

“Income Inequality and Monetary Policy: An Analysis on the Long Run Relation”

Karen Davtyan



Institut de Recerca en Economia Aplicada Regional i Públic
Research Institute of Applied Economics

WEBSITE: www.ub-irea.com • CONTACT: irea@ub.edu



Grup de Recerca Anàlisi Quantitativa Regional
Regional Quantitative Analysis Research Group

WEBSITE: www.ub.edu/aqr/ • CONTACT: aqr@ub.edu

Universitat de Barcelona

Av. Diagonal, 690 • 08034 Barcelona

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Abstract

The distributional effect of monetary policy is estimated in the case of the USA. In order to identify a monetary policy shock, the paper employs contemporaneous restrictions with ex-ante identified monetary policy shocks as well as log run identification. In particular, a cointegration relation has been determined among the considered variables and the vector error correction methodology has been applied for the identification of the monetary policy shock. The obtained results indicate that contractionary monetary policy decreases income inequality in the country. These results could have important implications for the design of policies to reduce income inequality by giving more weight to monetary policy.

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Karen Davtyan: AQR Research Group-IREA. University of Barcelona, Av. Diagonal 690, 08034 Barcelona, Spain. E-mail: karen.davtyan@ub.edu

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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 redistributive effects have not extensively been discussed. The objective of the paper is to contribute to this discussion by evaluating the effect of monetary policy on income inequality.

Redistributive 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). In the political economy arguments, the redistribution of income is implied to be implemented through fiscal policy by taxation and government spending. However, income is redistributed 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 redistribution, and are in constant interaction with each other.

High inflation can create uncertainty, raise expectations of future macroeconomic instability, disrupt financial markets, and lead to distortionary economic policies (Romer and Romer, 1998). 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 inequality in earnings and total income in the USA.

The estimated effect of monetary policy could depend on the inequality measure used in the empirical analysis. That is, the estimated effects might differ if the inequality measure is from another data source and it does not represent the whole income share of population, particularly the top one percent. In the USA,

the dynamics of income inequality has mainly been driven by the variation in the upper end of distribution since early 1980's (Congressional Budget Office, 2011). The paper evaluates the distributional effect of monetary policy in the USA by using the inequality measure that covers the whole income distribution, including the top one percent.

The paper finds a cointegration relation among real output, prices, the federal funds rate, and Gini index of income inequality. Consequently, vector error correction and equivalent vector autoregression models are used for the analysis of the relationship. In order to identify a monetary policy shock, the paper employs contemporaneous identification with ex-ante identified monetary policy shocks and log run identification. In particular, the vector error correction methodology is applied for the identification of the monetary policy shock. The obtained results show that contractionary monetary policy reduces the overall income inequality in the country.

The rest of the paper is organized as follows. Section 2 reviews the related academic literature while 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

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. However, there are some insightful papers discussing different aspects of distributive effects of monetary policy and they are discussed thoroughly below. In addition, these distributive effects, which are evaluated in the considered literature, are summarized in Table 1.

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 Høeften (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.

Galbraith et al. (2007) show that in the USA, 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.

Coibion et al. (2012) find 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) finds 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.

Nakajima (2015) claims that while monetary policy affects prices and real economic activity, it also has redistributive impact. In order to control for these main effects of monetary policy, the paper includes prices and real GDP into the considered models. As a monetary policy tool, the federal funds rate is used. 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 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. In particular, the paper compliments the work by Coibion et al. (2012) in evaluating the distributive effect of monetary policy by considering the measure of income inequality when it includes the top one percent of income distribution. The results show that the choice of the inequality measure has substantial impact on the evaluation of the distributive effect of monetary policy.

Table 1: The Estimated Effects of Contractionary Monetary Policy on Economic Inequality in the Literature

Cross-Country Evidence	Time Series Evidence for a Country
- (66 countries; Romer and Romer, 1999)	+ (USA; Romer and Romer, 1999)
- (75 countries; Bulir, 2001)	+ (USA; Galbraith et al., 2007; Coibion et al., 2012)
- (46 countries; Li and Zou, 2002)	- (Japan; Saiki and Frost, 2014)
- (51 countries; Albanesi, 2007)	- (Mexico; Villarreal, 2014)

3. Empirical Methodology

The examination of the distributional effects of monetary policy is implemented through multiple time series analysis. This analysis allows tackling the endogeneity problem among the variables and studying their interrelations. The considered vector autoregression of the order p , VAR(p), is the following¹:

$$y_t = A_1 y_{t-1} + \dots + A_p y_{t-p} + u_t, \quad (1)$$

where y_t is the vector of endogenous variables, A_i s are (4×4) coefficient matrices and $u_t = (u_{1t}, \dots, u_{4t})'$ is an error term. It is assumed that the error term is a zero-mean independent white noise process with positive definite covariance matrix $E(u_t u_t') = \Sigma_u$. That is, error terms are independent stochastic vectors with $u_t \sim (0, \Sigma_u)$. In the specification of the model, the vector of endogenous variables y_t consists of real GDP, prices, the federal funds rate, and income inequality measure: $y_t = (Y_t, P_t, R_t, Z_t)'$.

For the cointegrated variables, the equivalent vector error correction model of order $p-1$, VECM(p-1), should be used:

¹ The notations are in line with the representations used by Lütkepohl (2005).

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

where Δy_t denotes the first order differences of y_t , $\Gamma_i = -(A_{i+1} + \cdots + A_p)$ for $i = 1, \dots, p-1$, $\Pi = -(I_k - A_1 - \cdots - A_p)$. 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 Γ_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):

$$A_1 = \Gamma_1 + \Pi + I_k, A_i = \Gamma_i - \Gamma_{i-1} \text{ for } i = 2, \dots, p-1; A_p = -\Gamma_{p-1}. \quad (3)$$

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

$$y_t = \mu_t + x_t \quad (4)$$

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.

Reduced-form disturbances are linear combinations of structural shocks:

$$u_t = B\varepsilon_t \quad (5)$$

where ε_t is a (4×1) vector of structural innovations 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. There are different identification approaches that require out of sample information. The identification approaches used in the paper are presented below.

One of the most commonly employed identification approaches is Cholesky decomposition. It imposes the following contemporaneous restrictions on the matrix B :

$$\begin{pmatrix} u_Y \\ u_P \\ u_R \\ u_Z \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 \\ b_{21} & 1 & 0 & 0 \\ b_{31} & b_{32} & 1 & 0 \\ b_{41} & b_{42} & b_{43} & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_Y \\ \varepsilon_P \\ \varepsilon_R \\ \varepsilon_Z \end{pmatrix} \quad (6)$$

Analogously, long run restrictions (Blanchard-Quah, 1989) on the following total impact matrix are also low-triangular:

$$(I_4 - A_1 - \dots - A_p)^{-1} B \quad (7)$$

The zeros in these low-triangular matrices provide 6 required restrictions for just identification.

