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Monetary Policy and Income Inequality in the United States: The Role of Labor Unions*

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Abstract

There is a growing body of literature investigating if and how monetary policy impacts income inequality. Labor unions are generally found to mitigate income inequality and recent literature highlights that changing labor market structures, such as de-unionization, may be important for monetary policy. This paper tests whether labor unions influence the impact of monetary shocks on income inequality in the United States over the period 1970-2008, and the channels this effect runs through. This is the first paper to identify variations in unionization rates as a moderator of the impact of monetary policy on income inequality. I measure income inequality and unionization at the state level and can therefore exploit that unionization rates vary both within and across states while monetary shocks are common to all states. The main finding is that contractionary monetary shocks increase income inequality, but the impact is weaker with a higher union density. A one percentage point monetary shock increases the Gini coefficient by 5.4% when union density is 5%, while it increases the Gini coefficient by 1.7% when union density is 15%. I find evidence that both wages and employment are two channels explaining how unions mitigate the monetary policy and income inequality relationship. These findings suggest that unions make adjustments to monetary shocks more even across workers, rather than mitigating the aggregate effect of the shocks.

JEL Classification: C23, D31, E24, E32, E52, J51.
Keywords: Monetary policy, monetary policy shocks, income inequality, labor unions, panel data, employment, wages.

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1 Introduction

Income inequality has soared in the United States (Piketty and Saez, 2003; Saez and Zucman, 2016; Piketty and Saez, 2006). Even though income inequality is not a direct concern for central banks, there is a growing literature investigating whether monetary policy impacts the income distribution. Monetary policy is considered neutral in the long run, but with interest rates trending downwards and central banks using unconventional policies after the financial crisis, there is a discussion on potential distributional impacts of monetary policy. Several papers find that monetary policy impacts income and earnings inequality (Coibion et al., 2017; Mumtaz and Theophilopoulou, 2017). Simultaneously with the upsurge in income inequality, there have been large structural shifts in US labor markets. The most prominent change is the decline in union density (Farber et al., 2018). Research partially links the increase in income inequality to the lower union density (Card et al., 2020).[1] Policymakers highlight that changing labor market structures is important for monetary policy as well (see e.g. address by Haldane at the 2018 Jackson Hole conference).

This paper studies whether labor unions influence the impact of monetary policy shocks on income inequality. The hypothesis is that labor unions impact the response of income to monetary shocks that in turn may impact the distribution of incomes. For example, when monetary policy contracts and aggregate demand falls, the impact of monetary shocks on income may be less severe when unions are stronger. Previous papers find that unions are particularly important for creating downward wage rigidity (Holden, 2004; Dickens et al., 2007). They may also be important for providing job security (see e.g., DiNardo and Lee, 2004, who does not find that unionization results in lower employment which is the conventional view). I propose that this potential interactive impact of monetary policy and labor unions on income inequality is related to union influence over labor earnings, which in turn can have implications for how monetary policy impacts income inequality.

This is the first paper to identify variation in unionization rates as a potential moderator of the impact of monetary policy on income inequality. Previous papers focus on how monetary policy decisions affect income in different groups of the population (see e.g. Coibion et al., 2017; Amberg et al., 2022; Andersen et al., 2021). This paper focuses on the role of labor unions. Theoretically, monetary policy impacts income inequality through an *earnings heterogeneity* channel — the sensitivity of labor income to monetary shocks may differ over the income distribution — and an *income composition* channel — individuals have multiple income sources (such as labor and capital) that react to monetary shocks in different magnitudes (Colciago et al., 2019). However, there is limited empirical testing of these channels. My contribution to the literature is to fill this gap by studying how the *earnings heterogeneity* channel works. I do this by testing whether unions impact the relationship

[1] The literature also brings forward technological change (Acemoglu, 2002) and globalization (Jaumotte and Osorio Buitron, 2015) as potential causes of the increase in income inequality.

between monetary policy and income inequality, and the channels through which unions should impact this relationship.[2]

I employ a dynamic panel data model with state-level data for the United States between 1970-2008. The United States is an ideal case for several reasons. First, there are large variations in union density over time and across states (Hirsch et al., 2001). Historically, union density has been highest in northeast states (such as Michigan and New York), and lowest in southern states (such as Texas and Louisiana). The decline in union density is most prominent in the north and northeast states (mostly due to the decline in the manufacturing sector). Second, union density measures union membership rates exclusively in the United States compared to Scandinavian countries for example, where many workers are covered by collective agreements but are not members of a union. Third, by using state-level data there should be fewer endogeneity issues from unobservable heterogeneity, compared to using countries as cross-sections. Fourth, the monetary shocks are the same across all states. I use monetary shocks to measure monetary policy, since other measures such as the actual federal funds rate are endogenous with economic activity. My main shock measure is the Romer and Romer (2004) monetary shocks which is widely used in the literature (see e.g., Coibion, 2012; Tenreyro and Thwaites, 2016; Miranda-Agrippino and Rey, 2020; Leahy and Thapar, 2019; Doniger, 2019, for other applications on economic outcomes)[3]. This paper focuses on short-term variation in inequality since monetary shocks cannot explain trends in income inequality.

The baseline model estimates the interaction effect between monetary shocks and union density on income inequality. The main measure of income inequality is the Gini coefficient on total income. I use state fixed effects to achieve identification using within-state variation in union density. I therefore compare the impact of monetary shocks on income inequality for various levels of union density within a state. To understand *how* unions influence the monetary policy and income inequality relationship, I estimate the interaction effect of monetary shocks and union density on two channels: real wages and employment. There is a possibility that unions impact income inequality through capital income as well, but due to the lack of capital income data, the focus of this paper will be on labor income. I perform various robustness checks including top income shares as measures of inequality, and monetary shocks identified with a high-frequency approach.

My results confirm previous findings that positive, or contractionary, shocks to monetary policy increase income inequality. However, the strength of the effect of a monetary shock on income inequality falls with higher union density. A one percentage point increase in monetary shocks increases the Gini coefficient (henceforth Gini) by 0.029 units, or 5.4% compared to the mean Gini, for a union member density of 5%, while it increases Gini by 0.009 units, or 1.7% compared to the

[2]Amberg et al. (2022) provides evidence of both channels when studying income responses to monetary shocks in Sweden. My contribution is to study a potential determinant of the earnings heterogeneity channel, by incorporating the role of labor unions.

[3]I use the series on monetary shocks updated through 2008 by Coibion et al. (2017).

mean, for a union member density of 15%. Hence, unions mitigate the inequality effect of monetary shocks, and they do so for a union density of up to 16%. This finding is robust to alternative measures of monetary shocks, income inequality and subsample periods. Impulse responses from a local projection method show that monetary shocks increase income inequality up to three years after the shock hits, and that the relationship between monetary shocks and income inequality is weaker for higher levels of union density for all horizons up to five years. I also find evidence of asymmetric effects — unions have a stronger influence on the relationship between monetary shocks and income inequality for contractionary compared to expansionary shocks.

In a second set of results, I confirm that both wages and employment are two channels through which unions influence the monetary policy and income inequality relationship. I find that union density significantly reduces the positive impact on both real wages and employment when the economy is hit by a contractionary monetary shock. Connecting these findings with the baseline results on inequality, I argue that unions make the effect of monetary shocks more evenly spread out among workers, rather than mitigating the aggregate effect of the shocks on income inequality.

