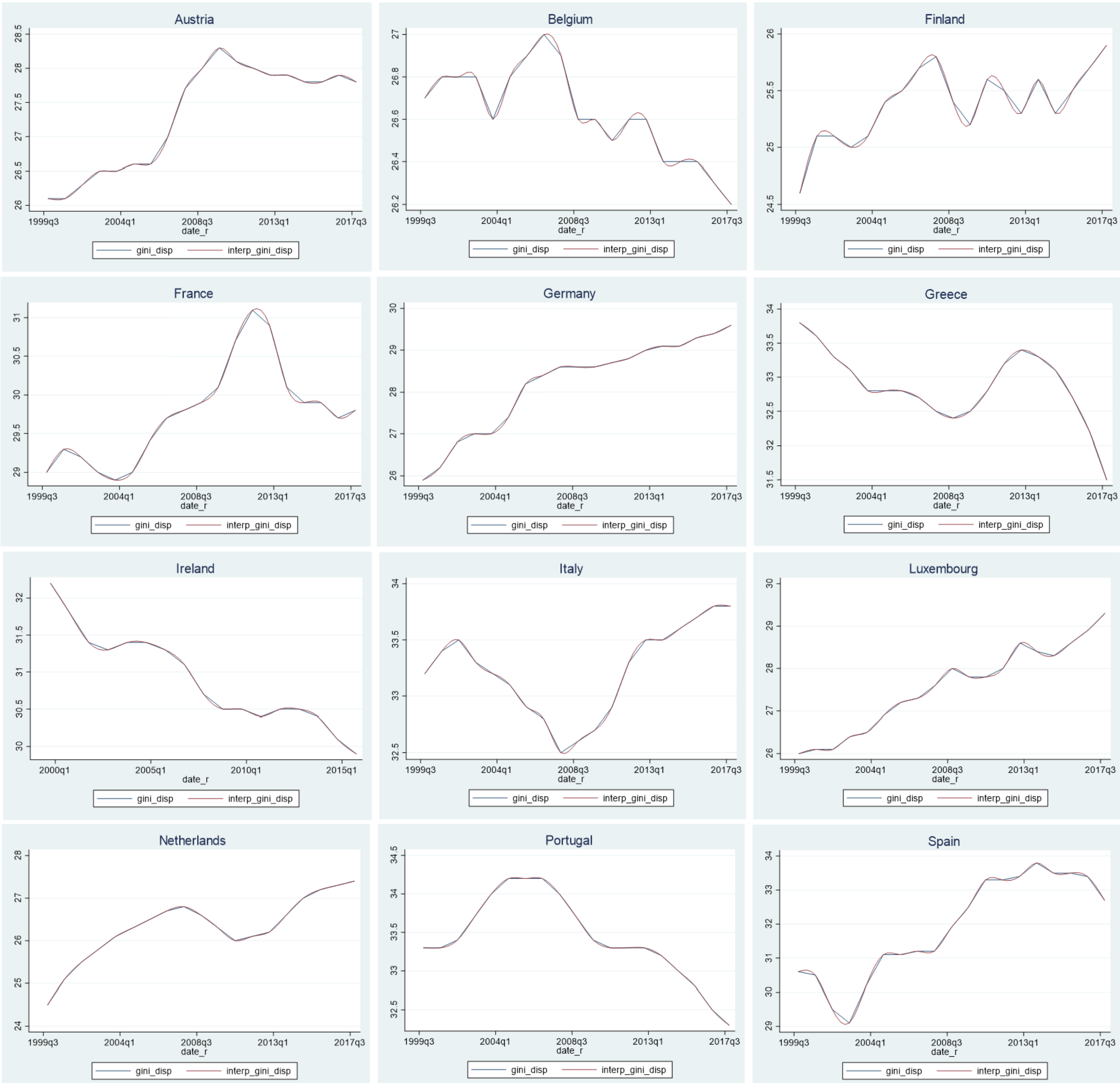

Figure A.8 – Net Gini Interpolated using Denton's method

Note: Gross Gini coefficients for each country interpolated using Denton's method of interpolation with real GDP as our indicator variable. All series are similar to their non-interpolated counterpart, with some smoothing of the curves from quarterly interpolation.