In the case of VECM, restrictions for identification are placed on the contemporaneous impact matrix and the long run impact matrix (Lütkepohl, 2005). There can be at most r shocks with zero long run impact (transitory effects) and at least $(4-r)$ shocks with permanent effects. Contemporaneous and long restrictions for transitory and permanent shocks provide enough restrictions for just identification.

As shown in the next section, there is only one cointegration relation among the variables. Therefore, there is only one shock with transitory effects (Lütkepohl, 2005). 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. For the final required restriction (6 in total), 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:

$$B = \begin{pmatrix} * & * & * & * \\ * & * & * & 0 \\ * & * & * & * \\ * & * & * & * \end{pmatrix} \quad and \quad \Xi B = \begin{pmatrix} * & 0 & * & * \\ * & 0 & * & * \\ 0 & 0 & * & 0 \\ * & 0 & * & * \end{pmatrix} \quad (8)$$

As a robustness check for these 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 monetary policy shocks. In line with the previous identification restrictions, it is also assumed that prices do not have permanent impact on real output. Thus, no restriction is imposed on the contemporaneous impact matrix. Since there is only one shock with transitory effects that is not necessary (Lütkepohl, 2005). That is, only restrictions on the long run impact matrix are imposed:

$$B = \begin{pmatrix} * & * & * & * \\ * & * & * & * \\ * & * & * & * \\ * & * & * & * \end{pmatrix} \text{ and } EB = \begin{pmatrix} * & 0 & * & 0 \\ * & * & * & 0 \\ 0 & 0 & * & 0 \\ * & * & * & 0 \end{pmatrix} \quad (9)$$

4. Data

The empirical analysis is implemented for the USA. One of the major difficulties for empirical analyses of the distributional effects of monetary policy is the scarcity of the data on income inequality. Therefore, a lot of attention is paid in the paper to the usage of consistently measured comparable data on income inequality. As an inequality measure, Gini coefficient is used since it provides the broadest coverage across time. The data source is the OECD. Gini coefficients are expressed in percent and they are for disposable income. The usage of Gini coefficients for disposable income (i.e., after taxes and transfers) allows controlling for the distributional effects of fiscal policy. The time series of Gini index is available only on the yearly frequency and, consequently, the series for the other variables are also considered on the annual basis.

Gini index for income inequality (GINI)² is measured for total population. In this respect, the paper compliments the work by Coibon et al. (2012) in

² In the parentheses, the abbreviated versions of the variables are mentioned in line with their usage in the empirical analysis.

evaluating the distributive effects of monetary policy by considering the measure of income inequality when it includes the top one percent of income distribution. The results show that this augmentation of inequality measure has substantial impact on the evaluation of distributive monetary policy effects.

The definitions and the sources of the other variables are as following. The real GDP (GDP60)³ is computed by using the data for nominal GDP and deflator from the World Bank, WB, and Federal Reserve Economic Database, FRED, respectively. For GDP deflator (GDPDX60) and CPI (CPIX60), base indices are used. The source for GDP deflator and CPI is FRED. The effective federal funds rate (FFR) is computed as an annual average. It is expressed in percent, and its source is FRED.

For the period from 1979 to 2012 (as it is available in the OECD database for the consistently measured index), the graphical representation of Gini coefficients is presented in Figure 1. Gini coefficients have an upward trend from around 1983. To present the dynamics of Gini coefficients before 1979, Gini coefficients from UNU-WIDER database are also employed from 1960 to 1978. To obtain a comparable series, Gini coefficients from UNU-WIDER database are adjusted towards the series from the OECD database. 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 UNU-WIDER, it is simply shifted towards the series from the OECD. The added values of the series of Gini coefficients are depicted in the same Figure 1. It is clearly observable a structural break in the series in around 1983.

The evolutions of the other variables are presented in Figures 2 to 5. There was a visible structural break in around 1983 in almost all the time series except of the series for real GDP. Literature (e.g., Cutler and Katz, 1991; Galli and von der Hoeven, 2001) also states that there was a structural break in the relationship between income inequality and macroeconomic variables in the USA in around 1983. For actual estimations, the paper uses the sample values for the period from 1983 to 2012. In addition, pre sample values (for the period 1981-1982, as it turns out during the analysis) are also used to preserve some degrees of

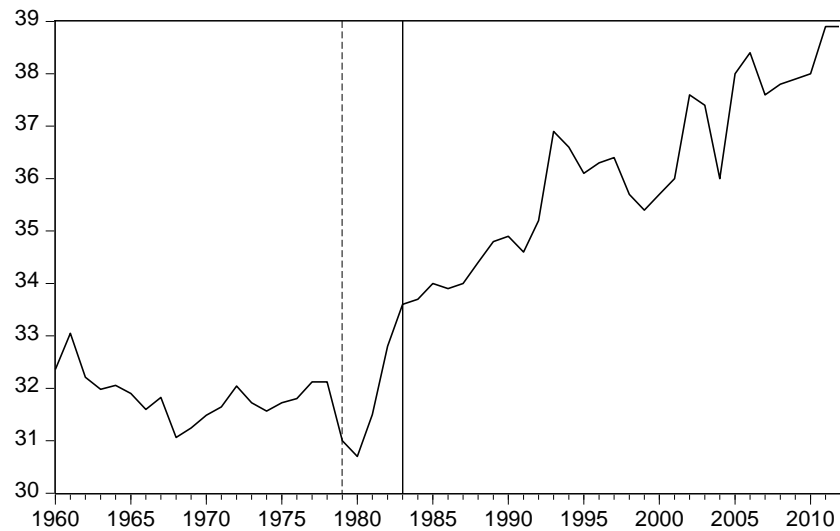
³ The number mentioned in the abbreviation is the last two digits of the base year.

freedom of the estimated models given the relatively short sample period. To observe the dynamics of the variables with respect to the beginning of the period, the base year for real GDP, CPI, and GDP deflator has been shifted to 1983.

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 (Figures 3 and 4). In order to describe the general statistical characteristics of the variables used in the empirical analysis, they are presented in Table 2.

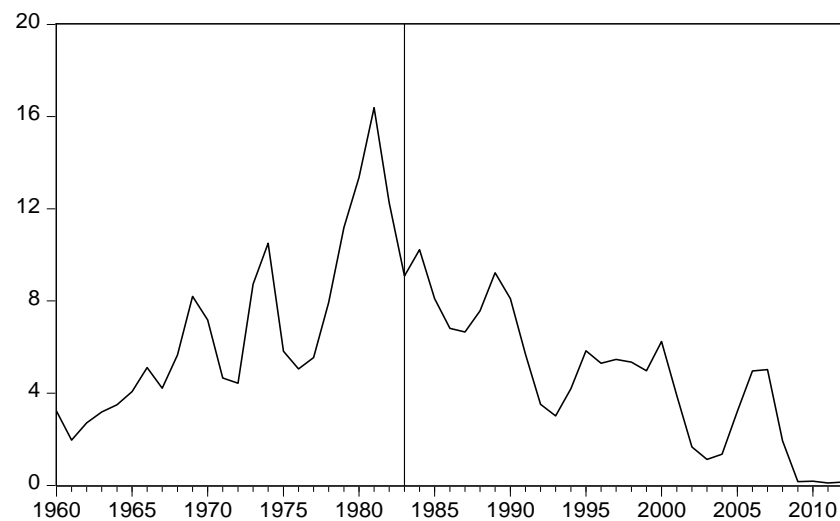
Thus, taking into account that the frequency of the data is yearly, the standard contemporaneous assumptions would be too strong. Therefore, the identification of a monetary policy shock is implemented by using the contemporaneous identification with ex-ante identified monetary policy shocks. In addition, a monetary policy shock is also identified by imposing long run restrictions. All these are discussed in detail in the next section.

