The Distributive Effect of Monetary Policy:
The Top One Percent Makes the Difference

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Abstract

The paper evaluates the distributional effect of monetary policy. The empirical analysis is implemented for the USA, where the dynamics in income inequality is mainly driven by the variation in the top one percent of the income distribution. The paper uses the inequality measures that represent the whole income distribution. The distributive effect of monetary policy is evaluated in the cases of different frequency data. To identify a monetary policy shock, the paper applies the contemporaneous and the long run identification methods. In particular, a cointegration relation is determined among the considered variables and the vector error correction methodology is used for the identification. The obtained results indicate that contractionary monetary policy decreases income inequality. These results can have important implications for the design of policies to reduce income inequality by giving more weight to monetary policy.

JEL classification: C32, D31, E52
Keywords: income inequality; monetary policy; cointegration; identification

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1. Introduction

Nowadays there are widespread concerns regarding growing income inequality and different fiscal policy measures are discussed to address it. However, monetary policy can also affect the distribution of income although its distributive effect is not extensively discussed. The objective of the paper is to contribute to this discussion by evaluating the effect of monetary policy on income inequality.

Distributive mechanisms are usually described through political economy arguments that specify some transmission channels between income inequality and economic growth (Acemoglu and Robinson, 2008; Benabou, 2000; Muinelo-Gallo and Roca-Sagales, 2011; Neves and Silva, 2014). According to these arguments, the distribution of income is implied to be implemented through fiscal policy. However, income is distributed also via monetary policy. Economic activities are regulated by macroeconomic policies, which include both types of policies. Though fiscal and monetary policies are used for comparatively different macroeconomic objectives (commonly to increase aggregate output and to control inflation, respectively), they also affect the same economic activities, such as the distribution of income.

Monetary policy can affect the income distribution through different transmission mechanisms. Inflation has a direct effect on income inequality through changes in the real valuation of financial and non-financial assets. In the case of the USA, studies show that inflation hits richer and older households whose asset holdings are typically imperfectly insured against surprise inflation (Doepke and Schneider, 2006; Doepke et al., 2015). Inflation is especially harmful for the poorest parts of the population. This is because poorer households tend to hold a larger fraction of their financial wealth in cash, implying that both expected and unexpected increases in inflation make them even poorer. Moreover, high inflation can create expectations of future macroeconomic instability and lead to distortionary economic policies (Romer and Romer, 1999). According to Bulir (2001), preceding inflation raises income inequality in following periods. As Albanesi (2007) demonstrates, a higher inflation rate is accompanied by greater income inequality. Accordingly, Villarreal (2014) shows that contractionary monetary policy decreases income inequality in Mexico. On the contrary, Coibion et al. (2012) find that contractionary monetary policy tends to raise economic inequality in the USA.

The estimated effects of monetary policy could depend on the representativeness of the inequality measures used in the empirical analysis. That is, the estimated effects might differ if they do not represent the whole income share of population, particularly the top one percent. In the USA, the dynamics of income inequality is mainly driven by the variation in this upper end of the distribution (Atkinson et al. 2011; Congressional Budget Office, 2011; Kenworthy and Smeeding, 2013). Moreover, it also affects the top shares of the world income distribution and, consequently, the world income inequality (Atkinson, 2007) since top income shares can be a proxy for inequality across the distribution (Leigh, 2007). Therefore, the paper evaluates the distributional impact of monetary policy in the USA by using the inequality measures that cover the whole income distribution, including the top one percent.

The distributive effect of monetary policy is evaluated in the cases of quarterly and annual data. Correspondingly, the paper applies the contemporaneous and the long run identification of a monetary policy shock. In the latter case, the vector error correction methodology is used for the identification of a monetary policy shock since the paper finds a cointegration relation among the considered variables. The obtained results show that contractionary monetary policy reduces the overall income inequality.

The rest of the paper is organized as follows. Section 2 reviews the related literature and the distribution channels. Section 3 discusses the empirical methodology. Section 4 describes the data and Section 5 provides the results. Section 6 contains the concluding remarks.

2. Literature Review and Distribution Channels

2.1. Literature Review

There are not many empirical papers devoted to the examination of the effect of monetary policy on income inequality in academic literature (Coibion et al., 2012; Saiki and Frost, 2014; Villarreal, 2014). The distributive impact of fiscal policy has been considered in the literature (among others, Afonso et al., 2010; Doerrenberg and Peichl, 2014; Wolff and Zacharias, 2007) more than the distributive effect of monetary policy. Nevertheless, there are some insightful papers discussing different aspects of distributive effects of monetary policy and they are discussed thoroughly below.

Using cross-country data, Bulir (2001) provides evidence that preceding inflation raises income inequality in following periods. He argues that the total impact of inflation on inequality takes some time to be revealed. His analysis indicates that the positive effect of price stability on income inequality is nonlinear. That is, the initial decline in hyperinflation substantially reduces inequality whereas the further effects of the reductions in lower levels of inflation consecutively decrease. Bulir (2001) concludes that price stabilization is beneficial for reducing income inequality not only via its direct effect but also indirectly through boosting money demand and preserving the real value of fiscal transfers.

Using cross-country panel data, Li and Zou (2002) find that inflation deteriorates income distribution and economic growth. They also show that inflation increases the income share of the rich and insignificantly reduces the income shares of the middle class and the poor.

Albanesi (2007) provides cross-country evidence of positive correlation between inflation and income inequality. She also builds a political economy model in which income inequality is positively related to inflation in equilibrium because of a distributional conflict in the determination of fiscal and monetary policies. The model implies that in equilibrium low income households have more cash as a share of their total consumption, in line with empirical evidence (Erosa and Ventura, 2000). Therefore, low income households are more exposed to inflation. Particularly, Easterly and Fischer (2001) bring empirical evidence, using data from 38 countries that the poor are more probably than the rich to indicate inflation as a top national concern. The model built by Albanesi (2007) also implies that households with more income have a greater power in the political process. As a result, for the government it is easier to finance its spending through positive seigniorage than via increased taxation, which requires parliamentary approval. Thus, according to Albanesi (2007), this leads to inflation in equilibrium and to its positive relation with income inequality.

Romer and Romer (1999) consider the influence of monetary policy on poverty and inequality in the short run and the long run. Using single equation time series evidence for the USA, they find that expansionary monetary policy is associated with better conditions for poor (decreased inequality) in the short run. On the contrary, examining the cross-section evidence from a large sample of countries, Romer and Romer (1999) show that tight monetary policy resulting in low inflation and stable aggregate demand growth are associated with the enhanced well-being of the poor (reduced inequality) in the long run.

Galli and von der Hoeven (2001) claim that there is a non-monotonic long run relationship between inflation and income inequality. Particularly, they argue that the relationship is U-shaped – inequality declines as inflation rises from low to moderate rates but inequality increases when inflation further grows from moderate to high levels. Their empirical analysis is implemented for the USA and a sample of 15 OECD countries.

For the USA, Galbraith et al. (2007) show that, earnings inequality in manufacturing is influenced by monetary policy. The latter is captured by the yield curve measured as the difference between 30-day Treasury bill and 10-year bond rate. They find that the earnings inequality is directly influenced by monetary policy in addition to indirectly being affected by inflation and unemployment, and by recessions in general. In particular, Galbraith et al. (2007) indicate that tight monetary policy raises the inequality of earnings while expansionary monetary policy reduces it.

The Bank of England (2012) states that while, through unconventional monetary policy measures, it could overcome the financial crisis, these measures might also increase income inequality. The Bank of England (2012) has implemented unconventional monetary policy almost entirely through the purchases of gilts. The implementation of this unconventional monetary policy has also increased the prices of other assets, such as corporate bonds and equities. As a result, this has raised the value of the financial wealth of households who hold them, and the owners can gain capital income by selling the assets. Consequently, it can also increase income inequality because the top 5% of households possesses 40% of the assets (the Bank of England, 2012).

Coibion et al. (2012) provide evidence that monetary policy shocks account for a significant component of the historical variation in economic inequality in the USA. Their measures of economic inequality are based on the Consumer Expenditures Survey, which does not include the top one percent of the income distribution. They show that contractionary monetary policy raises inequality in labor earnings, total income, consumption, and total expenditures. In particular, the results show that the shock most significantly affects expenditure and consumption inequality. Coibion et al. (2012) also explores different channels through which monetary policy affects economic inequality.

For Korea, Kang et al. (2013) find that inflation improves economic inequality in the short run but it has no significant impact on inequality in the long run. They also show that GDP growth decreases economic inequality. Their results indicate that there is no significant relation between real interest rate and inequality though real interest rate and poverty are positively correlated.

Saiki and Frost (2014) provide evidence that unconventional monetary policy raises income inequality in Japan in the short run. In particular, they show that by increasing the monetary base, unconventional monetary policy widens income inequality through resulting higher asset prices, benefiting the rich who usually hold these equities and acquire capital gains. Saiki and Frost (2014) conclude that while unconventional monetary policy tends to help to overcome the global financial crisis, it could have a side effect in terms of increased income inequality.

Villarreal (2014) shows that contractionary monetary policy decreases income inequality in Mexico. He uses different identification schemes for monetary policy shocks. Generally, all his results indicate that an unanticipated increase in nominal interest rate reduces income inequality over the short run. Villarreal (2014) interprets the differences of his results for Mexico from the ones obtained by Coibion et al. (2012) for the USA by the existence of such a level of financial frictions in Mexico that the benefits of inflation stabilization are higher than its costs.

For the UK, Mumtaz and Theophilopoulou (2016) find that contractionary monetary policy increases economic inequality while quantitative easing raises it. In the baseline case, they apply sign restrictions for the identification of a monetary policy shock. Mumtaz and Theophilopoulou (2016) use survey data to construct inequality measures, which are trimmed by cutting out the top and the bottom one percent of their distributions. Nevertheless, the dynamics of income inequality is mainly driven by the variation in the top one percent in the UK (Atkinson et al., 2011; Equality Trust, 2013; Leigh, 2007). This feature is also attributable to the case of the USA (Atkinson, 2007; Congressional Budget Office, 2011; Kenworthy and Smeeding, 2013) since the behavior of income inequality is similar in both the UK and the USA.

2.2. Distribution Channels

The overall distributive impact of monetary policy depends on the different channels through which monetary policy can affect income inequality. Coibion et al. (2012) classify five such channels, which are also considered by other authors (e.g., Saiki and Frost, 2014). These channels are the following:

1. The income composition channel refers to the heterogeneity in primary sources of income across households. Many households depend mainly on wages whereas others acquire their income from business and financial gains. So, if expansionary monetary policy increase profits more than labor earnings, the owners of assets and firms benefit more. Taking into account that they are usually wealthier, expansionary monetary policy shocks might lead to higher income inequality via this channel.

2. The financial segmentation channel implies the reallocation of income towards the agents involved in financial markets who can benefit from expansionary monetary policy shocks. Considering the fact that these agents generally earn more income than the agents not engaged in financial markets, expansionary monetary policy would raise inequality through this channel.