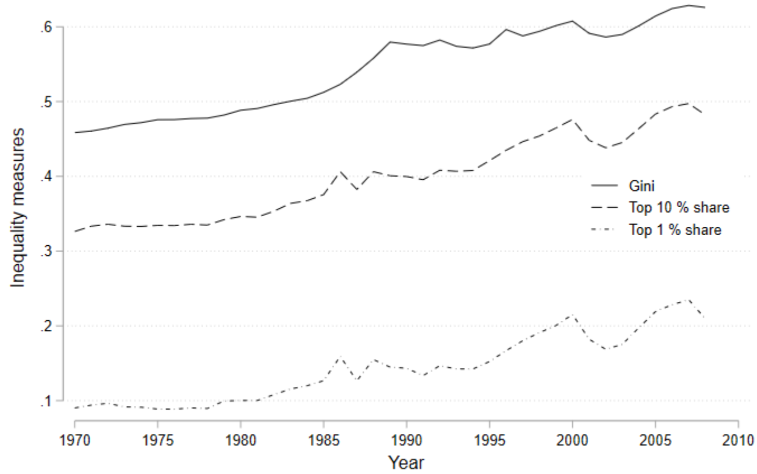
The outline of the paper is as follows. Section 2 provides a short background to income inequality and labor unions in the United States. Section 3 provides a link between monetary policy, labor unions and income inequality through a brief review of theoretical links and previous research. Section 4 provides the empirical strategy including the model specification and a description of the data. Section 5 provides the results, including the baseline results on income inequality, extensions and robustness checks, as well as findings for the potential channels on wages and employment. Lastly, section 6 concludes.

2 Income inequality and labor unions in the United States

2.1 Income inequality in the United States

Figure 1 illustrates the development of Gini and top income shares between 1970 and 2008 in the United States. Gini increased by 36.5% from 0.46 to 0.63 in this period. Top income shares increased with similar magnitude. The literature provides various explanations to this upsurge in inequality, including skill-biased technological change (Heathcote et al., 2020), trade and financial globalization (Jaumotte et al., 2013; Dabla-Norris et al., 2015), as well as changes in labor market institutions (Western and Rosenfeld, 2011; Jaumotte and Osorio Buitron, 2015). More recently, monetary policy has been added to the list (Coibion et al., 2017; Furceri et al., 2018). The increase in income inequality from the 1970s is often explained by an upsurge in top incomes. Before World War II, top incomes were mainly comprised of capital income. In the 1970s a shift occurred and top incomes mainly comprised of labor income (Piketty and Saez, 2003, 2006). More recent evidence shows another shift as capital income explains the upsurge of top incomes since the late 1990s (Piketty et al., 2018; El Herradi and Leroy, 2021). The left tail of the income distribution is important as well. Heathcote et al. (2020) find that the bottom half of the male earnings distribution has been the main driver of higher inequality since 1968, explained by declining hours worked in recessions.

Figure 1: Income inequality in the United states, 1970-2008

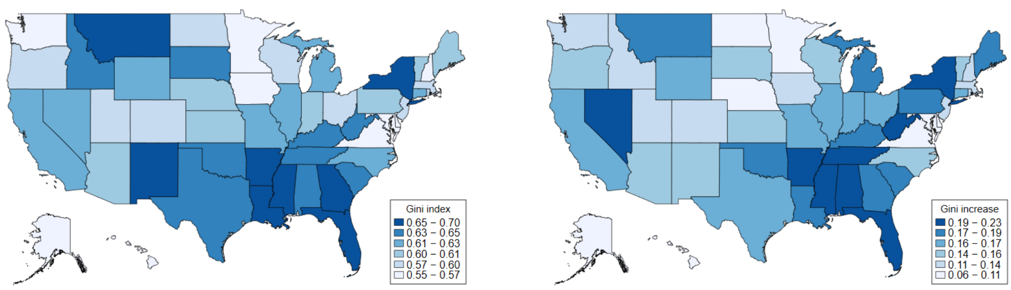


Notes: Inequality measures are from Frank (2009). The Gini index measures income inequality for total income using data from individual tax filing data. Total income includes wages and salaries, capital income and entrepreneurial income. The top 10% and 1% share is the pre-tax national income share held by the highest 10% and 1% of income earners.

Figure 2 Panel (a) presents the level of income inequality in each state in 2008. Inequality was higher in southern states such as Florida, Georgia and Arkansas, but also in some northern states such as Montana and New York. Figure 2 Panel (b) illustrates that these states also had the largest increase in inequality between 1970-2008.

Figure 2: State-level income inequality measured with the Gini index

(a) Gini in 2008 **(b)** Increase in Gini 1970-2008

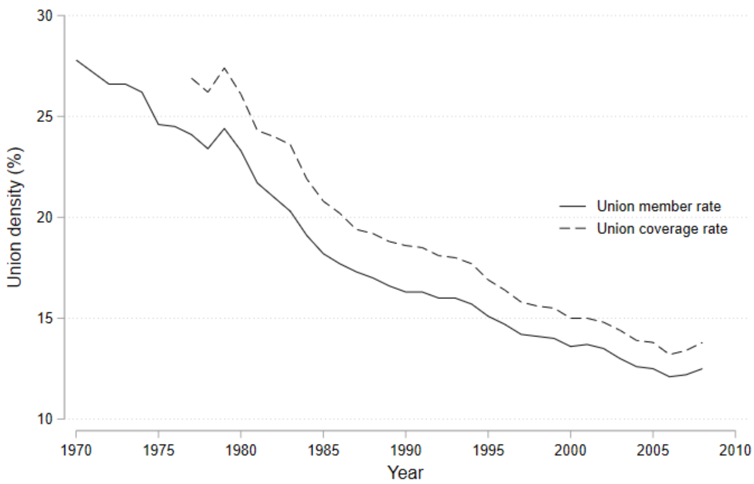


Notes: Panel (a) shows Gini by state for 2008, which is the last year of the sample. Panel (b) shows the unit increase in Gini between 1970 and 2008.

2.2 Labor unions in the United States

Union membership rates have declined since the 1970s. Figure 3 plots the union membership rate between 1970 and 2008 (solid line). The sharpest decline is from 1980 and forward where average union density declined from 23% to 13% in 2008.[4] A potential explanation is the changing composition of jobs in the United States, including the decline of the union-dominated manufacturing sector, and the rise of service jobs with historically low unionization (Heathcote et al., 2010; Farber and Western, 2001). There is also *right-to-work* legislation that allows states to prohibit agreements between labor unions and employers that require union membership and/or union fees (Jacobs and Dixon, 2006).[5] This means that all employees have the same rights as union members, whether they are members of a union and/or pay union fees. So far, 27 states have adopted *right-to-work* legislation (as of 2022), including states where unions have been historically strong such as Michigan and Wisconsin (Fortin et al., 2022).[6]

Figure 3: Union density for the United States, 1970-2008



Notes: The union member rate measures the percentage of non-agricultural wage and salary employees, including public employees, who are union members. The union coverage rate includes workers who are covered by a union contract. The data is from Hirsch et al. (2001).

The union membership rate and the union coverage rate in Figure 3 do not differ to any large extent, and the gap between the two series has not increased over time. Hence, my main measure of union strength, the union member density, exclusively measures the bargaining power of union members.[7]

[4] I refer to the union membership rate and union density interchangeably throughout the paper.

[5] The right of employees to create and engage in labor unions and collective bargaining is protected in the National Labor Relations Act (NLRA), signed into law in 1935. *Right-to-work* legislation was founded in the Taft-Hartley Act in 1947, aimed at restricting the power of labor unions, and has since been enacted by separate states (Jacobs and Dixon, 2006; Fortin et al., 2022).