Figure 1: Gini Coefficients (GINI)



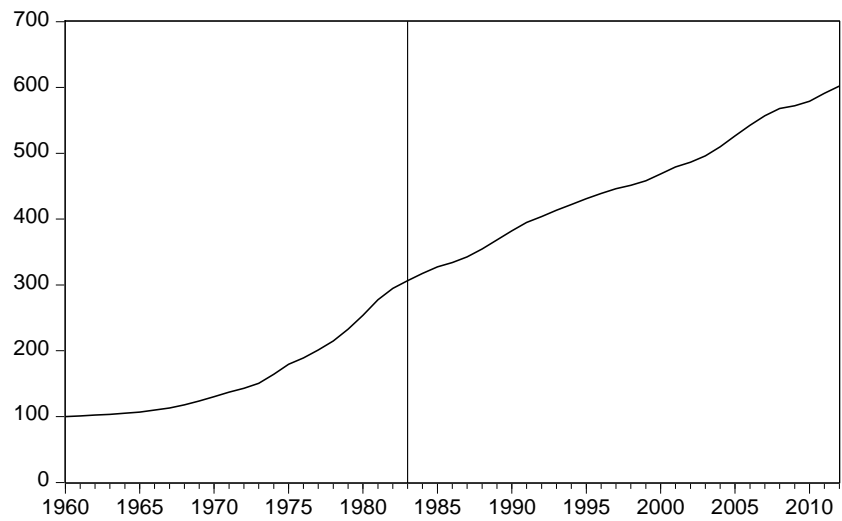
Note: The Gini coefficients are expressed in percent. They are for disposable income and total population. From 1960 to 1978, the data from UNU-WIDER are used and adjusted towards the series from the OECD for the period from 1979 to 2012.

Figure 2: The Effective Federal Funds Rate (FFR)



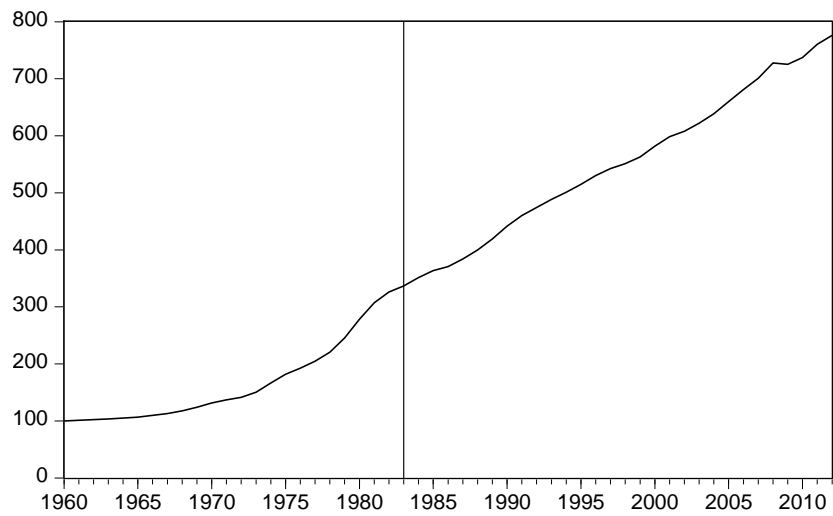
Note: The effective federal funds rate is computed as an annual average. It is expressed in percent, and its source is FRED.

Figure 3: GDP Deflator (GDPDX60)



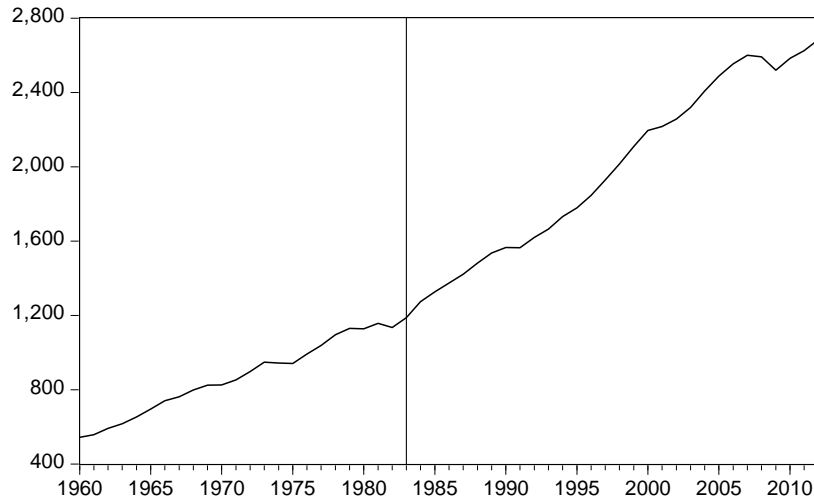
Note: The base year of the GDP deflator has been changed to 1960. The source for the initial data is FRED.

Figure 4: CPI (CPIX60)



Note: The base year of CPI has been shifted to 1960. The initial data are from FRED.

Figure 5: Real GDP (GDP60)



Note: The real GDP is in bln USD, and it is based on the prices of 1960. It is computed by using the data for nominal GDP and deflator from the WB and FRED, respectively.

Table 2: The Descriptive Statistics of the Variables, 1983-2012

Variables	Mean	Max.	Min.	SD
Real GDP (GDP83) (billions of USD, based on the prices of 1983)	6073.59	8227.14	3638.14	1482.36
Real GDP Growth (GRGDP) (annual percent change)	2.93	7.25	-2.77	1.86
GDP Deflator (GDPDX83) (annual average index, 1983=100)	147.63	196.46	100	28.99
GDP Deflator (GDPD) (annual percent change)	2.41	3.93	0.76	0.85
The Federal Funds Rate (FFR) (effective, annual average, in percent)	4.64	10.23	0.1	2.92
Gini Coefficient (GINI) (in percent)	36.16	38.9	33.6	1.63

5. Empirical Analysis

5.1. Cointegration Analysis

Natural logarithmic transformations are implemented for the variables: real GDP (GDP83L), GDP deflator (GDPDX83L) except for Gini coefficient (GINI) and the federal funds rate (FFR)⁴. 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 to check that and determine the orders of integration of the series. The test is carried out as for the levels of the variables as well as for their first differences⁵. The results are provided in Tables 3 and 4. The results of the augmented Dickey-Fuller test reveal that all the time series are not stationary⁶ and that they are integrated of order one.