3. The redistribution of income based on the structure of owned assets is represented by the portfolio channel. Normally, low income households have mainly currency whereas upper income households tend to possess various securities. Therefore, by causing inflation and financial market booms, expansionary monetary policy would harm low income households and benefit upper income households via this channel, leading to the increase in inequality.

4. The impact of unexpected inflation on nominal contracts is expressed by the savings redistribution channel. The unexpected increase in inflation would benefit borrowers and would hurt savers. Considering that usually savers are wealthier than borrowers, expansionary monetary policy shocks would reduce inequality through this channel.

5. The earnings heterogeneity channel describes the tendency that the labor income of the poorest population is mostly exposed to business cycle fluctuations. At the same time, low income households usually receive a bigger share of their income from government transfers, and government transfers are normally countercyclical. Generally, expansionary monetary policy might decrease income inequality via this channel.

Thus, monetary policy could have different distributional effects through these channels. As mentioned above, through the first three channels, expansionary monetary policy increases income inequality and reduces it via the last two channels. That is, the total distributive impact of monetary policy is not certain. This is also pointed out by O'Farrell et al. (2016). The objective of the paper is to assess the overall effect of all the channels.

Talking into account that monetary policy affects as prices as well as real economic activity[1], Nakajima (2015) specifies two general distributive channels of monetary policy: inflation and income channels. They incorporate the channels specified by Coibion et al. (2012). Inflation channel contains the financial segmentation channel, the portfolio composition channel, and the savings redistribution channel. Income channel includes the income composition channel and the earnings heterogeneity channel. Considering these aggregated channels, the paper uses prices and real output as the general distributive channels of monetary policy[2]. As a monetary policy tool, the federal funds rate is employed. Besides, these three variables are commonly incorporated in monetary policy models (Bernanke and Mihov, 1998; Christiano et al., 1996; Peersman and Smets, 2001; Uhlig, 2005). To assess the overall distributional effect of monetary policy, a measure of income inequality is also included in the analysis.

The paper aims to contribute to the existing literature on the valuation of the distributive effect of monetary policy by employing different identification methods for a monetary policy shock and by using an inequality measure that represents the whole income distribution. In particular, the paper detects a cointegration relation among the considered variable and it applies the vector error correction methodology for the identification of a monetary policy shock. This is a novel approach for the evaluation of the distributive effect of monetary policy. Also, the paper considers consistently measured data on income inequality that cover the top one percent of income distribution, complementing the work by Coibion et al. (2012). As the results show, the representativeness of the inequality measure has substantial impact on the evaluation of the distributive effect of monetary policy.

3. Empirical Methodology

The examination of the distributional effects of monetary policy is implemented through multiple time series analysis, which allows tackling the endogeneity problem among the variables and studying their interrelations. Two equivalent representations of multiple time series models are considered in the paper. One of the employed models is the vector autoregression, which for order p, VAR(p), can be expressed as following[3]:

$$y_t = A_1 y_{t-1} + \cdots + A_p y_{t-p} + u_t \tag{1}$$

where y_t is the vector of endogenous variables, A_is are (4×4) coefficient matrices and $u_t = (u_{1t}, \cdots, u_{4t})'$ is an error term. It is assumed that the error term is a zero-mean independent

[1] The mandate of the Federal Reserve includes the promotion of maximum employment.
[2] The considerations of the variables for the empirical analysis are analogous to Paper 3.
[3] The notations are in line with the representations used by Lütkepohl (2005).

white noise process with positive definite covariance matrix $E(u_t u_t') = \Sigma_u$. That is, the error term is an independent stochastic vector with $u_t \sim (0, \Sigma_u)$. In the specification of the model, the vector of endogenous variables y_t includes real output, prices, the federal funds rate, and an income inequality measure: $y_t = (Y_t, P_t, R_t, Z_t)'$. In the baseline case, y_t consists of real GDP, GDP deflator, the federal funds rate, and Gini index.

In the paper, the other considered model is the equivalent representation of VAR(p) in the case of the cointegrated variables. It is the vector error correction model of order *p-1*, VECM(p-1):

$$\Delta y_t = \Pi y_{t-1} + \Gamma_1 \Delta y_{t-1} + \cdots + \Gamma_{p-1} \Delta y_{t-(p-1)} + u_t \tag{2}$$

where Δy_t denotes the first order differences of y_t, $\Gamma_i = -\left(A_{i+1} + \cdots + A_p\right)$ for $i = 1, \dots, p-1$, $\Pi = -\left(I_4 - A_1 - \cdots - A_p\right)$, where I_4 is an identity matrix of dimension 4. The rank of $\Pi = \alpha\beta'$ equals to the number of cointegration relations *(r)*. α and β are matrices of loading and cointegration parameters, respectively. The term $\alpha\beta' y_{t-1}$ is the long run part, and $\Gamma_j s$ $(j = 1, \dots, p-1)$ are short run parameters.

Analogously, it is possible from the parameters of VECM(p-1) to determine the coefficients of VAR(p):

$$A_1 = \Gamma_1 + \Pi + I_4,\ A_i = \Gamma_i - \Gamma_{i-1} \text{ for } i = 2, \dots, p-1;\ A_p = -\Gamma_{p-1} \tag{3}$$

In both cases, deterministic terms could be included in the models as following:

$$y_t = \mu_t + x_t \tag{4}$$

where μ_t is a deterministic part and x_t is a stochastic process that can have a VAR or VECM representation. As a deterministic part could be such terms as a constant, a linear trend, or dummy variables.

If the variables are stationary, the process (1) can be represented through an infinite order polynomial in the lagged values of the innovation u_t by a moving average representation:

$$y_t = \Phi_0 u_t + \Phi_1 u_{t-1} + \Phi_2 u_{t-2} + \cdots \tag{5}$$

where Φ_0 is an identity matrix and $\Phi_s = \sum_{j=1}^{s} \Phi_{s-j} A_j$, $s = 1, 2, \dots$, can be computed recursively from the reduced form coefficients of (1). The elements of the matrices Φ_j are the impulse responses of the components of the vector y_t with respect to the innovations of the vector u_t. In this case when the vector y_t is stationary, the values of the matrices Φ_j converge to zero as *j* gets large.

In general, the reduced form disturbances are the linear combinations of structural innovations:

$$u_t = B\varepsilon_t \tag{6}$$

where ε_t is a (4×1) vector of orthogonal shocks and B is a (4×4) matrix of parameters. That is, $4^2 = 16$ parameters are required for identification. $\frac{4^2}{2} + \frac{4}{2} = 10$ restrictions are given by estimation. $\frac{4(4-1)}{2} = 6$ restrictions are necessary for just identification and they require out of sample information. The vector of the variables y_t can be expressed through structural shocks by substituting (6) in (5): $y_t = B\sum_{j=0}^{\infty} \Phi_j \varepsilon_{t-j}$.

One of the most commonly used identification approaches is Cholesky decomposition. Based on this approach, the following contemporaneous restrictions are imposed on the matrix B:

$$\begin{pmatrix} u_Y \\ u_P \\ u_R \\ u_Z \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 \\ b_{21} & 1 & 0 & 0 \\ b_{31} & b_{32} & 1 & 0 \\ b_{41} & b_{42} & b_{43} & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_Y \\ \varepsilon_P \\ \varepsilon_R \\ \varepsilon_Z \end{pmatrix} \tag{7}$$

Real output, prices, and the federal funds rate are commonly incorporated in monetary VAR models and the contemporaneous restrictions among them are placed following the related literature (Christiano et al., 1996; Peersman and Smets, 2001). In particular, it is assumed that the policy rate has no contemporaneous impact on output and prices. However, it contemporaneously responds to changes in output and prices. In addition, it is assumed that income inequality does not contemporaneously affect the policy rate and the other variables while all of them have contemporaneous impact on it. All these restrictions require the usage of such a frequency of data for estimations that the assumptions behind the restrictions are plausible. As it is later stated in the next section, the data on income inequality are available only on an annual basis. The assumptions for the restrictions related to the interactions between income inequality and the other variables can still be realistic in the case of the annual data frequency. However, the assumptions behind the restrictions related to the interactions between the policy rate and the macroeconomic variables are not generally plausible in the case of this data frequency. Therefore, the paper interpolates the data on income inequality and uses quarterly data within the contemporaneous identification scheme. To use the actual annual data on income inequality, the paper also relies on long run identification.

In the case when the variables are not stationary and y_t is expressed through (2), the process still has a moving average representation (Johansen, 1995):

$$y_t = \Xi \sum_{i=1}^{t} u_i + \sum_{j=0}^{\infty} \Xi_j^* u_{t-j} + y_0^* \tag{8}$$

where y_0^* contains the initial values, Ξ_j^* are coefficient matrices that converge to zero as j gets large. That is, the matrices Ξ_j^* represent transitory effects of shocks and they are analogous to the matrices Φ_j from (5). The long run effects of shocks are captured by the matrix $\Xi = \beta_\perp\left[\alpha_\perp'\left(I_4 - \sum_{i=1}^{p-1}\Gamma_i\right)\beta_\perp\right]^{-1}\alpha_\perp'$. By replacing the reduced form residuals u_t in (8) by the linear combinations of structural shocks (6), the first two terms of the expression can be represented as follows: $\Xi B\sum_{i=1}^{t}\varepsilon_i$ and $B\sum_{j=0}^{\infty}\Xi_j^*\varepsilon_{t-j}$. That is, the long run and the transitory effects of structural shocks are given by ΞB and B, respectively. Therefore, restrictions for identification are placed on the long run impact matrix ΞB and the contemporaneous impact matrix B.

The matrix ΞB has rank *4-r* and, consequently, it could have at most *r* columns of zeros (transitory effects) and there might be at least *4-r* shocks with permanent effects (Lütkepohl, 2005). Talking into account the reduced rank of the matrix, each column of zeros counts as only *4-r* independent restrictions. Together all contemporaneous and long run restrictions for transitory and permanent shocks provide enough restrictions (6 in total) for just identification. Nevertheless, considering the issues regarding contemporaneous restrictions in the case of the annual data, the required restrictions are generally placed on the long run impact matrix.