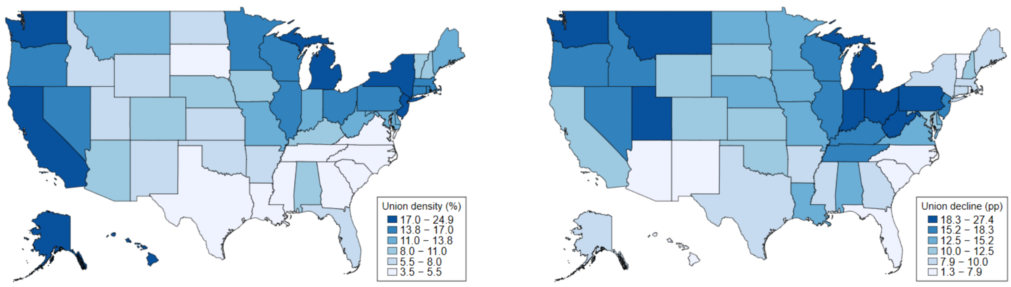
[6] Proponents argue that *right-to-work* legislation increase workers' rights, and opponents argue that they encourage freeriding and undermine unions in a workplace by reducing membership numbers (Jacobs and Dixon, 2006).

[7] The coverage rate refers to both union members and workers who report no union affiliation

Figure 4 illustrates state-level union densities. Panel (a) shows the union density for each state in 2008. There are 27 states with union membership rates below the US average of 12.5%, and 24 states above it. New York has the highest union membership rate of 24.9%. Five states have union membership rates below 5% including Georgia, Texas, Louisiana, Virginia, South Carolina and North Carolina (which has the lowest rate of 3.5%). All these states have enacted *right-to-work* legislation. Panel (b) shows the percentage point decline in union density between 1970 and 2008. There is a larger decline in the northern states compared to the southern states, likely reflecting the decline in manufacturing industries in the north (Carnevale et al., 2019).

Figure 4: State-level union densities in the United States

(a) Union density in 2008 **(b)** Decrease in union density 1970-2008



Notes: Panel (a) shows union density by state for 2008, which is the last year of the sample. Panel (b) shows the percentage point change in union density between 1970 and 2008. Union density measures the percentage of non-agricultural wage and salary employees (including public sector employees) who are union members. The data is from the Current Population Survey (CPS) and the Bureau of Labor Economics (BLS) (retrieved from Hirsch et al., 2001).

3 Linking monetary policy, labor unions and income inequality

How labor unions can moderate the relationship between monetary policy and income inequality links to three strands of literature. I provide a literature overview, and the theoretical channels, for all three; how monetary policy impacts income inequality, how labor unions impact income inequality and how the interaction between monetary policy and labor unions can impact income inequality.

3.1 Monetary policy and income inequality

Addressing inequality is not a direct objective of monetary policy, but both theoretical and empirical papers have pointed to (unintended) distributional consequences of monetary policy. Monetary policy can impact inequality in opposite directions, since there are several transmission channels (see Coibion et al. (2017) and Colciago et al. (2019) for a full review of the redistribution channels). It is possible to decompose the monetary policy transmission on inequality into direct and indirect effects

but whose jobs are covered by a union or an employee association contract (Hirsch et al., 2001).

(Ampudia et al., 2018). Theoretically, higher policy rates should increase income inequality directly through higher interest incomes, since high-income households hold relatively more financial assets and deposit savings (Ampudia et al., 2018; Colciago et al., 2019).

The literature distinguishes between two theoretical channels relating to the indirect impact on prices, wages, output and employment from monetary shocks (Samarina and Nguyen, 2019; Ampudia et al., 2018). The *income composition channel* refers to heterogeneity across individuals in their primary income sources. If monetary shocks are expansionary, income inequality should increase since high-income households receive relatively more business and capital income, which tends to rise relative to wages after an expansion (Coibion et al., 2017). The *earnings heterogeneity channel* captures that earnings for high- and low-income households may respond differently to monetary shocks (see also Auclert, 2019). Heathcote et al. (2010, 2020) show that changes in relative wages mainly affect earnings at the top of the distribution, while changes in hours worked or unemployment mainly affect earnings at the bottom. If unemployment falls mainly on low-income groups after a contractionary shock, then inequality should increase.[8] Similar effects may arise from wage rigidities across the income distribution, caused by unions or different skill sets among the employees (Coibion et al., 2017).

Among papers that empirically test if monetary policy impacts income inequality, one early contribution comes from Romer and Romer (1999) who find that expansionary monetary policy mitigates poverty in the short run. More recent evidence points towards higher income inequality from contractionary monetary shocks. Coibion et al. (2017) use individual-level survey data for US households and find that contractionary monetary shocks significantly increase inequality in labor earnings, total income, consumption and total expenditures. Similar results are found for the UK (Mumtaz and Theophilopoulou, 2017) and the euro area (Guerello, 2018; Samarina and Nguyen, 2019). Furceri et al. (2018) find that contractionary monetary shocks increase income inequality in the short- and medium-term for a panel of advanced and emerging economies.

There is evidence pointing to monetary policy effects in the opposite direction. Inui et al. (2017) find that expansionary monetary shocks result in higher income inequality in Japan. Cloyne et al. (2016) find that expansionary shocks raise incomes for mortgage holders more than for other groups in the US and the UK, which can increase inequality. Andersen et al. (2021) use Danish administrative data and show that lower policy rates increase inequality in disposable incomes, by raising income shares at the top and reducing them at the bottom. A similar study by Amberg et al. (2022) using Swedish administrative register data shows that expansionary monetary shocks increase incomes for low- and high-income individuals relative to middle-income individuals, resulting in small inequality effects when using aggre-

[8] Heathcote et al. (2020) show that declining hours worked and unemployment mainly occur in recessions and do not fully recover when the business cycle changes. Therefore recessions, cycles, cause a persistent increase, trend, in earnings inequality.

gate inequality measures such as Gini. The response of labor income is strongest at the bottom of the distribution, which can be accounted for by the earnings heterogeneity channel, while the response of capital income is strongest at the top of the distribution, which can be accounted for by the income composition channel.[9] Similarly, O'Farrell and Rawdanowicz (2017) find no distributional effects of monetary shocks for a panel of OECD countries.

In summary, monetary policy can impact income inequality in both directions. If monetary policy is contractionary, the direct effect should be lower income inequality through lower interest income. Indirect effects point to lower inequality from the *income composition channel* but higher inequality from the *earnings heterogeneity channel*. The empirical literature finds evidence of the latter channel in the US, UK and euro area, as contractionary policy seems to increase income inequality. However, there is empirical evidence of the opposite direction as well.

3.2 Labor unions and income inequality

There is a well-documented inverse relationship between income inequality and union membership in the United States. A large literature studies the relationship between labor unions and income inequality (see e.g., Card et al., 2020; Farber et al., 2018, for a review of the literature).

Labor unions can impact the level and distribution of income in several ways. First, unions have a direct impact on member wages through pay bargaining, which can result in a higher wage premium (the differential in wages between union and similar non-union workers) (Fortin et al., 2022). Unions can also limit wage reductions in economic downturns, relative to uncovered workers (Freeman, 1980). Second, unions can decrease within-group wage inequality by reducing the spread of wages among union members with similar characteristics, and between-group wage inequality by reducing educational and occupational inequality among union members (Kristal and Cohen, 2017). Third, unions can impact non-union members' wages through spillover effects and threats, or setting a fairness norm in an industry (Western and Rosenfeld, 2011).[10] Fourth, unions can impact the compensation of management and returns to capital, thus reducing inequality by lowering compensation in the upper part of the income distribution (Lee and Mas, 2012; DiNardo et al., 2000). Fifth, unions can impact income inequality through political mechanisms through lobbying (Acemoglu and Robinson, 2013). Previous papers highlight that the effect of unions on wages and income inequality depends on the skill group, gender, and sector analyzed (Lemieux, 1998; Card, 1996, 2001; Card

[9]Contrary to most previous studies, Andersen et al. (2021) and Amberg et al. (2022) use household-level outcomes instead of summary measures of inequality such as the Gini coefficient (since they use administrative register data).