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

Variables	Det. Terms	Lags	Test Values	Critical Values			P-Values
				1%	5%	10%	
GDP83L	c, t	1	-1.67	-4.30	-3.57	-3.22	0.74
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

Note: 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.

⁴ In the parentheses, the notations of the variables are mentioned as they are used in the empirical analysis. The letter L indicates the performed natural logarithmic transformation.

⁵ Similar results are obtained by applying Phillips – Perron test (Phillips and Perron, 1988).

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

Variables	Det. Terms	Lags	Test Values	Critical Values			P-Values
				1%	5%	10%	
GDP83L	c	0	-4.20	-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

Note: 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.

If the time series are cointegrated, they should be modeled through the error correction methodology or the corresponding VAR representation. Particularly, VECM will be employed if they are cointegrated because the paper aims to explore the dynamic interactions among the variables. Johansen methodology (Johansen, 1995) is carried out in order to check whether the series are cointegrated. To implement the cointegration test, the order of VECM or the corresponding VAR model should be determined since they are equivalent representations if there are no restrictions imposed on the cointegration relation. The order of VECM is one less than the order of VAR model.

Since the considered sample is relatively short, the specification approach is to determine the most parsimonious model possible. The order of VAR/VECM is selected based on the statistical analysis of the residuals. That is, the order is specified in such a way that VAR/VECM provides an adequate representation of the underlying data generation process. Tests for residual autocorrelation, non-normality, conditional heteroskedasticity, and stability are performed. Based on the results of these tests, VAR(2) (or, equivalently VECM(1)) is specified. For the cointegration test, it is also necessary to specify 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.

Table 5: Johansen Cointegration Maximum Eigenvalue Test

Hypothesized No. of CEs	Det. Terms	Lags	Eigen- values	Test Values	5% Critical Values	P-Values
None*			0.73	38.97	28.59	0.00
At most 1	c in CE	1	0.49	20.18	22.30	0.10
At most 2			0.35	12.75	15.89	0.15
At most 3			0.16	4.38	9.16	0.36
None*			0.72	38.41	27.58	0.00
At most 1	c in CE	1	0.43	16.94	21.13	0.17
At most 2	and in VAR		0.30	10.73	14.26	0.17
At most 3			0.11	3.57	3.84	0.06
None*			0.73	39.31	32.12	0.01
At most 1	c, t in CE	1	0.49	19.98	25.82	0.24
At most 2	and c in VAR		0.35	12.80	19.39	0.34
At most 3			0.11	3.59	12.52	0.80

Note: The following abbreviations are used: CE-cointegrating equation, c-constant, t-linear trend.

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. Based on the statistical features, a constant is considered in models as a deterministic term. It is included in the cointegration equation of

VECM or VAR, which are the benchmark models of the paper. For modeling the relations among the variables, VECM methodology is employed by applying Johansen's maximum likelihood (ML) approach (Johansen, 1995). Alternatively, the corresponding VAR model in levels is also used with ordinary least squares (OLS) estimations. As an empirical tool to explore the dynamic interactions among the variables, impulse response functions of the considered models are examined. In the paper, the provided impulse response functions (IRFs) are for the responses of variables to one standard deviation increase in the shock of the considered variable. In particular, the IRFs of contractionary monetary policy shocks are considered. 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.

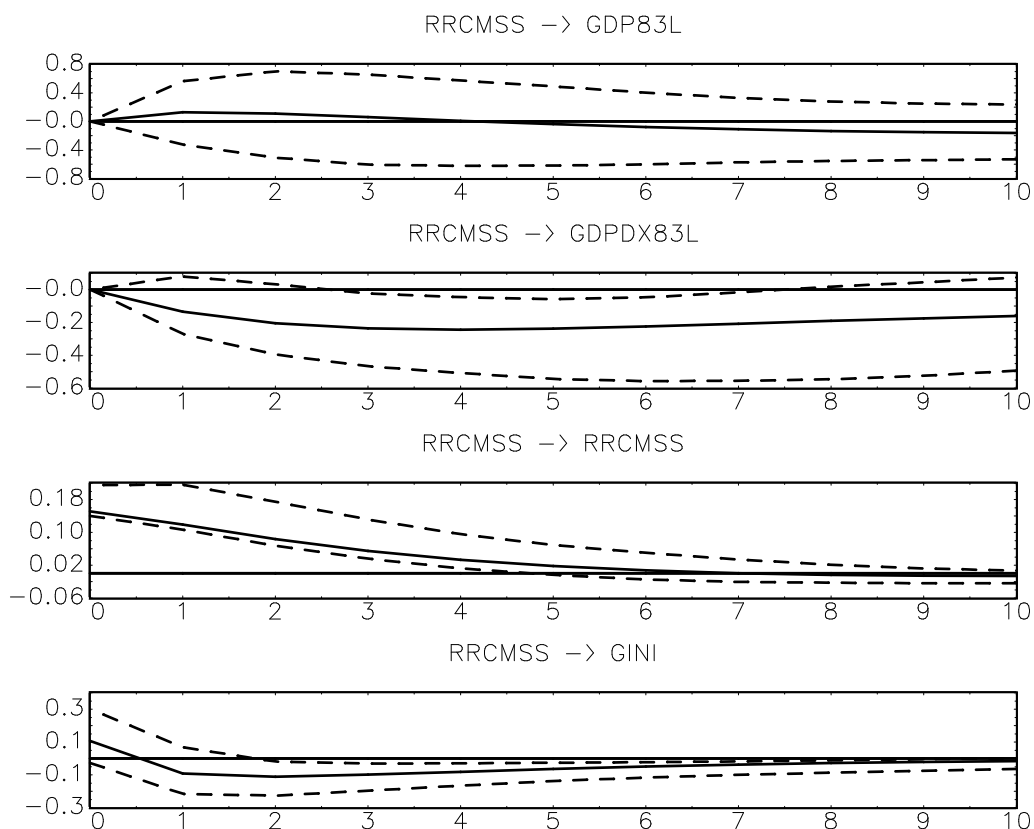
5.2. Contemporaneous Identification

The standard identification approach in the literature is Cholesky identification for VAR models. So, the empirical analysis is initially carried out using this identification procedure. It is necessary to impose contemporaneous restrictions discussed in Section 3 in order to implement that identification scheme. Taking into account that the yearly data are used in the analysis, the contemporaneous assumption that a monetary policy shock does not affect output and prices within a year is very strong in this case. Therefore, Cholesky identification is used with the exogenous monetary policy shocks proposed by Romer and Romer (2004). The series for these monetary policy shocks have been updated by Coibion et al. (2012) and they are used in the estimation of the IRFs. For the usage in the current analysis, they have been averaged across years. Then, following Coibion et al. (2012), they have been accumulated (RRCMSS) and placed instead of the federal funds rate in the VAR model estimated with a yearly lag.