As shown in the next section, there is only one cointegration relation among the variables. Therefore, there is only one shock with transitory effects. Following Duarte and Marques (2009), it is assumed that prices have transitory effects on the other variables. That is, the elements of the column of price shocks in the long run impact matrix are zeros. Taking into account that the matrix is singular, it only counts for 3 independent restrictions. In addition, it is also assumed that income inequality and real GDP do not have permanent effects on monetary policy rule. That is, it is assumed that, in the long run, the monetary policy shock is driven only by its own shock. For the final required restriction, it is assumed that inequality does not contemporaneously affect prices. Thus, the restrictions placed on the contemporaneous impact matrix and the long run impact matrix are the following[4]:

$$B = \begin{pmatrix} * & * & * & * \\ * & * & * & 0 \\ * & * & * & * \\ * & * & * & * \end{pmatrix} \quad and \quad \Xi B = \begin{pmatrix} * & 0 & * & * \\ * & 0 & * & * \\ 0 & 0 & * & 0 \\ * & 0 & * & * \end{pmatrix} \tag{9}$$

As an alternative set of restrictions, another identification scenario is also considered in the empirical analysis. In order not to restrict long run effects of monetary policy and its channels on income inequality, it is now assumed that inequality has temporary impact on the other variables. Again, it is assumed that, in the long run, the policy rule is solely driven by a monetary policy shock. In line with the previous identification restrictions, it is also assumed that prices do not

[4] Asterisks stand for unrestricted elements.

have permanent impact on real output. Thus, no restriction is imposed on the contemporaneous impact matrix. This is not necessary since there is only one shock with transitory effects (Lütkepohl, 2005). That is, only restrictions on the long run impact matrix are imposed:

$$B = \begin{pmatrix} * & * & * & * \\ * & * & * & * \\ * & * & * & * \\ * & * & * & * \end{pmatrix} \quad and \quad \Xi B = \begin{pmatrix} * & 0 & * & 0 \\ * & * & * & 0 \\ 0 & 0 & * & 0 \\ * & * & * & 0 \end{pmatrix} \tag{10}$$

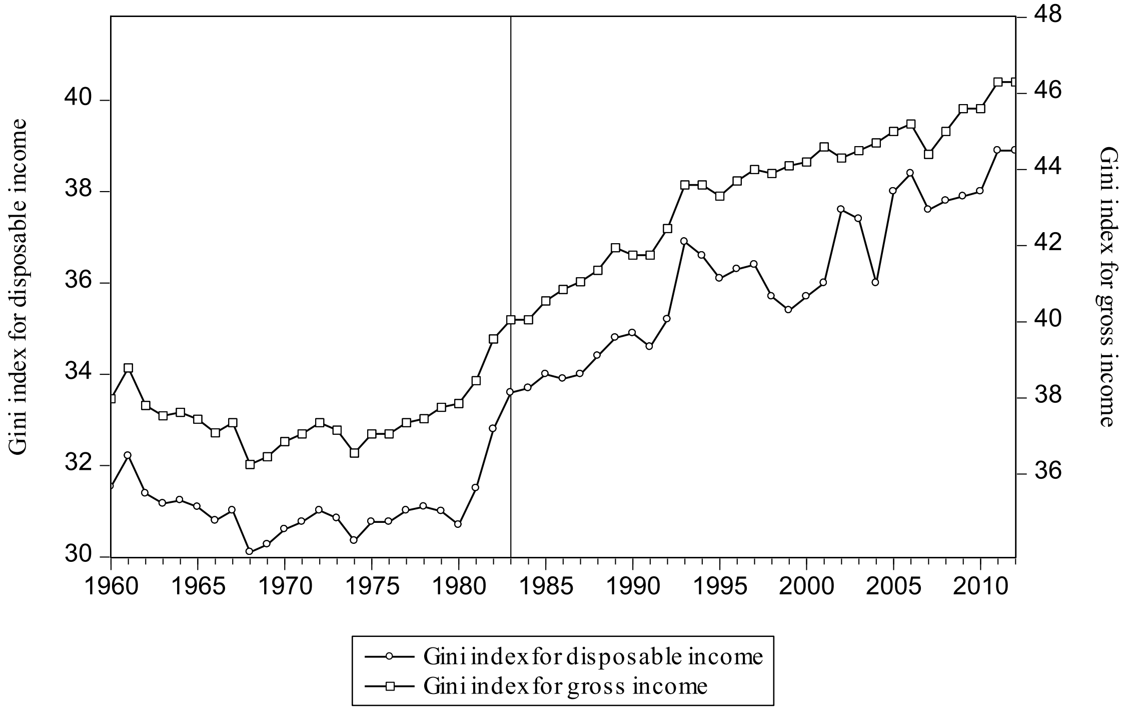
4. Data

The empirical analysis is implemented for the USA, where top income shares have the same dynamics as in other Anglo-Saxon countries (Atkinson and Leigh, 2013). It is a leading country whose income distribution affects the world income inequality (Atkinson, 2007). One of the major difficulties for the empirical analyses is the scarcity of the data on income inequality. Because of it, researchers often have to mix different classifications of data together. According to Knowles (2005), mixing various data classifications is not appropriate and it leads to results, which are not robust. Therefore, a lot of attention is paid in the paper to the usage of consistently measured comparable data on income inequality. The data source is the OECD. As an inequality measure, Gini index is used since it provides the broadest coverage across time. Nevertheless, the time series of Gini index is available solely on a yearly frequency. Gini index is expressed in percent and it is for disposable income. The usage of Gini index for disposable income (i.e., after taxes and transfers) allows controlling for the distributional effects of fiscal policy.

For Gini index, the longest consistently measured data series with the same income definition are available only for the period from 1979 to 2012. To present the dynamics of Gini index before 1979, additional data from Atkinson and Morelli (2014) are used for the period from 1960 to 1978. These data for Gini index are available only for gross income, as in other sources for that period. For the graphical representation solely, the paper combines Gini indices based on gross and net income, and from two different sources. Nevertheless, the paper beforehand adjusts the series from Atkinson and Morelli (2014) for the period from 1960 to 1978 towards the series from the OECD database to obtain comparable indices to some extent. The adjustment is implemented based on the averages of the overlapping values of the series. That is, keeping the same dynamics of the series from Atkinson and Morelli (2014), it is simply shifted towards the series from the OECD.

The prolonged series for Gini index based on disposable income is presented in Figure 1. The series for Gini index based on gross income is also provided in the same figure. Both Gini indices have an upward trend since around 1983. It is clearly observable a structural break in the series in around 1983. Further on, Gini index based on only disposable income is considered in the paper to control for the distributional effects of fiscal policy, as mentioned above.

Figure 1: Gini Indices for Gross and Disposable Income



Gini index of income inequality is measured for total population. In this respect, the paper compliments the work by Coibion et al. (2012) in evaluating the distributive effects of monetary policy by considering a measure of income inequality that includes the top one percent of income distribution. The coverage of the top share of income distribution is important for the overall measure of inequality (Atkinson, 2007; Atkinson et al., 2011; Leigh, 2007) and, consequently, for the evaluation of the distributive impact of monetary policy. To visualize that, in the USA, the dynamics of inequality is driven by the top one percent of income distribution (Congressional Budget Office, 2011; Kenworthy and Smeeding, 2013), the series for Gini index and for the income share of the top one percent[5] are graphed together in Figure 2. As can be seen, the both series actually have the same trend. That is, the dynamics of Gini index largely reflects the variation in the top one percent of income distribution. The scatter plot of these series illustrates that more apparently in Figure 3.

5 The series for the income share of the top one percent is from the database created by Piketty and Saez (2003). In particular, the updated version of the database (June, 2016) is used.

Figure 2: Gini index and the Income Share of the Top One Percent

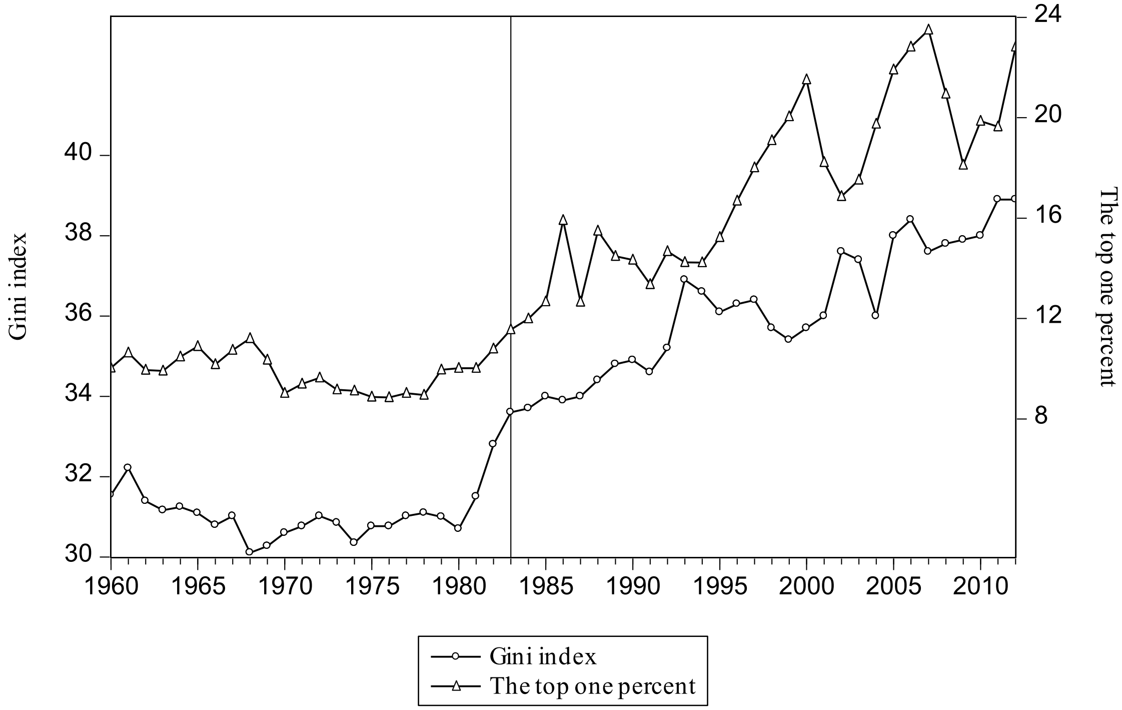
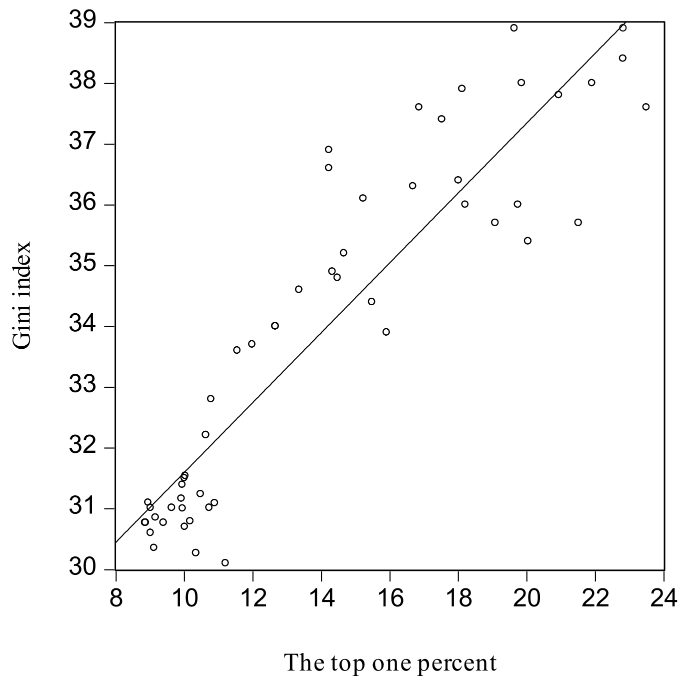
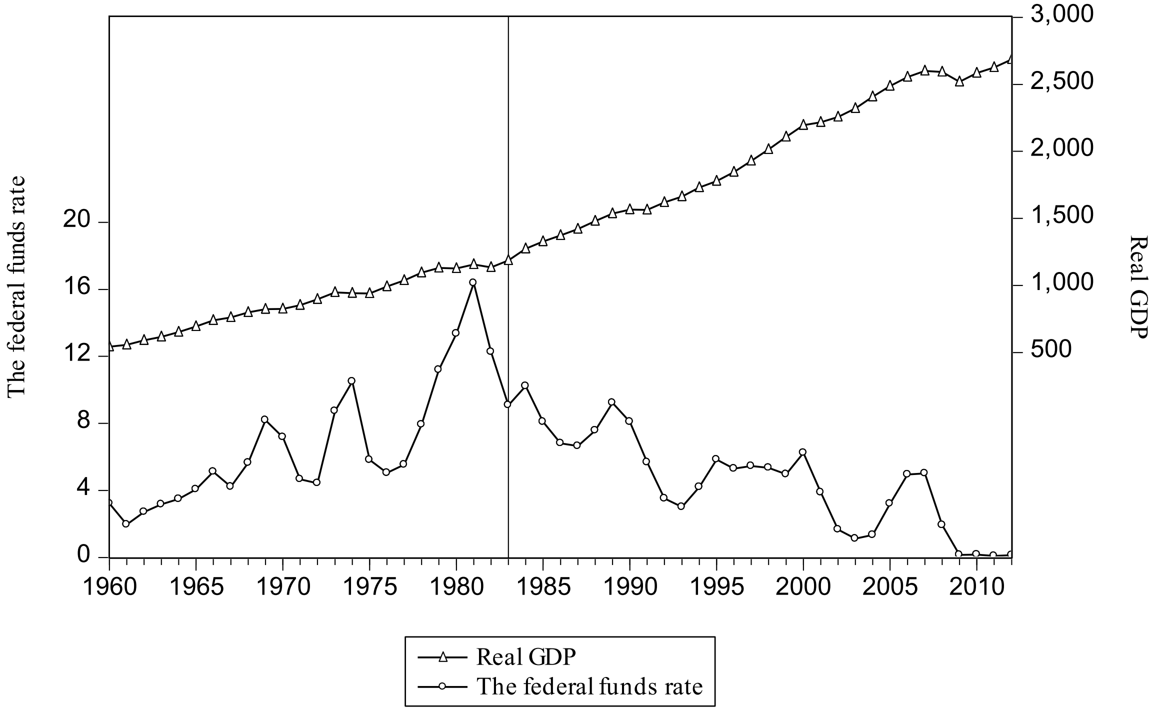