[10]Spillover effects refer to when unions increase wages and employers respond by lowering employment. Employees are then forced into the non-union sector where wages will fall as the labor supply increases. Threats refer to when non-union workplaces increase wages to the union level to avert the threat of unionization (Western and Rosenfeld, 2011; Farber et al., 2018).

et al., 2020).

Several papers establish that labor unions decrease income inequality (DiNardo et al., 1996; Card, 2001; Western and Rosenfeld, 2011; Farber et al., 2018). Jaumotte and Osorio Buitron (2015) study OECD countries and find that lower membership rates explain almost 30% of the increase in income inequality. Farber et al. (2018) finds similar results when analyzing the states in the United States. Some papers use individual-level data to estimate the impact of unions on workers' wages and income. This literature mainly focuses on estimating the union premium and generally finds a positive one (Farber et al., 2018; Card et al., 2020). The general challenge is to capture a causal effect of union membership on the level and distribution of income.[11] Card (1996) addresses the issue of selection into union membership by examining workers as they switch from a union sector to a non-union sector using panel data. He finds evidence of a union premium even when accounting for the selection. DiNardo and Lee (2004) use a regression discontinuity design on firm-level data, comparing firms where a union was formed with one marginal vote and firms where a union was not formed with one vote short. The effects of unions on wages are close to zero.

In summary, unions can impact the distribution of income directly through wage bargaining and indirectly via political mechanisms and fairness norms (impacting also non-union members). Previous findings show that the decrease in union membership rates partly explains the upsurge in income inequality since the 1970s.

3.3 The interaction between monetary shocks and union density

Several papers investigate heterogeneous responses to monetary policy. Previous research mainly focus on how household balance sheets and participation in credit markets affect the transmission of monetary policy (Cloyne et al., 2016; Kaplan et al., 2018; Auclert, 2019). However, there is less knowledge on whether the transmission of monetary policy depends on union density. Both policymakers and academics have raised that structural shifts of the US labor market, including the decline in union density and labor's share of GDP (Autor et al., 2020), should have implications for monetary policy (see address by Haldane, 2018). There is, however, limited research on how it would impact the transmission of monetary policy. I argue that union density should affect the transmission of monetary policy on income inequality. My motivation for this interaction effect of monetary shocks and unionization on income inequality is that unions can impact the response of labor income to monetary shocks, which in turn have implications for the distribution of income. My hypothesis is:

[11] The country-level studies suffer from omitted variable bias, since developments such as globalization and skill-biased technological change can explain both the increase in inequality and the decrease in union power (Acemoglu, 2002) The individual-level studies suffer from selection problems since people self-select into union membership, explaining differences in wages and income. Also, studies that estimate the union premium often assume that unions only affect its members' wages, which us unlikely (Farber et al., 2018).

H1: Unions moderate the relationship between monetary policy shocks and income inequality, and they do so by influencing labor income.

For example, when the economy is hit by a contractionary monetary shock and aggregate demand falls, the adjustment to both wages and employment may be less severe when union bargaining power is larger. Previous papers find that unions are particularly important to establish downward wage rigidity (Holden, 2004; Dickens et al., 2007). The impact of unions on employment is debated. On the one hand, unions can negotiate higher real wages, or less real wage drops in economic downturns, at the expense of fewer jobs (Jaumotte and Osorio Buitron, 2015). On the other hand, some papers find that unions increase productivity, which leads to lower unemployment (Barth et al., 2016). Unions also influence employment where the conventional view is that unionization decrease employment (due to higher wages). DiNardo and Lee (2004) show that the effect of unionization on employment is close to zero. The interaction of union density and monetary policy shocks may, therefore, show heterogeneous responses to income inequality stemming from different union densities.

One of the indirect channels through which monetary policy impacts income inequality captures responses to macroeconomic variables such as wages and employment (Samarina and Nguyen, 2019). If unions influence the monetary policy and income inequality relationship, the effect should go via this impact on labor income. There is some evidence of heterogeneous responses to wages and employment from monetary policy (but not related to union density). Doniger (2019) interacts monetary shocks with educational attainment and finds that responses to hourly earnings from monetary shocks do not differ for different educational attainment, while employment is more sensitive to monetary policy shocks for high school dropouts (compared to having a bachelor's degree).[12] Leahy and Thapar (2019) find that monetary policy is more effective (responses to income and employment are stronger when interest rates change), the greater the share of the middle-aged population and less effective the greater the share of the young population.

In summary, I motivate the interaction effect between monetary policy and union density on income inequality by unions' impacts on labor earnings, having implications for income inequality. Previous papers find that monetary policy impacts income inequality, but labor unions may influence this relationship through its impact on earnings. I identify both wages and employment as the channels through which the interaction effect should operate.

[12] Wages increase with higher education using a measure of allocative wage, similar to a user cost.

4 Empirical strategy

4.1 Model specification

To test if labor unions influence the effect of monetary policy shocks on inequality, I employ a dynamic panel data model with annual state-level data for the United States between 1970-2008. I use annual data since inequality measures are not available for higher frequencies. Appendix Figure A1 and A2 present impulse responses for output, prices, consumption and unemployment, when hit by a monetary shock, at a monthly and yearly frequency. The direction and magnitude of the effects do not change much when going from a monthly to yearly frequency, showing that an annual frequency is possible to use.[13]

I use state-level data because there is large variation in union density both across and within states over time (see Figure 4 in Section 2.2). Also, the monetary shocks are identical for all states, making them normalized across states. Lastly, state-level data reduces endogeneity problems from unobservable heterogeneity across states within the country (compared to using countries as cross-sections). I estimate the following baseline model:

$$
\begin{aligned}
Y_{i,t} = {} & \beta_0 Y_{i,t-1} + \beta_1 MP_{t-1} + \beta_2 Union_{i,t-1} + \beta_3 MP_{t-1} \times Union_{i,t-1} \\
& + \beta_4 X_{i,t-1} + \beta_5 F_{t-1} + \alpha_i + \epsilon_{i,t}
\end{aligned} \tag{1}
$$

where $Y_{i,t}$ measures income inequality for state i and year t. The main coefficient of interest is β_3 since it captures the interaction effect between monetary policy shocks and union density. MP_{t-1} is the national level exogenous shocks to monetary policy and $Union_{i,t-1}$ is the state level union density. In the baseline model, I measure income inequality with the Gini index for total income, monetary policy with the Romer and Romer (2004) monetary shocks, and union density with the percentage of employees who are union members. I lag all variables by one year, as it is unlikely that there is an contemporaneous impact on inequality.

The vector $X_{i,t-1}$ includes controls for potential determinants of income inequality. To capture the relationship between human capital and income inequality (see e.g. Autor et al., 2008), I include the share of *college* graduates of the total state population. To capture fiscal policy, I use total *taxes* as a proxy for income taxation and *expenditure of GDP* which measures the total expenditure by the state government.[14] To capture asset prices, I include a house price index, *HPI*. Lastly, I control for the size of each state by including total *population* by state. As a

[13] I also run the monetary shocks on state-level employment at a monthly, quarterly and annual frequency, shown in Appendix Table A1. The shock coefficient increases with an annual frequency compared to monthly and quarterly frequencies. This is reasonable given that the variation in the data is smaller at a lower frequency and that the shock effect is not instant. Signs and significance levels are the same regardless of the frequency.

[14] The data on income taxation is missing for some states and the data on total and income taxation are highly correlated.

robustness check, I add total *exports* by state, since trade is brought forward as one explanation to the increase in income inequality in the United States.[15] I also control for the state-level minimum wage as one robustness check. All controls are in growth rates to induce stationarity.