The IRFs derived using the exogenous monetary policy shocks are provided in Figure 6. As can be seen, a contractionary monetary policy shock insignificantly increases income inequality on impact. However, the shock then gradually decreases inequality significantly up to around 0.1 percentage points in a period

and it generally stays at that level for several years until the effect fades away. The monetary policy shock also reduces prices while its effect on real output is not significant. Thus, contractionary monetary policy decreases income inequality similar to the estimated IRFs by Villarreal (2014) for Mexico and on the contrary to the results obtained by Coibion et al. (2012) for the USA. As it will be shown in the next paper, this distributive effect of monetary policy is preserved when the quarterly data are also used as in Coibion et al. (2012). Therefore, the differences in the obtained results lie in the data source and the measure of income inequality used in the empirical analysis. In the current work, the measure of inequality represents the whole distribution of income. Coibion et al. (2012) employ inequality measures that do not cover the top one percent of income distribution, which has substantially influenced the dynamics of income inequality in the USA over the considered period (Congressional Budget Office, 2011).

Figure 6: Contemp. Identification with Exogenous Monetary Policy Shocks

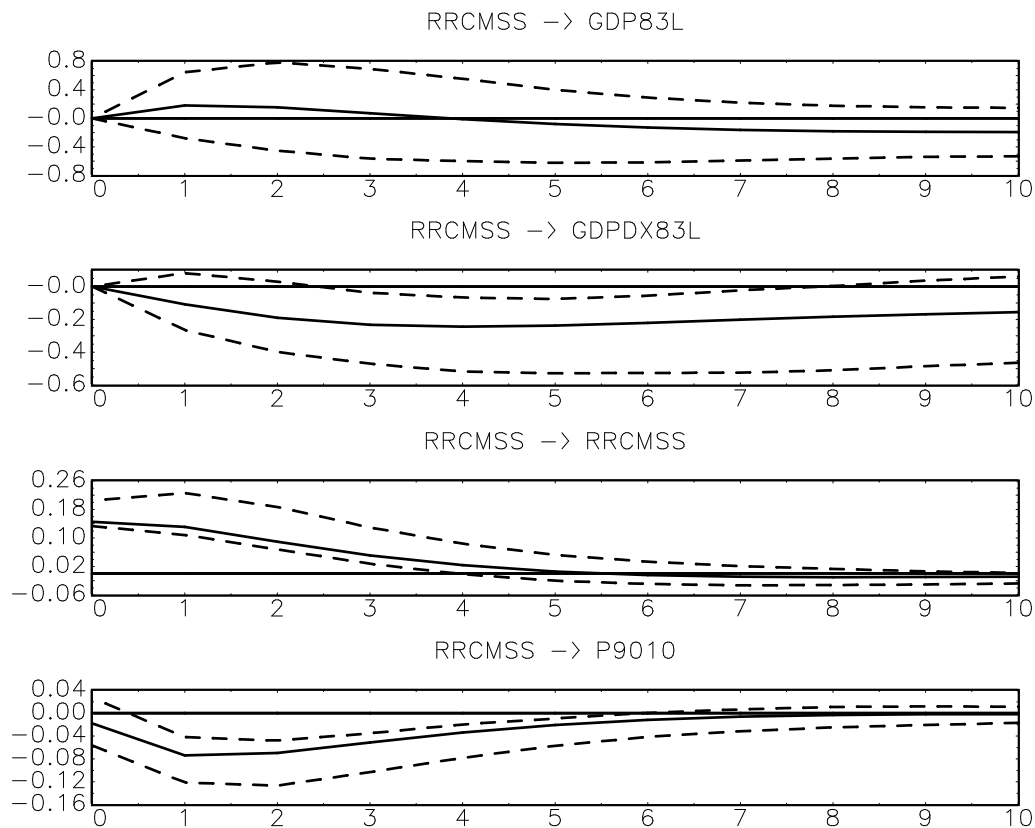


As another measure of income inequality, the ratio between the 90th percentile and the 10th percentile (thereafter, it is referred as the 90-10 ratio) is considered from the report by DeNavas-Walt and Proctor (2015). This percentile ratio is based on the data from the Current Population Survey (CPS) of the U.S. Census Bureau. It is a household survey which includes the resident civilian noninstitutionalized population of the USA. Besides, this inequality measure is based on income before taxes and it does not include noncash benefits (DeNavas-Walt and Proctor, 2015). However, the inequality measure could still be helpful in assessing the distributive effect of monetary policy and in performing a robustness check of the results.

In the considered VAR model for the contemporaneous identification, Gini index has been replaced by the 90-10 ratio (P9010). The resulting IRFs are presented in Figure 7. As can be observed from the obtained results, a contractionary monetary policy shock reduces inequality measured by the 90-10 ratio throughout the considered periods. The impact reaches its lowest point of the around 0.08 units decrease in the 90-10 ratio by the first period. The responses of the other variables to the monetary policy shock exhibit very similar behavior with the results provided in the previous case. Thus, the results are robust with regard to the usage of different inequality measures.

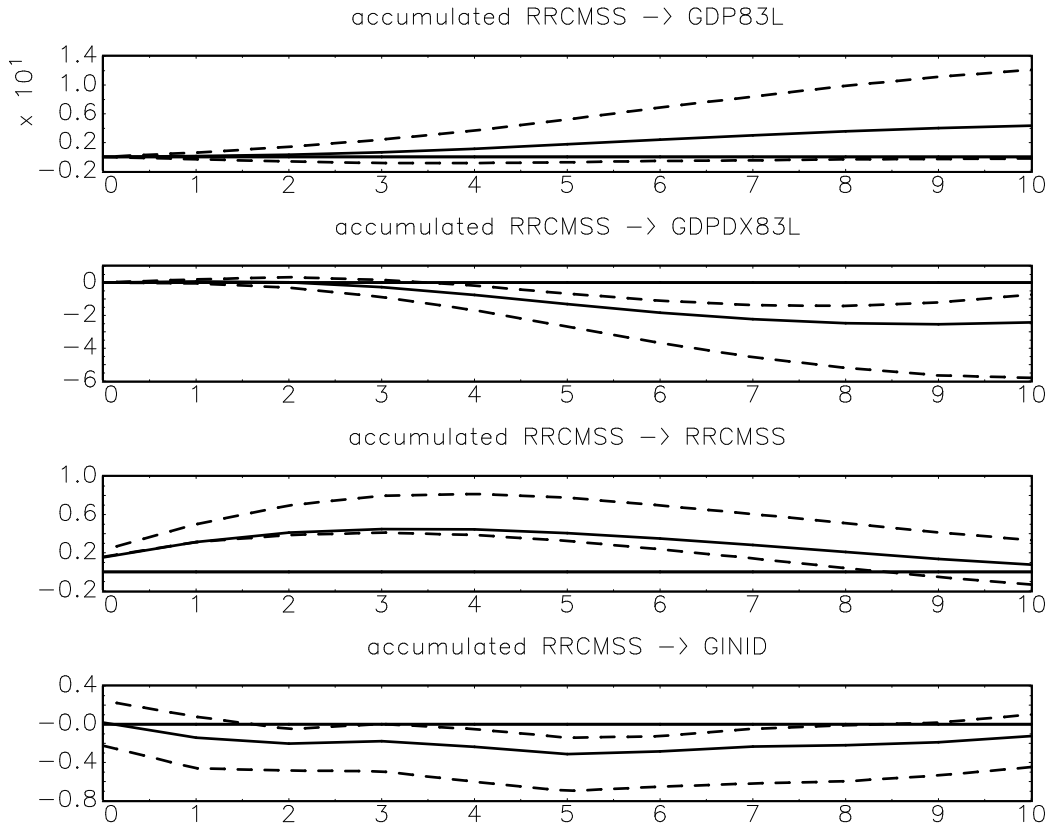
Before continuing the empirical analysis with the long run identification, the existence of the long run distributive effect of monetary policy is examined within the framework of the contemporaneous identification. This examination is implemented by following Born et al. (2015) and by considering the VAR model with Gini index in the first differences (calculated for the whole sample, GINID) and with the other variables in levels. Then, the total effect of monetary policy on inequality is checked for the significance based on the VAR of order two as specified in the previous subsection. This identification approach is line with the method proposed by Blanchard and Quah (1989).