Figure 3: The Relation between Gini index and the Income Share of the Top One Percent



The other variables: real GDP, CPI, GDP deflator, and the federal funds rate, are taken from Federal Reserve Economic Database, FRED. In particular, the federal funds rate is extracted as an annual average (i.e., it is an effective rate) and it is expressed in percent. Real GDP is computed by using nominal GDP and a base index for GDP deflator. For the graphical representation and the initial analysis, the base year of the variables is changed to the initial date of 1960. The graphs of the variables are depicted in Figures 4 and 5. There was a visible structural break in around 1983 in almost all the time series except the series for real GDP.

Figure 4: Real GDP and the Federal Funds Rate



The paper also uses Chow breakpoint test (Lütkepohl, 2005) to check the existence of the structural break in 1983. To implement the test, the baseline VAR model is first estimated by ordinary least squares (OLS) based on the sample 1960-2012. Then, the breakpoint test is conducted for the year of 1983 in the cases of the VAR models of order one and two. For the both cases, CPI is also considered instead of GDP deflator. The results of the test are presented in Table 1. As can be observed, all the p-values are zero. Consequently, the test rejects the null hypothesis that there is no a breakpoint. Thus, the results of the test confirm that there was a structural break in 1983.

It is also stated in the literature (Cutler and Katz, 1991; Galli and von der Hoeven, 2001) that there was a structural break in the relationship between income inequality and macroeconomic variables in the USA in around 1983. Therefore, for actual estimations, the paper uses the sample values from 1983 to 2012. In addition, presample values (for the period 1981-1982, as it turns out during the analysis) are also used to preserve some degrees of freedom of the estimated models given the relatively short sample period. To examine the dynamics of the variables with respect to the beginning of the period, the base year for real GDP, CPI, and GDP deflator is shifted to 1983.

Figure 5: GDP Deflator and CPI

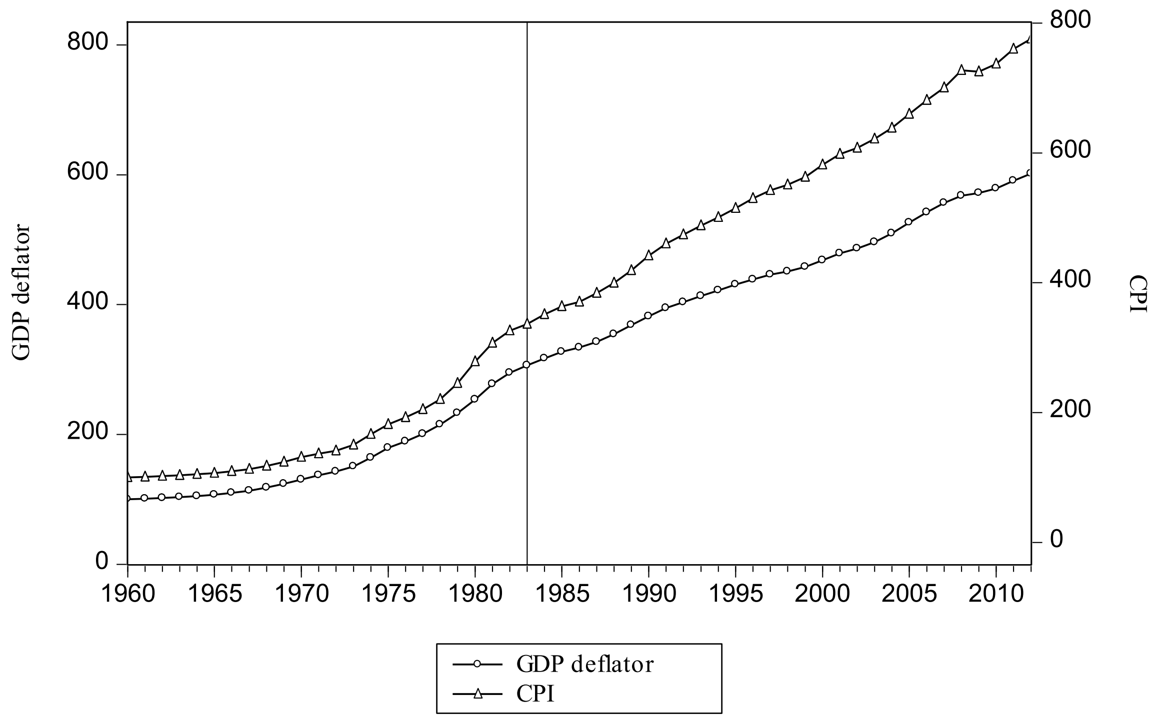


Table 1: Chow Breakpoint Test for the Year of 1983

Models	Test Values	p-Values	Bootstrap p-Values
VAR(1) with GDP deflator	109.50	0.00	0.00
VAR(2) with GDP deflator	154.93	0.00	0.00
VAR(1) with CPI	101.50	0.00	0.00
VAR(2) with CPI	155.80	0.00	0.00

Notes: The estimation sample is 1960-2012. Bootstrap p-values are based on 3000 replications.

Since during the period from 1983 to 2012, inflation in the USA was moderate, the relation between income inequality and inflation was probably linear. That is, that allows concentrating on the time dimension of the relationship between monetary policy and income inequality abstracting from the magnitude of the effect of inflation on inequality, which is claimed to be nonlinear along the levels of inflation by Galli and von der Hoeven (2001), and Bulir (2001). As a price index, GDP deflator is used in the empirical analysis because it measures the level of prices of all the goods and services produced in the economy. Nevertheless, the usage of CPI instead of GDP deflator would not make a significant difference since the both series are alike (Figure 5). To describe the general statistical characteristics of the variables used in the empirical analysis, they are provided for the estimation period from 1983 to 2012 in Table 2.

Table 2: The Descriptive Statistics of the Variables

Variables	Mean	Max.	Min.	SD
Real GDP (billions of USD, in prices of 1983)	6073.46	8223.14	3638.10	1482.16
Real GDP growth (annual percent change)	2.93	7.25	-2.77	1.86
GDP deflator (annual average index, 1983=100)	147.63	196.46	100	28.99
GDP deflator (annual percent change)	2.41	3.93	0.76	0.85
The federal funds rate (effective, annual average, in percent)	4.64	10.23	0.10	2.92
Gini index (in percent)	36.16	38.90	33.60	1.63

Notes: The statistical indicators are provided for the estimation period 1983-2012.

The time series of Gini index is available only on the yearly frequency in contrast to the others. Consequently, the series of the other variables are also considered on the annual basis in the case of the long run identification. This leads to a relatively short estimation sample, which raises concerns about the estimation results. Besides, in the case of the contemporaneous identification, the usage of the data on the yearly frequency would make the assumptions behind it too strong. That is, the application of the contemporaneous identification requires the usage of the data on a higher frequency. Therefore, to use the contemporaneous identification and a longer estimation sample, Gini index is interpolated from the annual frequency to the quarterly series[6]. The interpolation of Gini index is justifiable because its series has low variation.

[6] The interpolation is implemented by the specialized software ECOTRIM developed by Eurostat.

The disaggregation of Gini index is implemented by the index type. In particular, it is conducted in such a way that, for each reference period, the average of the interpolated series equals to the corresponding aggregate value. The series of Gini index is interpolated by the method proposed by Boot et al. (1967). The interpolation of the series by this method is curried out using the first difference approach. By implementing this interpolation procedure, the data for Gini index are disaggregated from the annual frequency to the quarterly series. The data for the other variables are available on the quarterly frequency. Thus, the interpolation of Gini index allows employing a longer time series and the contemporaneous identification in the paper using the quarterly data.

5. Empirical Analysis

5.1. Cointegration Analysis

As a standard approach, natural logarithmic transformation is implemented for the variables: real GDP (GDP83L), GDP deflator (GDPDX83L) except for Gini coefficient (GINI) and the federal funds rate (FFR)[7]. It is implemented to stabilize the variance of the time series, and, in the models, the coefficients of variables with logarithmic transformation are elasticities. Then, the visual inspection of the time series shows that they have apparent trends and, consequently, they cannot be stationary. The formal augmented Dickey-Fuller test (Dickey and Fuller, 1979) is implemented in the case of the annual data to check that and determine the order of integration of the series. The test is carried out as for the levels of the variables as well as for their first differences[8]. The results are provided in Tables 3 and 4. They indicate that all the time series are not stationary[9] and that the series are integrated of order one.

Taking into account that the time series are integrated of the same order, the paper applies Johansen methodology (Johansen, 1995) in order to check whether the series are cointegrated. To implement the cointegration test, the order of VECM is specified. Since the considered sample is relatively short, the specification approach is to determine the most parsimonious model possible. The order of VECM is selected based on the statistical analysis of the residuals. That is, the order is specified in such a way that VECM provides the adequate representation of the underlying data generation process. Based on this analysis, VECM(1) is specified.