Since there are some time-invariant differences across states, I include state fixed effects α_i. Identification is achieved using within-state variation in union density, comparing the impact of monetary shocks on income inequality for varying levels of union density (after controlling for observables) within states. The model excludes time-fixed effects because monetary shocks are omitted otherwise, being common to all states. Even though the shocks to monetary policy are exogenous by design, there may be some common developments over time that all states share. Another way to control for this is to include national-level control variables in the vector F_{t-1}. In the baseline model, I include the log change of GDP and unemployment at the national level. As a robustness test, I include principal components based on a large amount of macroeconomic time series at the national-level. The errors, $\epsilon_{i,t}$, are clustered at the state-level to correct for heteroskedasticity and autocorrelation.

I also estimate model (1) with $Y_{i,t}$ equal to the log change in average real wages (by worker) and employment (for wage and salary workers), respectively. Based on the discussion in section 3.3, these are the main channels through which the interaction between monetary shocks and union density should impact income inequality.

I use a dynamic model to account for persistence in the data. However, a bias arises with a one-way fixed effects model, as the demeaning process creates a correlation between the lagged dependent variable and the error term (Nickell, 1981; Bond, 2002). The inconsistency of the coefficient for the lagged dependent variable decreases with T though (Nickell, 1981; Flannery and Hankins, 2013). Since T is large ($= 38$) in this paper, the bias should be small (Bruno, 2005a). To take this into consideration, I use a bias-corrected least square dummy variable estimator (LSDVC). Instead of computing valid instruments, as with the Arellano and Bond (1991) difference GMM and the Blundell and Bond (1998) system GMM estimators, it computes a correction for the fixed effects bias (Kiviet, 1995; Bruno, 2005b). Monte Carlo results from Bruno (2005b) show that the corrected LSDV estimator outperforms the difference and system GMM estimators in terms of bias and errors even when N and T are equally large (see also Bruno et al., 2017; Buddelmeyer et al., 2008; Judson and Owen, 1999).[16] It requires a vector of coefficient starting values where the Anderson-Hsiao (AH), Arellano-Bond (AB) and Blundell-Bond (BB) estimators are common options. I only report results derived from AB estimates of the initial coefficient matrix, since the LSDVC estimates are robust to the initial matrix selection (Bun and Kiviet, 2001).

[15]This data is available from 1999. The reason for including exports, and not imports, is because the import data is only available from 2008.

[16]An issue with the difference and system GMM estimators is that the instrument count grows larger relative to the sample size when T increases, which can overfit endogenous variables and bias coefficient estimates (Roodman, 2009; Windmeijer, 2005).

4.2 Data

This paper combines several datasets. Inequality measures come from Frank (2009). The data includes annual state-level income inequality measures that is constructed from individual tax filing data from the International Revenue Services (IRS). Income includes wages, salaries, capital income and entrepreneurial income. The main measure of income inequality is the Gini index. As alternative measures of income inequality, I study the effect on the income share of the top 10% and top 1% from Frank (2009). For the channels, I use data on average real wages per worker and total employment. The wage and employment data are from the *Bureau of Economic Analysis* (BEA).

To measure monetary policy, I use the Romer and Romer (2004) shocks to monetary policy. Monetary shocks are necessary since conventional measures of monetary policy, such as the actual federal funds rate, is endogenous with economic activity (Christiano et al., 1999). I use the Romer and Romer (2004) shocks because they are widely used in the literature (see e.g., Coibion, 2012; Miranda-Agrippino and Rey, 2020; Tenreyro and Thwaites, 2016) and they are available from 1969. Romer and Romer (2004) identify the shocks by regressing intended federal funds rate changes on *Greenbook* forecasts of output growth, inflation, and unemployment, to separate the endogenous response of policy to information about future economic development from exogenous shocks.[17] The residuals are the resulting shocks to monetary policy.[18] These shocks capture over- or under-reactions, shifts in preferences of policymakers, or deliberately induced policy surprises (Romer and Romer, 2004; Cloyne and Hürtgen, 2016). I use the series on monetary shocks updated through 2008 by Coibion et al. (2017).[19]

As a robustness test, I use the monetary shocks estimated with a high-frequency identification approach by Nakamura and Steinsson (2018) (N&S) and Gertler and Karadi (2015) (G&K). Both papers measure monetary shocks as changes in interest rates of different maturities over a 30-minute window surrounding scheduled FOMC announcements.[20] The N&S shocks start in 2000 and the G&K shocks start in 1990. For the overlapping period, the N&S and G&K shocks are similar. They are, however, smaller in magnitude compared to the Romer and Romer shocks (see Appendix figure A3). For all three types of shocks, the annual shocks are a simple average of the original shocks at a quarterly frequency.

I use state-level data on union density from the Current Population Survey (CPS)

[17] They derive a series of intended funds rate changes around meetings of the Federal Open Market Committee (FOMC).

[18] Romer and Romer (2004) identify their monetary shocks ϵ_m, by estimating the following equation: $\triangle i_m = \alpha + \beta i_{m-1} + \sum_{i=-1}^{2} \gamma_i \tilde{x}_{mi} + \sum_{i=-1}^{2} \delta_i \triangle \tilde{x}_{mi} + \rho u_{m0} + \epsilon_m$, where $\tilde{x}_{mi}$ is a vector including forecasts on GDP growth and inflation at the FOMC meeting date m for horizon i, $\triangle \tilde{x}_{mi}$ is the corresponding change in the forecast since the previous meeting and u_{m0} is the current forecast for unemployment.

[19] These shocks are close to identical to the Romer and Romer (2004) shocks up until 1996.

[20] Nakamura and Steinsson use changes in a composite measure of interest rates including the federal funds rate, federal funds futures and eurodollar futures, while Gertler and Karadi use changes in either federal funds futures or eurodollar futures.

and the Bureau of Labor Economics (BLS) to measure union strength (retrieved from Hirsch et al., 2001). Union density measures the percentage of non-agricultural wage and salary employees (including public sector employees) who are union *members*. The Appendix Table A4 lists all variables included in the analysis, descriptive statistics and sources.

5 Results

When presenting the results, I start with a parsimonious model including only the monetary shocks. This facilitates a comparison with previous papers. I then present the results for the full model (1). I add several extensions and robustness checks, such as analyzing the interaction effect for positive and negative shocks separately, as well as using alternative measures for monetary shocks. I end the section by analyzing the wage and employment channels.

5.1 Baseline results

Table 1 presents the baseline regression results for model (1). I estimate specifications (1)-(5) with a standard FE-estimator, specification (6) with a LSDVC estimator using the AB estimator to estimate the initial coefficient matrix, and specification (7) with a system GMM.[21] I include all three estimators to verify that the bias from using a dynamic fixed effects model is small, since T is large. When describing the results, I refer to positive or contractionary shocks interchangeably (or negative and expansionary shocks).[22]

Columns (1)-(2) present results for the parsimonious model, estimating the impact of monetary shocks on income inequality. In (1), I find that a one percentage point increase in the monetary shock in the previous period, significantly increases within state inequality by 0.056 units in the current period. This finding is confirmed in specification (2), including a lagged dependent variable in line with Coibion et al. (2017). The shock coefficient decreases in (2), as I find that a one percentage point increase in the monetary shock significantly increases within-state inequality by 0.016 units. This corresponds to a 3% increase from mean Gini. The largest monetary shock in the sample is 0.168 percentage points, which corresponds to an increase in Gini of 0.6% from the mean.

These results confirm the findings in Coibion et al. (2017) and Furceri et al. (2018), that contractionary monetary shocks increase income inequality. Comparing the coefficient with previous findings is not straightforward, since most studies rely on impulse responses from local projections or VAR models when estimating the impact

[21] I collapse the instrument matrix to reduce the risk of instrument proliferation in line with Roodman (2009). The instrument count (= 60) is close to the number of cross-sections and the Hansen J-test of instrument validity does not show implausibly perfect p-value close to one. This indicates that the model is not subject to instrument proliferation (Roodman, 2009).