Figure 7: Contemporaneous Identification with Exogenous Monetary Policy Shocks (Gini Index is Replaced by the 90-10 Ratio)



The IRFs are accumulated and they are depicted in Figure 8. It can be seen that after a contractionary monetary policy shock, the accumulated changes in Gini index decrease up to 0.2 percentage points. Besides, the total distributive effect of monetary policy is generally significant. That is, monetary policy has a long run effect on income inequality and it is thoroughly examined in the next subsection. The responses of the other variables are consistent with the corresponding results of the previous estimations of the IRFs.

Figure 8: Contemporaneous Identification with Exogenous Monetary Policy Shocks (Gini Index is in the First Differences)

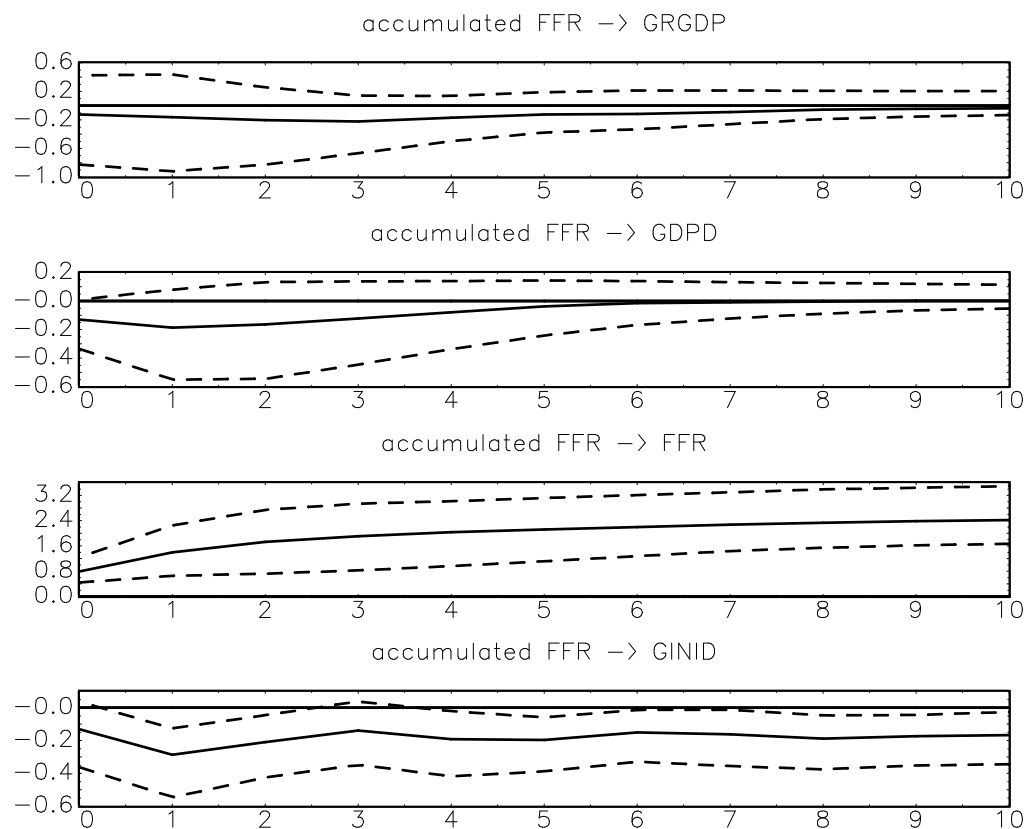


5.3. Long Run Identification

After revealing a long run relation between monetary policy and income inequality, the distributive effect of monetary policy is studied by the long run identification methods commonly used in the literature. First, the identification approach proposed by Blanchard and Quah (1989) is directly implemented in order to evaluate the distributive impact of contractionary monetary policy. Analogously with their approach, the VAR model is considered with real GDP growth (GRGDP), GDP deflator inflation (GDPD), the federal funds rate (FFR), and with the first order difference of Gini index (GINID). The VAR model is of the second order as in the benchmark case.

According to the identification method by Blanchard and Quah (1989), long run restrictions are imposed on the total impact matrix as discussed in Section 3. The accumulated IRFs are provided in Figure 9. As in the case with the usage of exogenous monetary policy shocks, the accumulated changes in Gini index decrease to around 0.2 percentage points after a contractionary monetary policy shock. The accumulated response of real GDP growth is insignificant as the response of real GDP in Figure 8. Though GDP deflator decreases following the contractionary monetary policy shock, the impact is not significant as in the case of the response of prices in Figure 8. However, compared to the results presented in Figure 8, the application of this identification method provides a very similar distributive effect of monetary policy, which is the focus of the current study.