[7] In the parentheses, the notations of the variables are mentioned as they are used in the empirical analysis. The number mentioned in the abbreviation is the last two digits of the base year while the letter "L" indicates the performed natural logarithmic transformation.
[8] Similar results are obtained by applying Phillips – Perron test (Phillips and Perron, 1988).
[9] Even if one or couple of the variables were initially stationary, the cointegration relation among the all variables could still hold within the more general definition of cointegration specified by Lütkepohl (2005).

Table 3: The Augmented Dickey-Fuller Test for the Levels of the Variables

				Critical Values			
Variables	Det. Terms	Lags	Test Values	1%	5%	10%	p-Values
GDP83L	c, t	1	-1.66	-4.30	-3.57	-3.22	0.75
GDPDX83L	c	2	-1.98	-3.68	-2.97	-2.62	0.29
FFR	c	2	-1.47	-3.68	-2.97	-2.62	0.53
GINI	c	2	-1.06	-3.68	-2.97	-2.62	0.72

Notes: Deterministic terms (c-constant and t-trend) are chosen according to the dynamics of the series. The order of the lagged differences is selected based on Schwarz information criterion.

Table 4: The Augmented Dickey-Fuller Test for the First Differences of the Variables

				Critical Values			
Variables	Det. Terms	Lags	Test Values	1%	5%	10%	p-Values
GDP83L	c	0	-4.19	-3.67	-2.96	-2.62	0.00
GDPDX83L	none	0	-2.47	-2.64	-1.95	-1.61	0.01
FFR	none	1	-4.91	-2.65	-1.95	-1.61	0.00
GINI	none	1	-5.89	-2.65	-1.95	-1.61	0.00

Notes: The inclusion of the deterministic term (c-constant) is associated with the dynamics of the series. The order of the lagged differences is selected based on Schwarz information criterion.

In order to carry out the cointegration test, the paper also specifies the deterministic terms to be included in the model. Since the series have trending behavior, all the most common cases of the deterministic terms are considered. Taking into account that in comparison to the maximum eigenvalue test, the trace test sometimes has more distorted sizes in small samples (Lütkepohl, 2005), the former is implemented as a cointegration test (Johansen, 1995). The results are presented in Table 5. All the results of the cointegration tests with different deterministic terms indicate that the time series are cointegrated, and there is one cointegrating relation among them.

Table 5: Johansen Cointegration Maximum Eigenvalue Test

Hypothesized No. of CEs	Det. Terms	Lags	Eigen-values	Test Values	5% Critical Values	p-Values
None*	c in CE	1	0.73	38.97	28.59	0.00
At most 1			0.49	20.19	22.30	0.10
At most 2			0.35	12.78	15.89	0.15
At most 3			0.14	4.38	9.16	0.36
None*	c in CE and in VAR	1	0.72	38.41	27.58	0.00
At most 1			0.43	16.94	21.13	0.17
At most 2			0.30	10.80	14.26	0.16
At most 3			0.11	3.55	3.84	0.06
None*	c, t in CE and c in VAR	1	0.73	39.31	32.12	0.01
At most 1			0.49	19.98	25.82	0.24
At most 2			0.35	12.83	19.39	0.34
At most 3			0.11	3.57	12.52	0.80

Note: The asterisk "*" denotes the rejection of the null hypothesis at the 5% significance level. The following abbreviations are also used: CE-cointegrating equation, c-constant, t-linear trend.

Based on the statistical features of the time series, a constant is considered in the models as a deterministic term. It is included in (1) and in the cointegration equation of (2). In particular, in the case of the usage of the quarterly data in the empirical analysis, the VAR models in levels are considered by estimating them with OLS. For modeling the relations among the variables in the case of the annual data, the VECM methodology is employed by applying Johansen´s maximum likelihood approach (Johansen, 1995).

As an empirical tool to explore the dynamic interactions among the variables, impulse response functions of the considered models are examined. The paper provides impulse response functions (IRFs) of the variables to one standard deviation increase in a monetary policy shock. Taking into account that, as a monetary policy indicator, the federal funds rate is considered, the monetary policy shock is contractionary. Hall´s (1992) 95% confidence bands based on 3000 bootstrap replications are provided for the IRFs. For the representation of the IRFs, solid lines are used while, for the demonstration of the confidence bands, dotted lines are drawn.

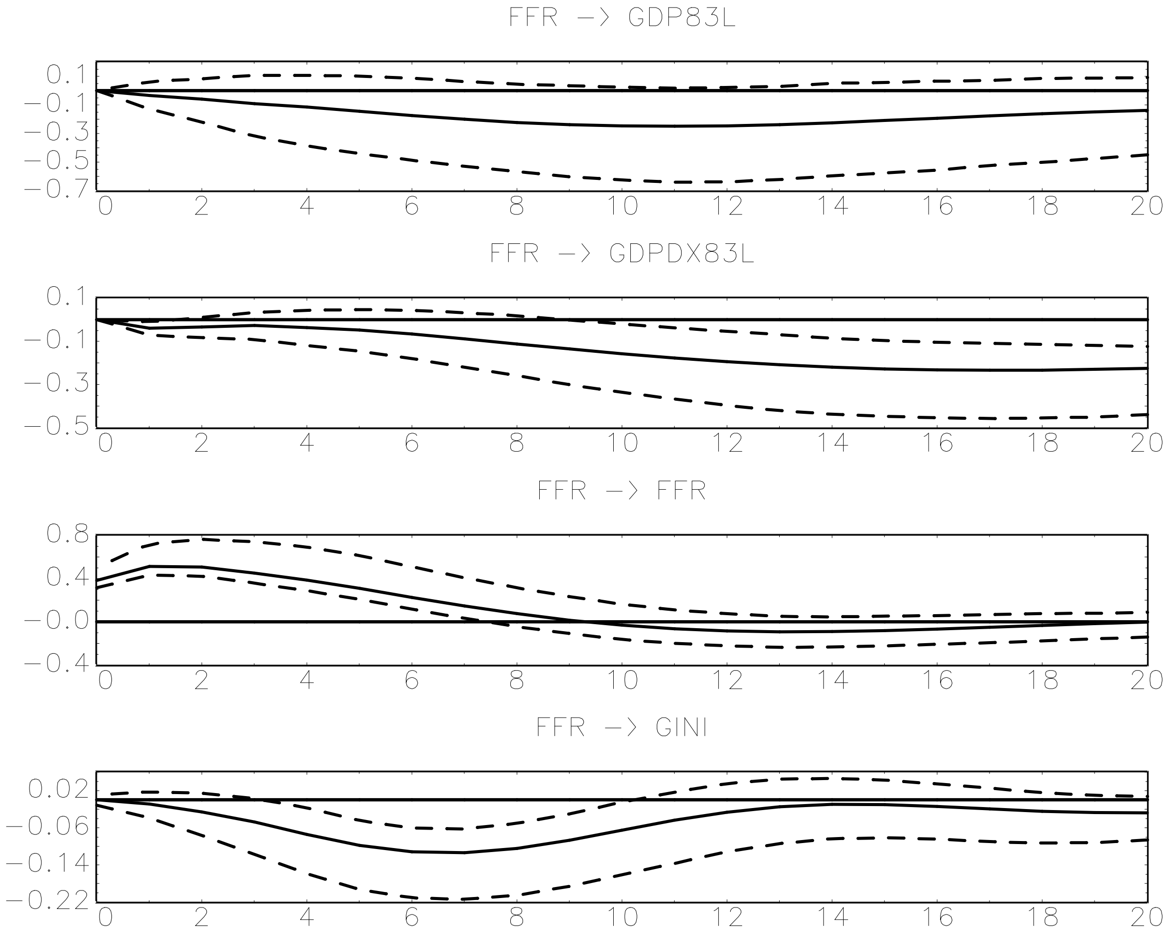
5.2. Contemporaneous Identification

The standard identification approach for a monetary policy shock is Cholesky decomposition of the covariance matrix of the VAR residuals. So, the empirical analysis is initially carried out using this identification scheme. In order to identify a monetary policy shock by this approach, the paper imposes the contemporaneous restrictions discussed in Section 3. However, as already mentioned, the contemporaneous assumptions are too strong in the case of the annual data. Therefore, by implementing the interpolation for Gini index, the data on the quarterly frequency are used for this identification scheme. In addition, the usage of the quarterly data allows addressing the concerns of the relatively short time series of the variables in the case of the annual data. By employing the longer series for the variables, the robustness of the analysis is checked with respect to the number of observations. In line with the related literature (Christiano et al., 1996; Coibion et al., 2012), the quarterly VAR model is estimated with four lags.

The estimated IRFs are provided in Figure 6. As can be seen, a monetary policy shock decreases real output up to 0.3 percent. This real impact of monetary policy is in accordance with the related literature (Christiano et al., 1996; Peersman and Smets, 2001). The monetary policy shock also reduces prices by around 0.2 percent. That is, the response of prices does not feature the issue of an increase in prices in response to tight monetary policy, known as a “price puzzle” in the literature (Bernanke and Blinder, 1992; Sims, 1992). Following the monetary policy shock, income inequality decreases up to 0.1 percentage points. This distributional effect of monetary policy is significant between the fourth and the tenth quarters after the shock.

As the results indicate, tight monetary policy decreases income inequality similar to the IRFs provided by Villarreal (2014) for Mexico and, on the contrary to the results obtained by Coibion et al. (2012) for the USA. Here the employed methodological approach is similar to the one applied by Coibion et al. (2012). Therefore, the differences in the obtained results lie in the data. In particular, they are related to the representativeness of the income inequality measures used in the empirical analyses. The current inequality measure represents the whole income distribution, including the top one percent, which has substantially influenced the dynamics of income inequality in the USA over the considered period (Atkinson et al. 2011; Congressional Budget Office, 2011; Kenworthy and Smeeding, 2013).

Figure 6: The IRFs to a Monetary Policy Shock
(The Baseline Case)



To check the robustness of the results with respect to the recent years of the estimation sample, the period when the federal funds rate reaches the zero lower bound is excluded from the sample. In particular, the IRFs are estimated for the sample period until 2008. The resulting IRFs are provided in Figure A1. As can be observed, the obtained results are similar to the IRFs provided in the previous case. The responses of real output, prices, and income inequality to a monetary policy shock have actually identical magnitudes. The significance levels of the responses of prices and income inequality are also analogous. The response of real output is just more significant in this case. Thus, the results are not affected by the exclusion or the consideration of the recent period after the financial crisis.

As alternative measures of economic inequality, Gini index of wage inequality and the ratio between the 90th percentile and the 10th percentile (thereafter, it is referred as the 90-10 ratio) are also considered. They are based on the data from the Current Population Survey (CPS) of the U.S. Census Bureau. It is a household survey which represents the resident civilian noninstitutionalized population of the USA, including also the upper part of income distribution. Though, this inequality measure is based on income before taxes and it does not include noncash benefits, it could still be helpful in assessing the distributive effect of monetary policy.