[22] Positive shocks reflect contractionary policy, while negative shocks reflect expansionary policy.

on Gini from monetary shocks using quarterly data.[23] Coibion et al. (2017) find that a cumulative shock of 1.5 percentage points increases Gini by 1.5 percentage points for the US. Mumtaz and Theophilopoulou (2017) find that a 10 basis point increase in the short interest rate increases Gini for incomes by 0.3% in one year for the UK. My results point to an effect size somewhere in between these previous findings. For long term effects, I estimate impulse responses from local projections in Section 5.1.1.

Table 1: Baseline results

	Dependent variable: Gini						
	(1) FE	(2) FE	(3) FE	(4) FE	(5) FE	(6) LSDVC	(7) SGMM
$\text{Gini}_{i,t-1}$		0.876***	0.840***	0.837***	0.834***	0.865***	0.899***
		(0.013)	(0.017)	(0.017)	(0.017)	(0.011)	(0.012)
$MPshock_{t-1}$	0.056***	0.016***	0.016***	0.050***	0.040***	0.039***	0.036***
	(0.005)	(0.003)	(0.004)	(0.007)	(0.007)	(0.009)	(0.009)
$Union_{i,t-1}$			-0.068***	-0.077***	-0.073***	-0.057***	-0.016**
			(0.010)	(0.011)	(0.012)	(0.012)	(0.007)
$MPshock_{t-1}$*$Union_{i,t-1}$				-0.180***	-0.205***	-0.201***	-0.190***
				(0.040)	(0.042)	(0.043)	(0.054)
$\text{Taxes}_{i,t-1}$					-0.004	-0.004	-0.002
					(0.004)	(0.004)	(0.005)
Expeditures of $\text{GDP}_{i,t-1}$					-0.018	-0.012	-0.012
					(0.048)	(0.039)	(0.052)
$\text{HPI}_{i,t-1}$					0.003	0.003	0.012**
					(0.006)	(0.005)	(0.006)
$\text{College}_{i,t-1}$					0.004	0.004	0.002
					(0.004)	(0.005)	(0.004)
$\text{Population}_{i,t-1}$					-0.094*	-0.093**	-0.060*
					(0.049)	(0.036)	(0.034)
GDP_{t-1}					-0.187***	-0.187***	-0.181***
					(0.024)	(0.030)	(0.024)
$\text{Unemployment}_{t-1}$					-0.488***	-0.496***	-0.491***
					(0.056)	(0.066)	(0.056)
Constant	0.485***	0.063***	0.095***	0.098***	0.106***		0.061***
	(0.001)	(0.006)	(0.010)	(0.010)	(0.010)		(0.007)
Observations	1950	1950	1950	1950	1950	1900	1950
Number of states	50	50	50	50	50	50	50
R-squared	0.767	0.947	0.948	0.949	0.950		
State FE	YES	YES	YES	YES	YES	YES	YES
Dummy 1987	YES	YES	YES	YES	YES	YES	YES
Adjusted R-square	0.767	0.947	0.948	0.949	0.950	0.950	0.953
DW	0.683	1.771	1.784	1.776	1.794		
AR(1) p-value						0.000	0.000
AR(2) p-value						0.472	0.265
# Bootstraps						1000	
# Instruments							60
Hansen J p-value							0.365

Note: Levels of significance: p<0.01, *** p<0.05, ** p<0.1 *. The sample consists of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. The dependent variable is the Gini index on total income. I report robust standard errors (in parentheses) for specifications (1)-(5) and bootstrap robust standard errors for specifications (6)-(7). In the bias corrected OLS estimation, in (8), I base the initialization on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. In the system GMM estimation in (9), I collapse the instrument matrix to reduce the risk of instrument proliferation. I include a dummy taking the value 1 from 1987 and forward to account for a structural break in the Gini index and union density. DW is the Durbin Watson statistic and is calculated for the FE estimator. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation, where the null is no serial correlation. Hansen J is a test for instrument validity, where the null is that instruments are valid/exogenous.

[23] I primarily rely on single equation estimations since my state-level data is available at an annual frequency only.

Column (3) adds union density and I find that a higher union density significantly lowers income inequality. This is in line with previous findings (see e.g., Farber et al., 2018). In column (4) I add the interaction between monetary shocks and union density. The results show that the positive relationship between monetary shocks and income inequality is significantly weaker when union density is higher. Hence, the unionization rate moderates the positive relationship between monetary shocks and inequality. The effect is robust to the inclusion of controls in (5) and different estimators in (7)-(8). When comparing the results for the different estimators in columns (5), (6), and (7), they are similar. Only the union density coefficient shows a lower level of significance in column (7), but the interaction effect is of a similar size and significance level.[24] Hence, the bias arising from the dynamic setting of the model is small (relating to the large T). Therefore, I only report results for the LSDVC-estimator in (6) further on. The Appendix Table A6 presents the findings for specification (6), but with contemporaneous impacts of all variables. The interaction variable is no longer significant. This confirms the choice of one lag in the model.

Only population significantly impacts income inequality among the state-level controls. A larger population in a state decreases Gini. The national level controls, GDP and unemployment, significantly lower Gini. I find no presence of serial correlation in the baseline model presented in column (7), since I cannot reject the null of no second-order serial correlation indicated by the AR(2) p-value.[25] In all specifications, I add a dummy variable to account for a structural break in both the Gini index and union density.[26] Adjusted R-square values are high in all specifications, relating to the inclusion of the lagged dependent variable and the dummy correcting for the structural break. If I exclude lagged Gini, the adjusted R-square drops to 0.77 and if I exclude the dummy, the adjusted R-square drops to 0.11.

Figure 5 presents the marginal effect of the monetary shocks given the union density. The coefficients are from specification (6) in Table 1. The coefficient for the interaction variable gives the slope of the curve and establishes my baseline finding: the significant positive relationship between monetary shocks and Gini is weaker with a higher union density. The positive impact of the shocks on Gini is significant

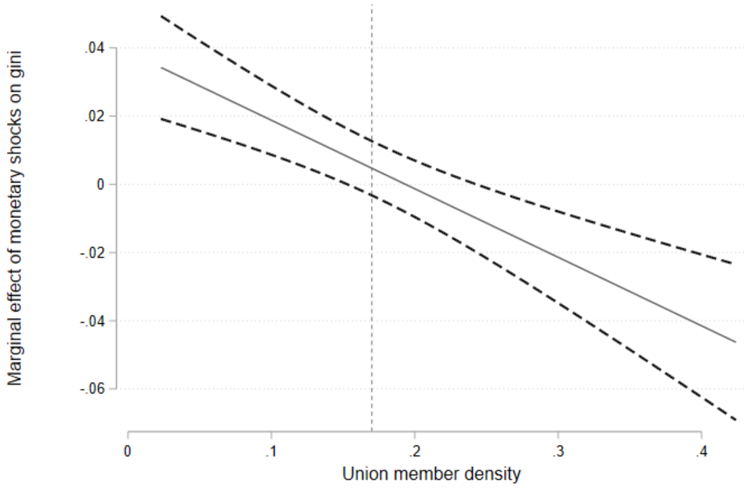
[24]Previous papers highlight that the relationship between union density and income inequality may suffer from endogeneity problems. Results do not change if I take this into account by instrumenting both lagged Gini and union density using a system GMM approach like the one in specification (7) in Table 1.

[25]The AR(1) p-value shows that we can reject the null of there being no first order serial correlation, which is expected since the estimation is done in differences. The AR(2) p-value shows that there is no second-order serial correlation in first differences, meaning that there is no first-order serial correlation in levels, which is good.