Figure 9: Long Run Identification by Blanchard-Quah Method



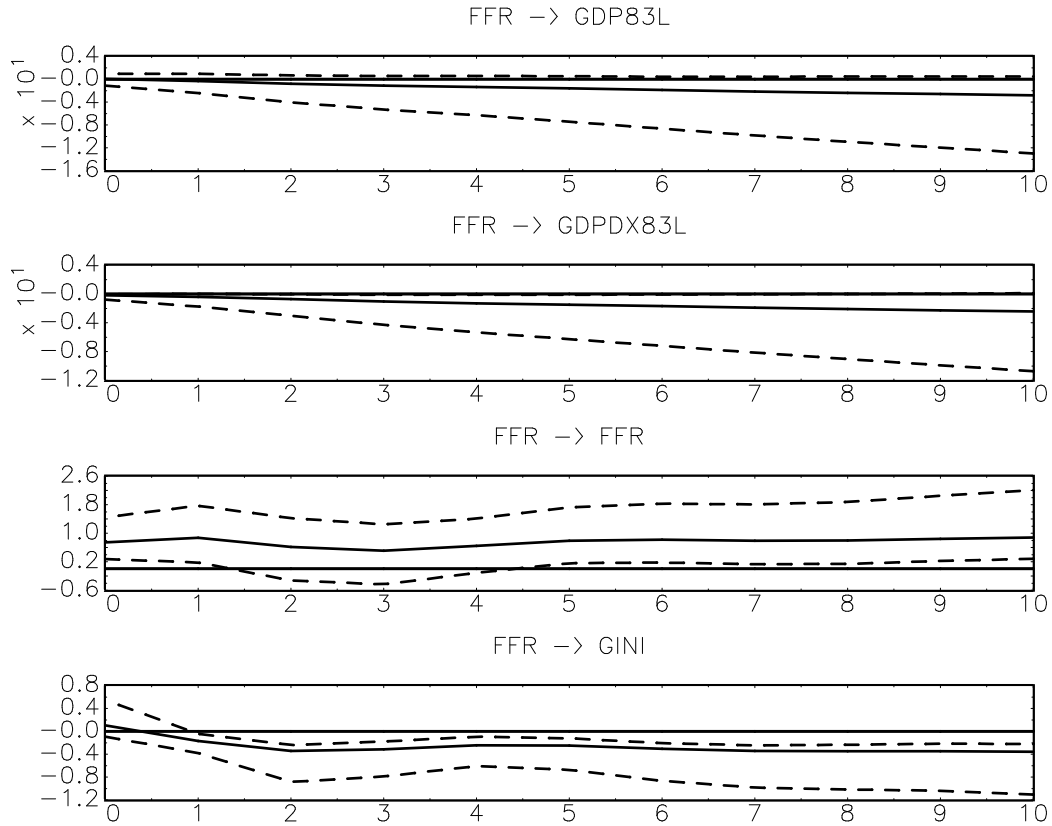
Since there is a cointegration relation among real GDP, prices, the federal funds rate, and Gini index, the IRFs can also be identified through the VECM methodology. As discussed in Section 5.1, the VECM of order one is specified with a constant included into the cointegration equation. They are identified by imposing restrictions on the contemporaneous impact matrix and the long run impact matrix as described in (8) of Section 3. The corresponding IRFs are presented in Figure 10. The impact of contractionary monetary policy shock is significant after one period when it reduces inequality by around 0.1 percentage points. Later, tight monetary policy decreases inequality by nearly 0.4 percentage points. Here the distributive impact of monetary policy is stronger than in the previous cases. After a contractionary monetary policy shock, the responses of prices and the federal funds rate are generally similar to the former results whereas real GDP significantly decreases following monetary policy tightening.

As a robustness check for the VECM identification, another set of restrictions is also imposed within this framework. As presented in (9) 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 11. Comparing them with the results presented in Figure 10, it can be observed that the IRFs to a monetary policy shock are actually identical in the both cases. In particular, a contractionary monetary policy shock decreases Gini index of income inequality up to around 0.4 percentage points.

In order to check the robustness of the results with respect to the estimation sample, the recent period when the federal funds rate reaches the zero lower bound is excluded from the sample. The VECM and the corresponding IRFs are re-estimated for this sample period until 2008 as in the case of the contemporaneous identification. The IRFs are identified by using the both sets of the restrictions of (8) and (9). The resulting IRFs are provided in Figures A1 and A2 in Appendix. As can be seen, the obtained results are generally very similar to the IRFs from Figures 10 and 11. Again, the estimated IRFs from the both identification schemes are almost identical. In this case of the shorter estimation sample, the responses of real output and prices to a monetary policy shock are just less significant. However, a contractionary monetary policy shock

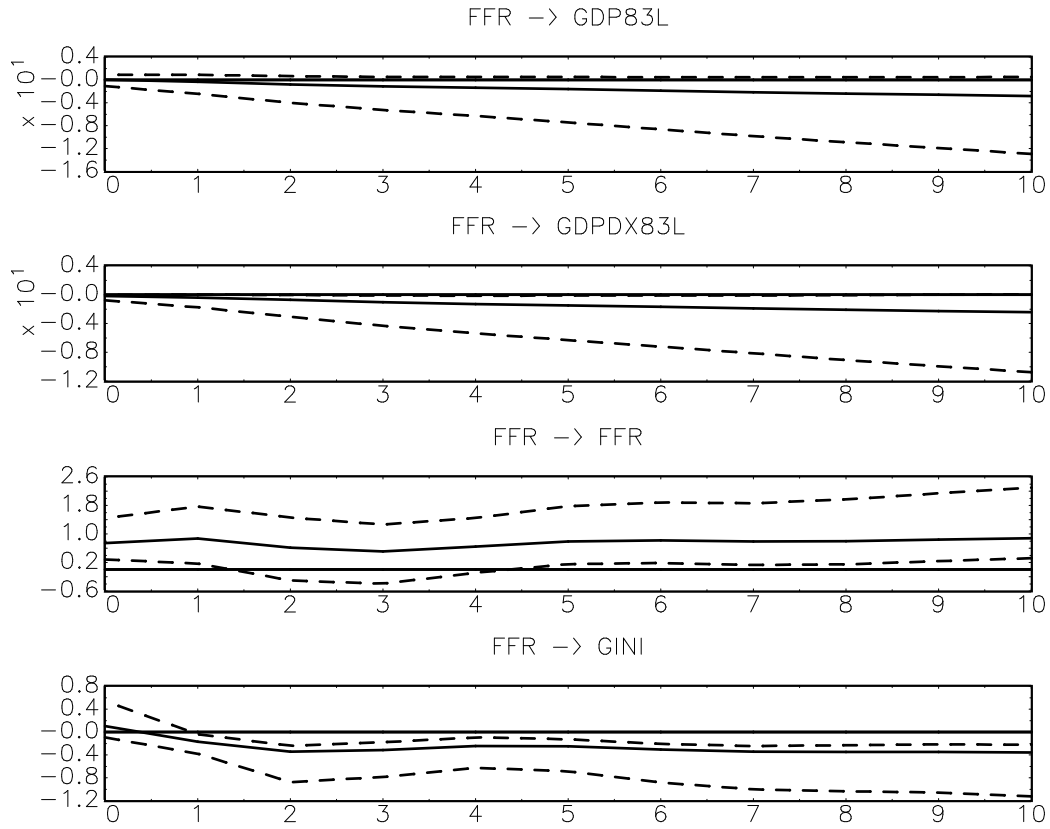
still significantly decreases Gini index of income inequality up to around 0.4 percentage points.

**Figure 10: Long Run Identification by Applying VECM Methodology
(Prices are Considered to Have Transitory Effects)**



Thus, in the all cases of the identification of a monetary policy shock, income inequality decreases following a tightening of monetary policy. This distributive effect of monetary policy is more pronounced in the case of long run identification with the VECM methodology, which is the benchmark analysis of this study. Gini index decreases up to around 0.4 percentage point after a contractionary monetary policy shock of one standard deviation. In addition, in the case of this identification, the responses of the other variables are better matched with theoretical implications and they are also significant.

**Figure 11: Long Run Identification by Applying VECM Methodology
(Income Inequality is Considered to Have Transitory Effects)**



6. Conclusion

The empirical analysis is implemented in accordance with the objective of the paper to evaluate the distributional effect of monetary policy. For the evaluation, the time series analysis for the USA is implemented using annual data. The inequality measure used in the paper represents the whole distribution of income. 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 and the equivalent VAR representation.