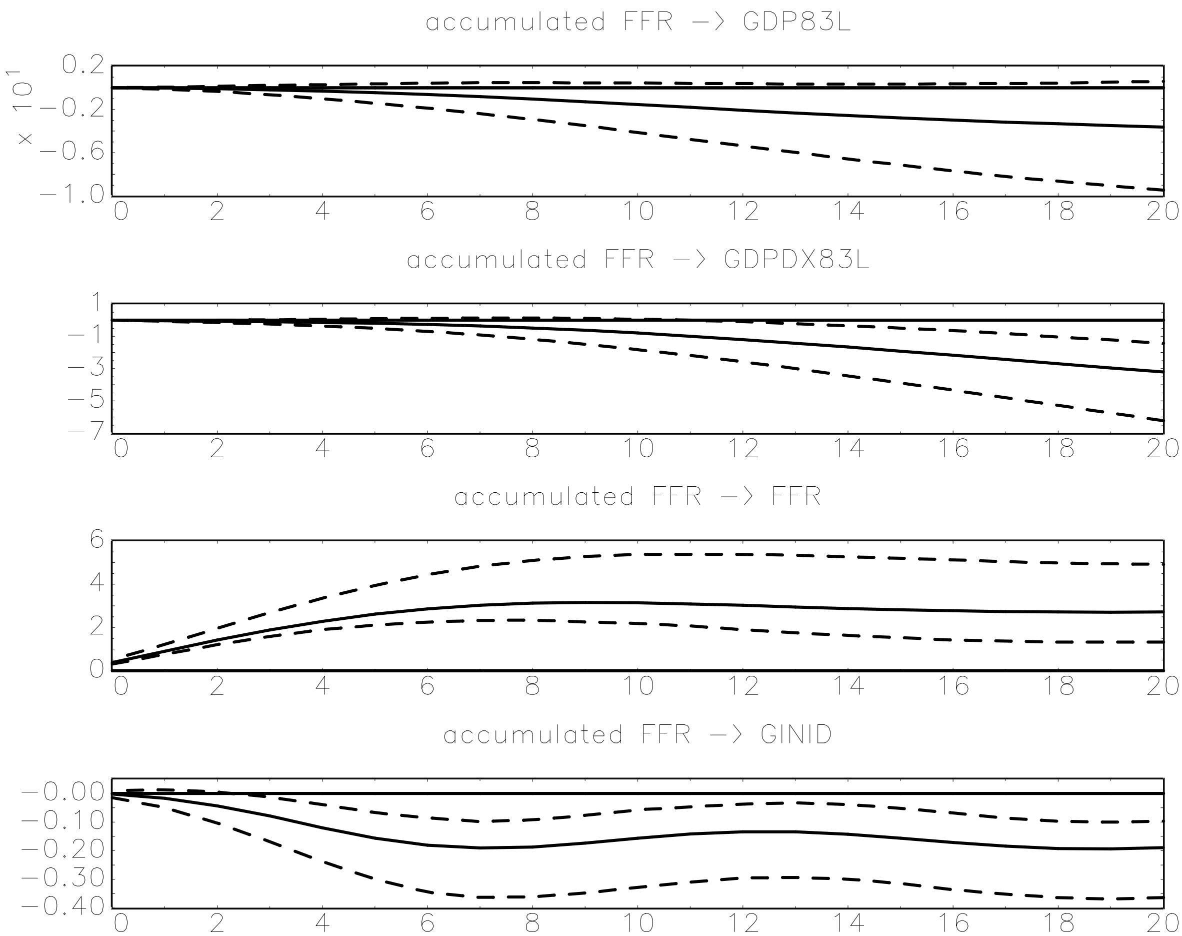
Gini index of wage inequality (GINIW) is computed by using the data provided in the Center for Economic and Policy Research (2016). In the VAR model, Gini index of income inequality is replaced by Gini index of wage inequality, and, analogously, the contemporaneous identification scheme is applied. The resulting IRFs are presented in Figure A2. As can be seen, the response of real output to a monetary policy shock is not significant in this case. However, the shock still leads to the significant reduction in prices. Besides, the shock decreases wage inequality by around 0.04 percentage points and the impact is significant since the third period until it vanishes. Following the shock, the reduction of wage inequality is smaller than the response of income inequality. This result is due to the fact that wages are just part of total income. That is, after a monetary policy shock, the stronger reduction of total income inequality is also related to the decrease in the inequality of business and financial income.

As another measure of income inequality, the 90-10 ratio (P9010) is also considered. It is from the CPS report by DeNavas-Walt and Proctor (2015) and available only on the yearly frequency. Therefore, analogously to Gini index of income inequality, the 90-10 ratio is interpolated to the quarterly frequency. Then, it is included in the VAR model instead of Gini index. The IRFs identified with the contemporaneous restrictions are provided in Figure A3. After a monetary policy shock, the responses of real output and prices are analogous to the results provided in the baseline case. The shock reduces the 90-10 ratio up to 0.03 units by the sixth quarter. The timing of this response displays the similar pattern with the corresponding result of the baseline case. Thus, the results are robust with regard to the usage of different inequality measures.

Before continuing the empirical analysis with the application of the VECM methodology, the existence of the long run distributive effect of monetary policy is examined within the framework of the contemporaneous identification. Following Born et al. (2015), this examination is implemented by considering the VAR model with Gini index of income inequality in the first differences (GINID) and with the other variables in levels. Then, the total effect of monetary policy on income inequality is checked for the significance. This evaluation of the long run effect is line with the method proposed by Blanchard and Quah (1989).

The IRFs are accumulated and they are depicted in Figure 7. As can be seen, the accumulated responses of real output and prices to a monetary policy shock are generally analogous to the corresponding results of the previous estimations of the IRFs. After the shock, the accumulated changes in Gini index decrease up to 0.2 percentage points. Besides, this total distributive effect of monetary policy is completely significant. That is, monetary policy has a long run effect on income inequality, and it is thoroughly examined in the next subsection.

Figure 7: The IRFs to a Monetary Policy Shock
(Gini Index in the First Differences)



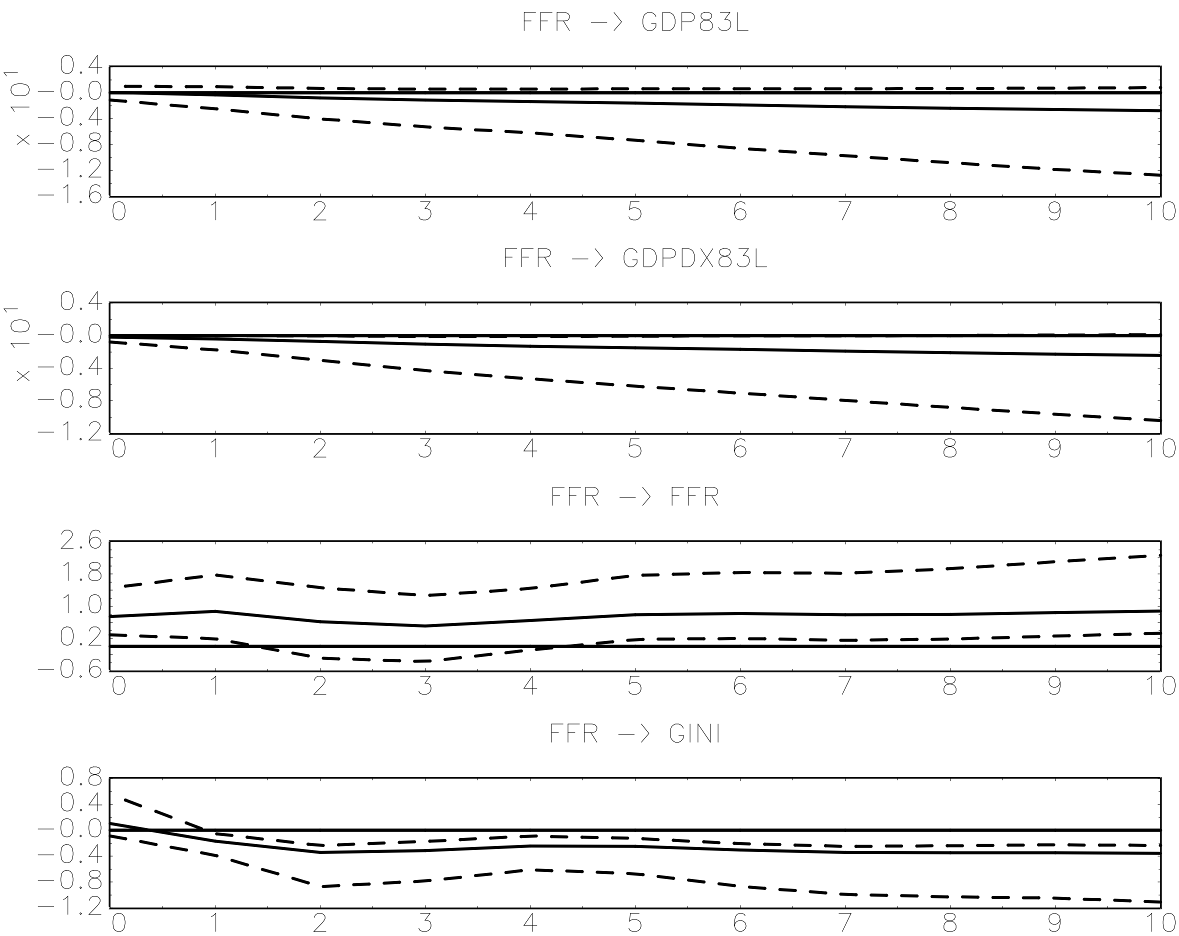
5.3. Long Run Identification

As revealed in Section 5.1, there is a cointegration relation among real GDP, prices, the federal funds rate, and Gini index of income inequality. Therefore, the distributive impact of monetary policy is also evaluated through the long run identification with the VECM methodology when the annual data are used. As discussed in Section 5.1, the VECM of order one is specified with a constant included into the cointegration equation. The IRFs are identified by imposing restrictions on the contemporaneous impact matrix and the long run impact matrix as described in (9) of Section 3.