[26]A set of unit root tests reveal that both the Gini index and union density are non-stationary. When regressing Gini on a constant in a pooled regression and running a test for a structural break, the null of no structural break is rejected at the year 1987. Therefore, I include a dummy variable equal to one for the period 1987 and forward. When regressing the dummy variable on Gini and union density separately, it is positive and significant on the 1% level. When I include the dummy variable, the unit root tests reject the null hypothesis of a unit root at the 1% level for the corrected series. See Appendix Table A2 and A3 for results on the structural break and stationarity test. The results are not sensitive to changing the start date for the dummy variable to any year between 1985-1989.

and decreases up to a union density of approximately 16% (which is close to the mean of the union density, illustrated with the vertical dashed line). As an example, a one percentage point contractionary shock in the previous period increases Gini by 0.029 units (5.4% compared to the mean Gini) in the current period, when the union member density is 5% in a state, while Gini increases by 0.009 units (1.7% compared to the mean Gini) when the average union member density is 15% in a state. Figure 5 also shows that the impact of the shocks on Gini is negative for high values of union density (above 25%). However, the number of observations with a union density that high are few (see Appendix Figure A4).

Figure 5: The impact of monetary shocks on Gini for varying union densities



Notes: Dashed lines give 95% confidence intervals. The vertical dashed line shows the mean density of union membership, which is equal to 17%. The results are based on the coefficients in specification (6) in Table 1.

Auclert (2019) argues that redistribution is not only a side effect of policy changes, but a channel through which monetary policy affects the macroeconomy (due to different marginal propensities to consume along the income distribution). These findings, therefore, not only inform the debate on the sometimes unintended consequences of monetary policy but also facilitate the understanding of how labor market structures, and more specifically unions, may impact the effectiveness of monetary policy.

5.1.1 Local projections

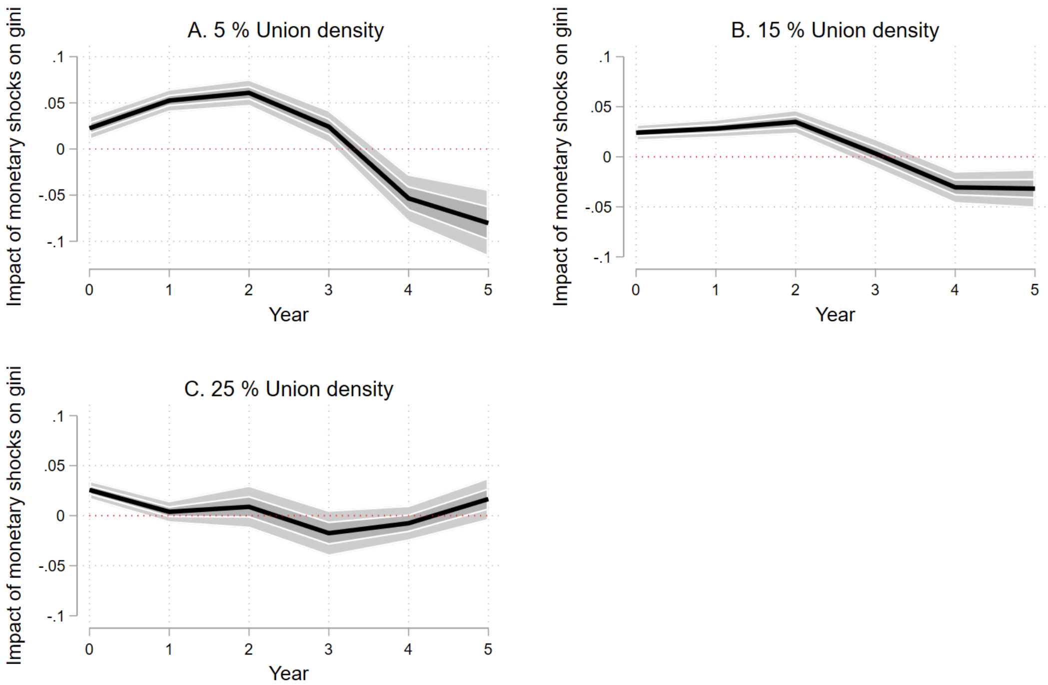
I follow previous literature and complement the analysis by estimating impulse responses of monetary shocks, union density, and the interaction variable at different horizons h. I follow Jordá (2005) and use a local projection method where I estimate the following model;

$$
\begin{aligned}
Y_{i,t+h} = \alpha_i^h &+ \sum_{j=1}^{J} c_{i,j}^h Y_{i,t-j} + \sum_{k=0}^{K} \beta_{i,k}^h MP_{t-k} + \sum_{k=0}^{K} \gamma_{i,k}^h Union_{i,t-k} \\
&+ \sum_{k=0}^{K} \rho_{i,k}^h MP_{t-k} \times Union_{i,t-k} + \sum_{m=0}^{M} \theta_{i,m}^h X_{i,t-m} + \sum_{m=0}^{M} \omega_{i,m}^h F_{t-m} + \epsilon_{i,t}^h
\end{aligned}
\tag{2}
$$

where $h = 0,...,5$ horizons. Y_i is the Gini index on total income and X_i and F includes the same control variables as in the baseline model (1). I follow Coibion et al. (2017) and use a lag structure of $J = 2$, $K = 5$ and $M = 2$. Impulse responses of the main variables (shocks, union density and the interaction variable) from estimates of model (2) only provide information about the sign of the coefficients since I include an interaction variable. The Appendix Figure A5 presents the impulse responses for monetary shocks (A), union density (B), and the interaction between the shocks and union density (C). The total impact of the monetary shocks however is the sum of the coefficient for the monetary shocks and the interaction variable, conditional on the union density.

To understand how the total impact of the shocks on income inequality differs depending on the union density, I present three graphs in Figure 6. These show the total marginal effect of the shocks on Gini for different horizons when the union density is 5% (Panel A), 15% (Panel B), and 25% (Panel C). Panel A and B illustrate a positive relationship between monetary shocks and income inequality up to three years after the shock, but this positive relationship is weaker with a higher union density (the impact of the shocks on Gini is smaller when the union density is 15%). The impact is strongest two years after the shock. Four to five years after the shock, there is a negative impact of the shocks on Gini, but this relationship is less negative with a higher union density. For a high union density of 25% in Panel C, the positive impact of the shocks on income inequality is even weaker and the relationship turns negative after three years. However, the confidence bands cross zero for all periods except the first.

Figure 6: Impulse responses of the impact of monetary shocks on inequality



Notes: The graph plots impulse responses of the total marginal effect of monetary shocks on Gini when the union density is 5% (A), 15% (B) and 25% (C). Light and dark grey shaded areas are 95% and 68% confidence bands.

Table 2 presents the coefficients and standard errors for each horizon and the three main variables: the monetary shocks, the union density and the interaction variable. Using the coefficients, it is possible to confirm the findings in Figure 6. Given a union density of 15%, a contractionary shock increases Gini by 0.03 units one year after the shock (=0.065-0.243*0.15) and by 0.04 units two years after the shock. Then the impact of the shocks on Gini decreases to 0.004 units three years after the shock, until the impact on Gini turns negative four years after the shock. For a higher union density, the impact of the shocks on Gini is weaker, as in baseline. Overall, the findings from the local projection analysis coincide with the baseline findings using single equation estimations. That the impact of the shocks on inequality turns negative three to four years after the shock, regardless of the level of union density, likely relates to the usage of annual data.