Different approaches are employed to identify a monetary policy shock and to analyze its impact on income inequality through the IRFs. First, exogenous monetary policy shocks (Romer and Romer, 2004; Coibion et al., 2012) are employed within the scheme of contemporaneous identification. Then, a long run identification approach proposed by Blanchard and Quah (1989) is applied in the analysis. The IRFs identified via these schemes show that contractionary monetary policy reduces income inequality, which is measured by Gini index and the 90-10 percentile ratio. Finally, taking advantage of the existence of the cointegration relation among the variables, the identification is implemented through the VECM framework. The obtained results indicate that a contractionary monetary policy shock decreases Gini index of income inequality up to 0.4 percentage points. Thus, the overall income inequality in the country could be reduced by implementing contractionary monetary policy and it might be considered as another effective policy instrument to decrease inequality.

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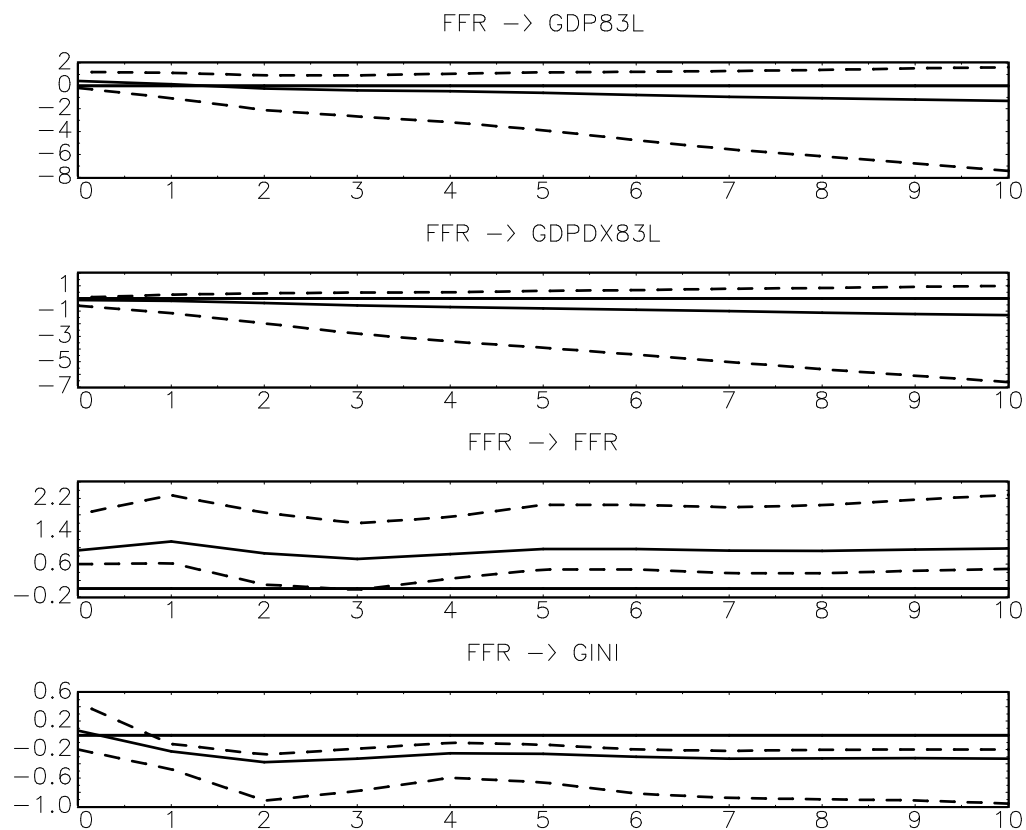
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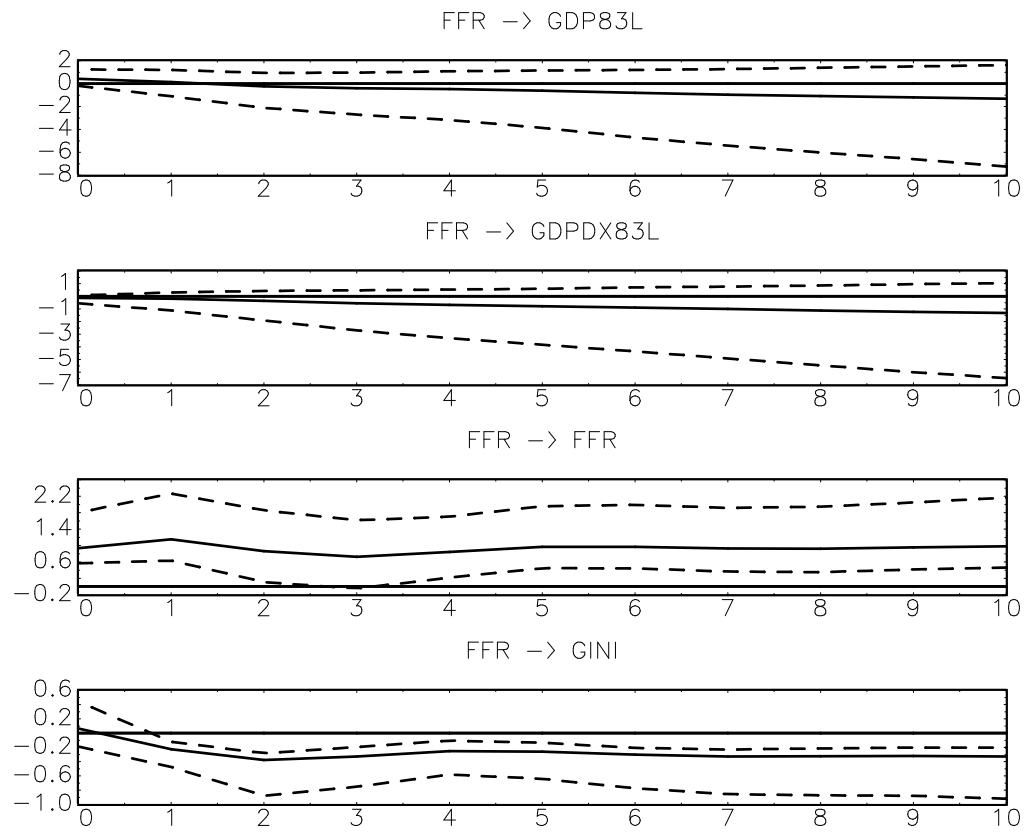
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Appendix: The IRFs Estimated by the VECM Identification in the Case of the Reduced Sample

Figure A1: Long Run Identification by Applying VECM Methodology (Prices are Considered to Have Transitory Effects); Reduced Sample



**Figure A2: Long Run Identification by Applying VECM Methodology
(Income Inequality is Considered to Have Transitory Effects);
Reduced Sample**





Institut de Recerca en Economia Aplicada Regional i Públic
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WEBSITE: www.ub-irea.com • **CONTACT:** irea@ub.edu



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