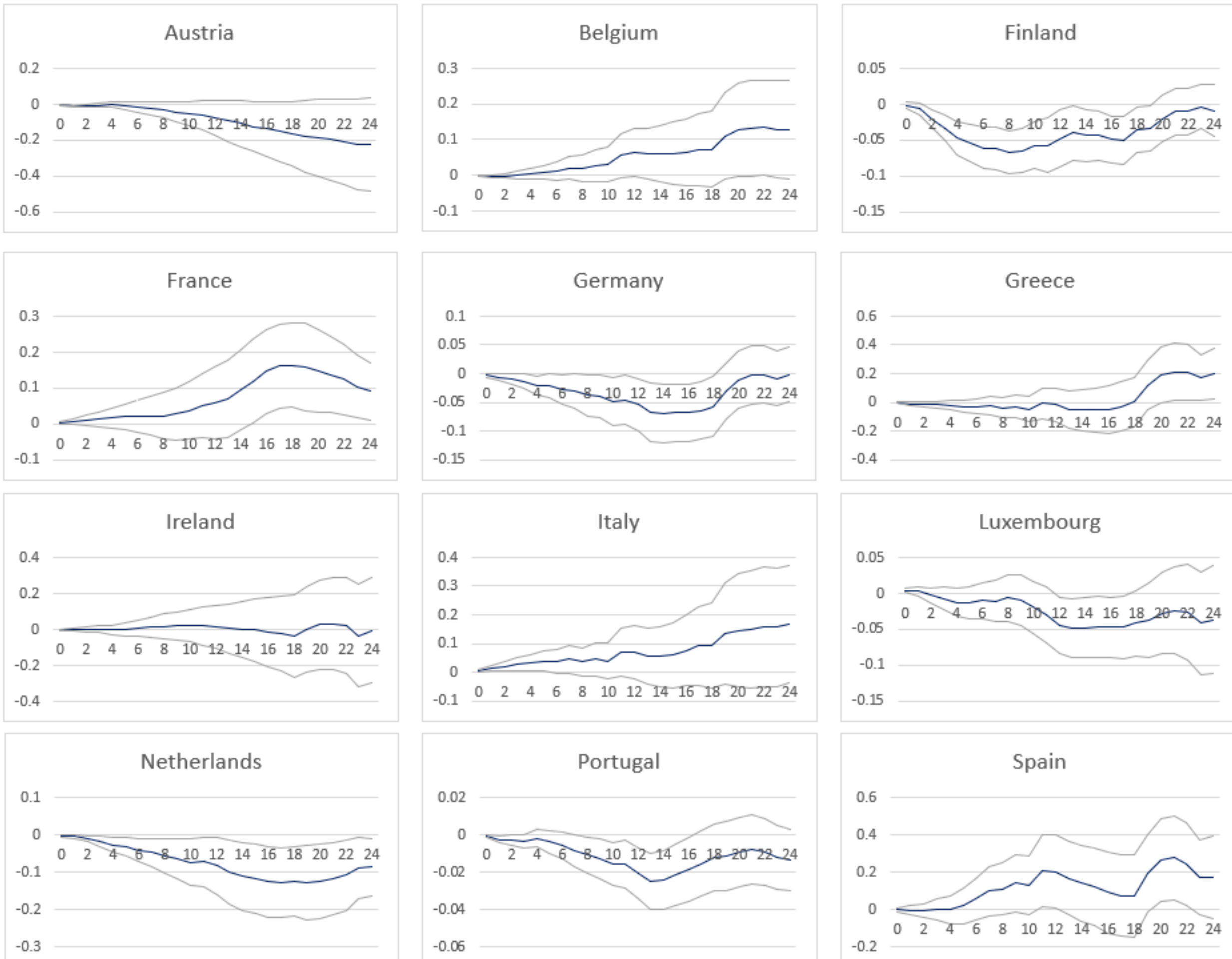

Figure A.9 – Gross Gini IRFs by Local Projection using alternative ordering

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

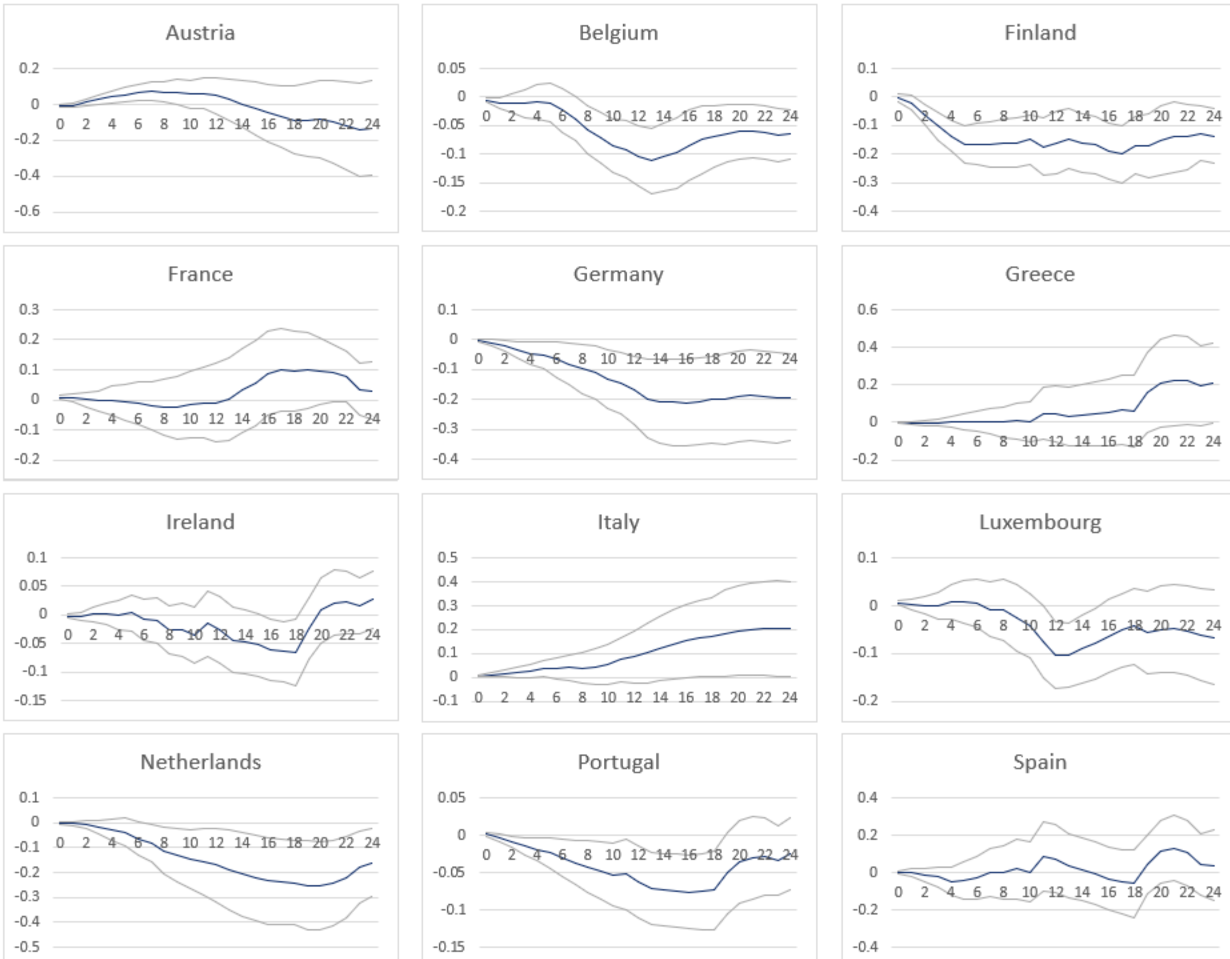

Figure A.10 – Net Gini IRFs by Local Projection using alternative ordering

Note: The X-axis represents the number of quarters since the 25-bps shock. The Y axis is the change in percentage points in income inequality since the shock occurred (t=0). The gray bands represent a one standard deviation error band. Notice that the Y axis varies from country to country (the magnitude of the IRF differs for each country).

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