The estimated IRFs are presented in Figure 8. As can be observed, a monetary policy shock significantly decreases real output and prices. Their responses are analogous to the previous results. The shock also significantly reduces income inequality by around 0.1 percentage points after the first period. Then, contractionary monetary policy decreases inequality by nearly 0.4

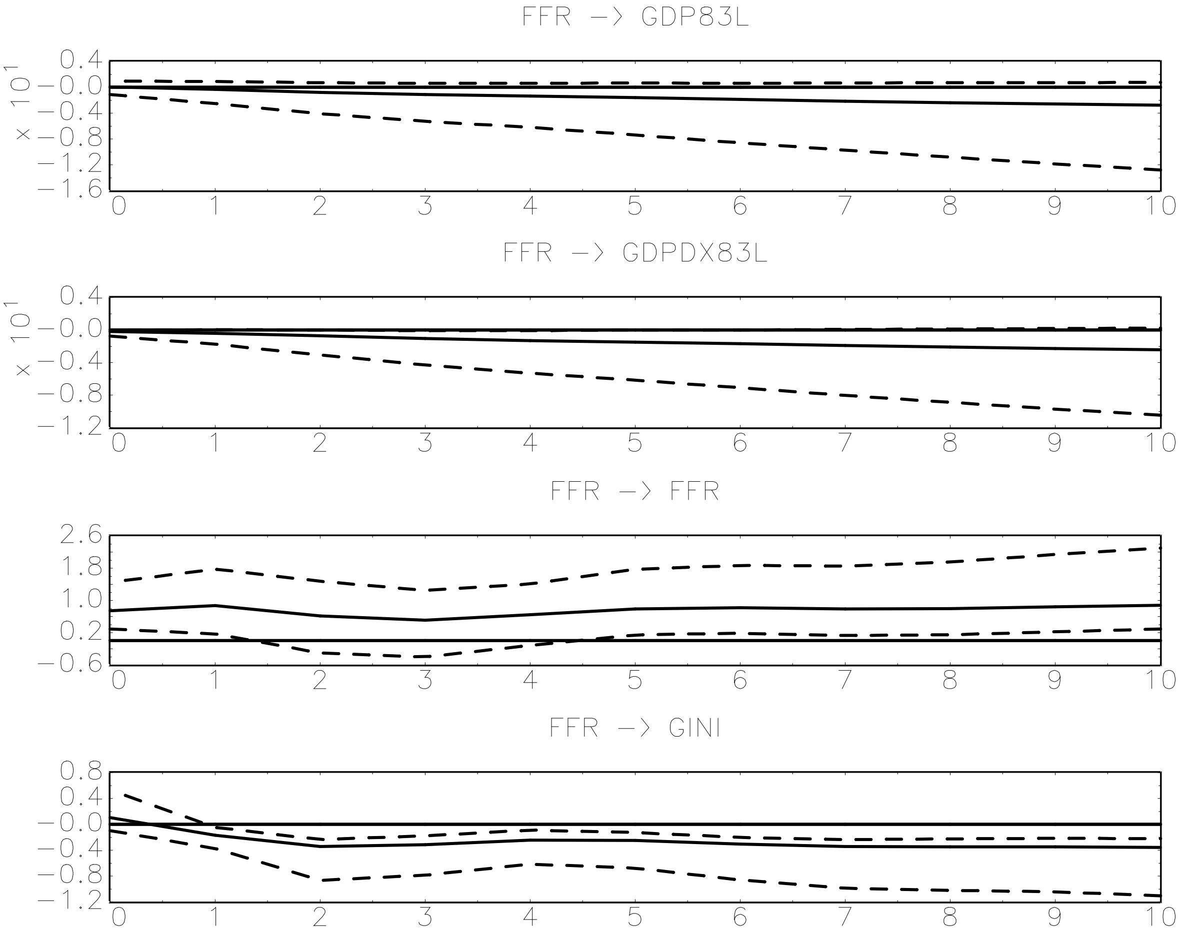
percentage points. Thus, the obtained IRFs are in line with the previous results provided in the case of the application of the other identification method with the quarterly data.

Figure 8: The IRFs to a Monetary Policy Shock
Identified Using the Restrictions in (9)



An alternative set of restrictions is also imposed within the VECM identification framework. As discribed in (10) of Section 3, no contemporaneous and long run restrictions are imposed on the impact of monetary policy and its channels on income inequality. The resulting IRFs are depicted in Figure 9. Comparing them with the results presented in Figure 8, it can be observed that the responses of the variables to a monetary policy shock are actually identical in the both cases. In particular, a monetary policy shock decreases Gini index of income inequality up to 0.4 percentage points.

Figure 9: The IRFs to a Monetary Policy Shock
Identified Using the Restrictions in (10)



Thus, in all the cases of the obtained results, income inequality decreases following a tightening of monetary policy. The distributive effects of monetary policy are in line with each other across different identification methods and frequency data. That is, the difference from the results obtained by Coibion et al. (2012) is in the representativeness of income inequality data, especially in the examination of the top one percent. This feature is evident from the results when the impact of monetary policy on income inequality is stronger than on wage inequality. These results are related not only to the composition of income but also to the tendency that the upper part of the income distribution includes the larger share of business and financial income.

6. Conclusion

The paper evaluates the distributional effect of monetary policy for the USA. A monetary policy shock is identified using different methods in the cases of quarterly and annual data. The inequality measures employed in the paper represents the whole income distribution, including the top one percent. The study period covers the time span after the structural break in the relationship between income inequality and the macroeconomics variables that occurred in around 1983. For the period after the structural break, a comprehensive cointegration analysis is carried out. The analysis determines a cointegration relation among real output, prices, the federal funds rate, and Gini index of income inequality. Therefore, the time series are modeled through the VECM along with the VAR representation.

A monetary policy shock is initially identified using the contemporaneous restrictions. Since the data on income inequality are available on the annual frequency, they are interpolated to the quarterly series. This allows applying the contemporaneous identification scheme and addressing the concerns regarding the relatively short time series of the variables when the annual data are used in the analysis. The obtained results indicate that a contractionary monetary policy shock leads to the significant decline in Gini index of income inequality by approximately 0.1 percentage points. As alternative economic inequality measures, Gini index of wage inequality and the 90-10 percentile ratio are also considered. Analogously, a contractionary monetary policy shock significantly reduces these measures of economic inequality.

Taking advantage of the existence of the cointegration relation among the variables, the identification of a monetary policy shock is also implemented through the VECM framework with the annual data. The application of two alternative sets of identifying restrictions gives actually identical results. A contractionary monetary policy shock decreases Gini index of income inequality up to 0.4 percentage points. That is, the results are in line with each other across all the considered identification approaches. Thus, contractionary monetary policy could reduce the overall income inequality in the country, and monetary policy might be considered as another effective policy instrument to decrease inequality.

References

Acemoglu, D. and Robinson, J. (2008), "Persistence of Power, Elites and Institutions," American Economic Review, 98 (1), pp. 267-293.

Afonso, A., Schuknecht, L., and Tanzi, V. (2010), "Income Distribution and Public Spending: an Efficiency Assessment," Journal of Economic Inequality, 8, pp. 367–389.

Albanesi, S. (2007), "Inflation and inequality," Journal of Monetary Economics, Elsevier, 54 (4), pp. 1088–1114.

Atkinson, A. (2007), "Measuring Top Incomes: Methodological Issues," In Top Incomes over the Twentieth Century: A Contrast between Continental European and English-Speaking Countries, Edited by A. Atkinson and T. Piketty, Oxford University Press, pp. 18–42.

Atkinson, A. and Leigh, A. (2013), "The Distribution of Top Incomes in Five Anglo-Saxon Countries Over the Long Run," The Economic Record, 89, pp. 31-47.

Atkinson, A. and Morelli, S. (2014), "Chartbook of Economic Inequality," ECINEQ, Society for the Study of Economic Inequality, Working Paper, 324.

Atkinson, A., Piketty, T., and Saez, E. (2011), "Top Incomes in the Long Run of History," Journal of Economic Literature, 49 (1), pp. 3–71.

Bank of England (2012), "The Distributional Effects of Asset Purchases," Quarterly Bulletin, 3, pp. 254-266.

Benabou, R. (2000), "Unequal Societies: Income Distribution and the Social Contract," American Economic Review, 90, pp. 96-129.

Bernanke, B. and Blinder, A. (1992), "The Federal Funds Rate and the Channels of Monetary Transmission," American Economic Review, 82, pp. 901–21.

Bernanke, B. and Mihov, I. (1998), "Measuring Monetary Policy," The Quarterly Journal of Economics, 113 (3), pp. 869-902.

Blanchard, O. and Quah, D. (1989), "The dynamic effects of aggregate demand and supply disturbances," American Economic Review, 79 (4), pp. 655–673.

Boot, J., Feibes, W., Lisman, J. (1967), "Further methods of derivation of quarterly figures from annual data," Applied Statistics, 16 (1), pp. 65–75.

Born, B., Müller, G. and Pfeifer, J. (2015), "Does Austerity Pay Off?" CEPR Discussion Paper, 10425, February.

Bulir, A. (2001), “Income Inequality: Does Inflation Matter?” IMF Staff Papers, 48 (1), pp. 139-159.

Center for Economic and Policy Research, (2016), “CPS ORG Uniform Extracts,” Version 2.1, Washington, DC.

Christiano, L., Eichenbaum, M. and Evans, C., (1996), "The Effects of Monetary Policy Shocks: Evidence from the Flow of Funds," The Review of Economics and Statistics, 78 (1), pp. 16-34.

Coibion, O., Gorodnichenko, Y., Kueng, L. and Silvia, J. (2012), “Innocent Bystanders? Monetary Policy and Inequality in the U.S,” NBER Working Paper 18170, National Bureau of Economic Research, Inc.

Congressional Budget Office (2011), “Trends in the Distribution of Household Income between 1979 and 2007,” Congress of the United States.

Cutler, D. and Katz, L. (1991), “Macroeconomic performance and the disadvantaged,” Brookings Papers on Economic Activity, 2.

DeNavas-Walt, C. and Proctor, B. (2015), “Income and Poverty in the United States: 2014,” Current Population Reports, U. S. Census Bureau, September.

Dickey, D. and Fuller, W. (1979), “Estimators for autoregressive time series with a unit root,” Journal of the American Statistical Association, 74, pp. 427–431.

Doepke, M. and Schneider, M. (2006), “Inflation and the Redistribution of Nominal Wealth,” Journal of Political Economy, University of Chicago Press, 114(6), pp. 1069–1097.

Doepke, M., Schneider, M. and Selezneva, V. (2015), “Distributional Effects of Monetary Policy,” Hutchins Center on Fiscal & Monetary policy at BROOKINGS, Working Paper, 14.

Doerrenberg, P. and Peichl, A. (2014), "The impact of redistributive policies on inequality in OECD countries," Applied Economics, 46 (17), pp 2066-2086.

Duarte, R. and Marques, C. (2009), "The dynamic effects of shocks to wages and prices in the United States and the euro area," Working Paper Series 1067, European Central Bank.

Easterly, W. and Fischer, S. (2001) “Inflation and the poor,” Journal of Money, Credit and Banking, Part 1, pp. 159–178.

Erosa, A. and Ventura, G. (2000), “On inflation as a regressive consumption tax,” Manuscript, University of Western Ontario.

Equality Trust (2013), “Income Inequality: Trends and Measures,” The Equality Trust Research Digest, http://bit.ly/etrust-2013.

Galbraith, J., Giovannoni, O. and Russo, A. (2007), "The Fed´s Real Reaction Function: Monetary Policy, Inflation, Unemployment, Inequality – and Presidential Politics," UTIP Working Paper, 42.

Galli, R. and von der Hoeven, R. (2001), "Is Inflation Bad for Income Inequality: the Importance of the Initial Rate of Inflation," ILO Employment Paper, 29.

Hall, P. (1992), "The Bootstrap and Edgeworth Expansion," Springer Series in Statistics, Springer-Verlag, New York.

Johansen, S. (1995), "Likelihood-based Inference in Cointegrated Vector Autoregressive Models," Oxford University Press, Oxford.

Kang, S., Chung, Y. and Sohn, S. (2013), "The effects of Monetary Policy on Individual Welfares," Korea and the World Economy, 14 (1), pp. 1-29.

Kenworthy, L. and Smeeding, T. (2013), "Growing Inequalities and their Impacts in the United States," GINI Country Report United States.

Knowles, S. (2005), "Inequality and Economic Growth: The Empirical relationship Reconsidered in the Light of Comparable Data," Journal of Development Studies, 41(1), pp. 135-159.

Lee, J. (2014), "Monetary policy with heterogeneous households and imperfect risk-sharing," Review of Economic Dynamics, 17 (3), pp.505-522.

Leigh, A. (2007), "How Closely Do Top Income Shares Track Other Measures of Inequality?" Economic Journal, 117 (524), pp. F619-F633.

Li, H. and Zou, H. (2002), "Inflation, Growth, and Income Distribution: A Cross-Country Study," Annals of Economics and Finance, 3, pp. 85-101.

Lütkepohl, H. (2005), "New Introduction to Multiple Time Series Analysis," Springer, Berlin.

Muinelo-Gallo, L. and Roca-Sagales, O. (2011), "Economic Growth, Inequality and Fiscal Policies: A Survey of the Macroeconomics Literature," In: Bertrand, R.L. (Ed.), Theories and Effects of Economic Growth, Chapter 4. Nova Science Publishers, Inc., pp. 99–119.

Mumtaz, H. and Theophilopoulou, A. (2016), "The Impact of Monetary Policy on Inequality in the UK. An Empirical Analysis," Working Papers 783, Queen Mary University of London, School of Economics and Finance.

Nakajima, M. (2015), "The Redistributive Consequences of Monetary Policy," Business Review, Federal Reserve Bank of Philadelphia Research Department, Second Quarter, pp. 9-16.

Neves, P. and Silva, S. (2014), “Inequality and Growth: Uncovering the Main Conclusions from the Empirics,” Journal of Development Studies, 50 (1), pp. 1-21.

O'Farrell, R., Rawdanowicz, L. and Inaba, K. (2016), “Monetary Policy and Inequality” OECD, Economics Department Working Papers, 1281.

Peersman, G. and Smets, F. (2001), "The monetary transmission mechanism in the euro area: more evidence from VAR analysis," Working Paper Series 0091, European Central Bank.

Phillips, P. and Perron, P. (1988), “Testing for a unit root in time series regression,” Biometrika 75, pp. 335–346.

Piketty, T. and Saez, E. (2003), “Income Inequality in the United States, 1913-1998,” The Quarterly Journal of Economics, 118 (1), pp. 1-39.

Romer, C. and Romer, D. (1999), “Monetary Policy and the Well-Being of the Poor,” Economic Review, Federal Reserve Bank of Kansas City, First Quarter, pp. 21-49.

Saiki, A. and Frost, J. (2014), “How Does Unconventional Monetary Policy Affect Inequality? Evidence from Japan,” DNB Working Paper 423.

Sims, C. (1992), “Interpreting the Macroeconomic Time Series Facts: the Effects of Monetary Policy,” European Economic Review, 36, pp. 975–1011.

Uhlig, H. (2005), "What are the effects of monetary policy on output? Results from an agnostic identification procedure," Journal of Monetary Economics, 52 (2), pp 381-419.

Villarreal, F. (2014), “Monetary Policy and Inequality in Mexico,” MPRA Paper 57074.

Wolff, E and Zacharias, A. (2007), "The Distributional Consequences of Government Spending and Taxation in the U.S., 1989 and 2000," Review of Income and Wealth, 53 (4), pp. 692-715.

Appendix

Figure A1: The IRFs to a Monetary Policy Shock
(The Reduced Sample)

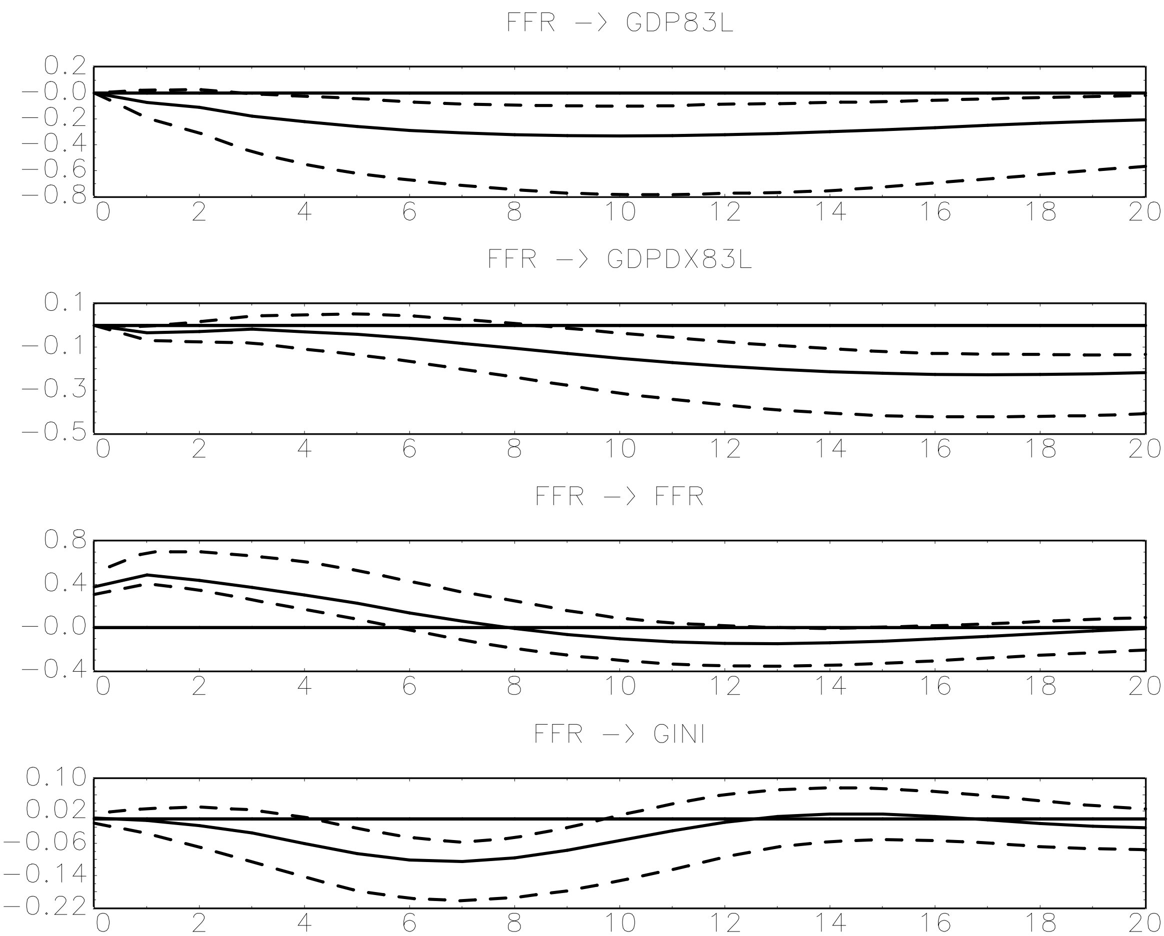


Figure A2: The IRFs to a Monetary Policy Shock
(Gini Index of Wage Inequality as an Economic Inequality Indicator)

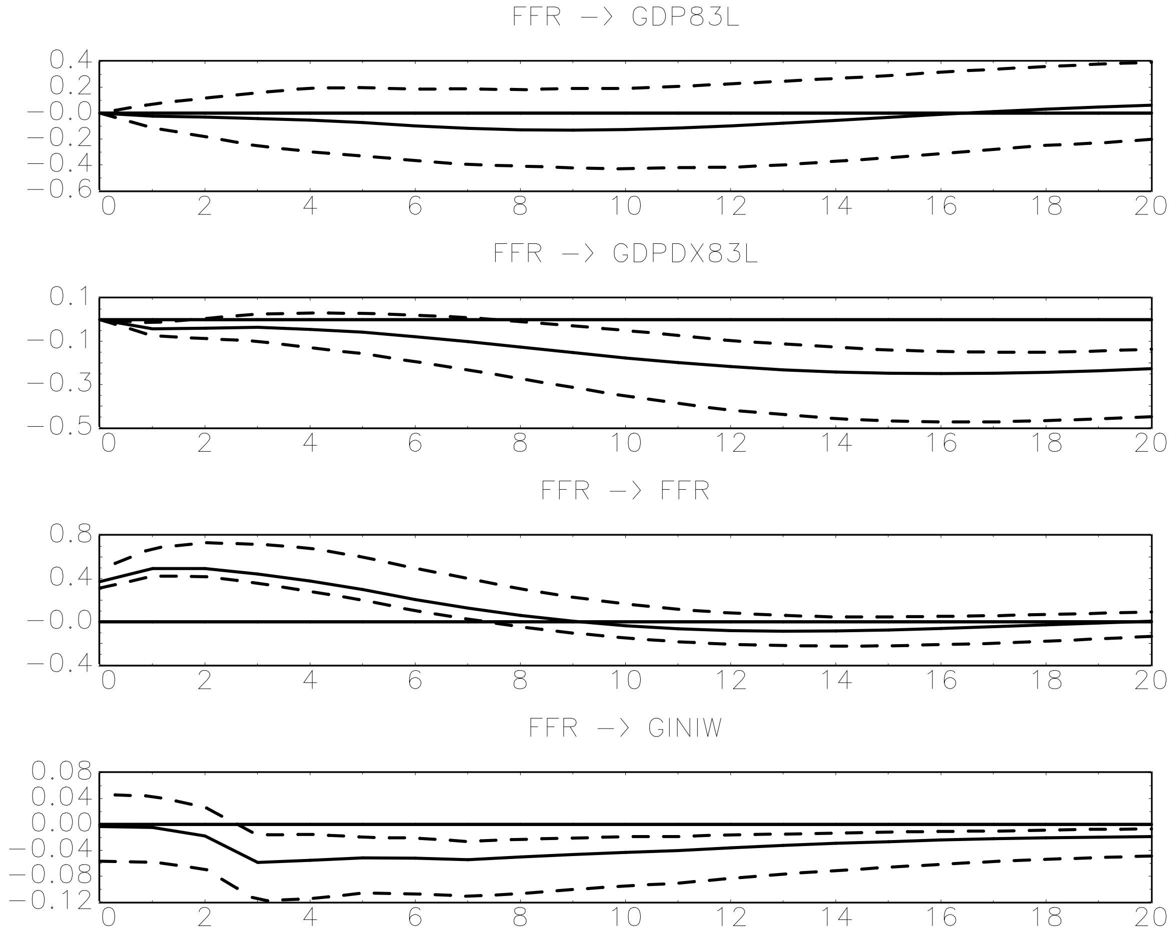


Figure A3: The IRFs to a Monetary Policy Shock
(The 90-10 Ratio as an Income Inequality Measure)