Table 2: Results for local projection analysis

	Year 0	Year 1	Year 2	Year 3	Year 4	Year 5
MP shock	0.022**	0.065***	0.074***	0.035***	-0.065***	-0.105***
	(0.008)	(0.008)	(0.010)	(0.012)	(0.017)	(0.024)
Union density	-0.072***	-0.094***	-0.148***	-0.227***	-0.230***	-0.239***
	(0.023)	(0.031)	(0.034)	(0.038)	(0.042)	(0.042)
MP shock*Union density	0.016	-0.243***	-0.260***	-0.208***	0.230***	0.485***
	(0.040)	(0.036)	(0.070)	(0.073)	(0.079)	(0.116)
Observations	1750	1750	1750	1750	1750	1750
Number of states	50	50	50	50	50	50
State FE	YES	YES	YES	YES	YES	YES
Dummy 1987	YES	YES	YES	YES	YES	YES
R^2	0.955	0.913	0.859	0.808	0.774	0.739

Notes: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The sample consists of all US states (except the District of Columbia due to data limitations) for the period 1970-2008. The dependent variable is the Gini index on total income. I report bootstrap robust standard errors in parentheses.

5.1.2 Asymmetric effects of monetary shocks

The shock effect may differ depending on monetary policy being contractionary or expansionary. Previous papers find that the effect of positive monetary shocks (contractionary monetary policy) on economic activity are larger than the effect of negative monetary shocks (expansionary monetary policy) (see e.g., Furceri et al., 2018). One explanation for this asymmetry in the monetary policy transmission relates to credit market imperfections, suggesting that higher interest rates result in less investment by less liquid firms (Bernanke and Blinder, 1992; Kashyap and Stein, 2000). I want to test whether the interaction effect between the shocks and union density differs for expansionary and contractionary shocks. The hypothesis is that unions should moderate the relationship between monetary shocks and inequality more strongly when shocks are contractionary compared to expansionary, since unions create downward wage rigidity (Dickens et al., 2007). This is tested by splitting the monetary shocks into contractionary and expansionary in the following model:

$$\begin{aligned} Y_{i,t} = {} & \beta_0 Y_{i,t-1} + \beta_1 MP_{t-1} + \beta_2 Union_{i,t-1} + \beta_3 MP_{t-1} \times Union_{i,t-1} \\ & + \beta_4 MP_{t-1} \times D_{i,t-1} + \beta_5 Union_{t-1} \times D_{i,t-1} \\ & + \beta_6 MP_{t-1} \times Union_{i,t-1} \times D_{i,t-1} + \beta_7 X_{i,t-1} + \beta_8 F_{t-1} + \alpha_i + \epsilon_{i,t} \end{aligned} \tag{3}$$

where $D_{i,t-1}$ is a dummy variable that takes value one if the shocks are contractionary (i.e., positive) and zero otherwise (i.e., negative). The vectors $X_{i,t-1}$ and F_{t-1} include the same variables as in baseline. Table **??** presents the regression output. When union density is equal to zero, I find that an expansionary shock to monetary policy decreases Gini (since the shock is negative) by 0.023 units (β_1), while a contractionary shock increases Gini by 0.072 units ($\beta_1 + \beta_4 = 0.023 + 0.049$). The impact on Gini is stronger for contractionary shocks compared to expansionary shocks, and my results confirm the findings in Furceri et al. (2018).

The results show that the interaction effect between the shocks and union density differs for expansionary and contractionary shocks. When monetary shocks are expansionary ($D_{i,t-1} = 0$), a one percentage point increase in the union density results in a 0.059% stronger impact of the shocks on Gini ($\beta_3 = 0.059$). The interaction effect is insignificant though. When shocks are contractionary ($D_{i,t-1} = 1$), a one percentage point increase in the union density results in a 0.49% lower impact of the shocks on Gini ($\beta_3 + \beta_6 = 0.059$ - 0.550).

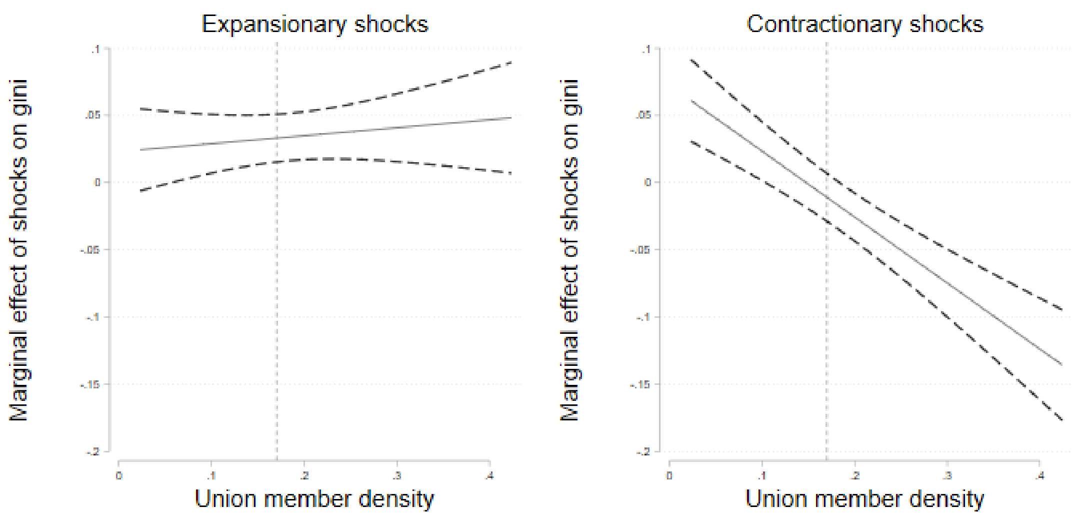
Table 3: Regression output for positive and negative shocks

Dependent variable: Gini	
$\text{Gini}_{i,t-1}$	0.866***
	(0.011)
$MPshock_{t-1}$	0.023
	(0.017)
$Union_{i,t-1}$	-0.023*
	(0.013)
$MPshock_{t-1}*Union_{i,t-1}$	0.059
	(0.079)
$MPshock_{t-1}*Dummy_{i,t-1}$	0.049
	(0.030)
$Union_{t-1}*Dummy_{i,t-1}$	-0.011**
	(0.005)
$MPshock_{t-1}*Union_{i,t-1}*Dummy_{i,t-1}$	-0.550***
	(0.158)
$\text{Taxes}_{i,t-1}$	-0.002
	(0.004)
Expenditures of $\text{GDP}_{i,t-1}$	0.006
	(0.038)
$\text{HPI}_{i,t-1}$	0.001
	(0.005)
$\text{College}_{i,t-1}$	0.004
	(0.005)
$\text{Population}_{i,t-1}$	-0.083***
	(0.037)
GDP_{t-1}	-0.232***
	(0.033)
$\text{Unemployment}_{t-1}$	-0.535***
	(0.068)
Observations	1900
Number of states	50
State FE	YES
Dummy 1987	YES
Adjusted R-square	0.952
AR(1) p-value	0.000
AR(2) p-value	0.178

Notes: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The sample consists of all US states (except the District of Columbia due to data limitations) for the period 1970-2008. The dependent variable is the Gini index on total income. $Dummy_{i,t-1}$ takes the value one if monetary shocks are positive and zero otherwise. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS (LSDVC) on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

Figure 7 illustrates the impact of expansionary and contractionary monetary shocks on Gini for different union densities. I use β_1, β_2 and β_3 to plot the impact of expansionary shocks in the left graph. A minus sign must be added to the y-axis for a correct interpretation of the effect, since the shocks are negative (or expansionary). I find that an expansionary monetary shock decreases Gini and that a higher union density makes this relationship between the shocks and inequality stronger. The curve is relatively flat though, so unions do not influence the monetary policy and inequality relationship to any large extent when shocks are expansionary (in line with the insignificant interaction coefficient (β_3) in Table **??**). I find a stronger relationship for contractionary shocks as the slope of the curve is steeper (-0.491 compared to 0.059 for expansionary shocks). A higher union density decreases the positive relationship between monetary policy and income inequality, when shocks are contractionary. These results show that unions have a stronger influence on the relationship between the shocks and inequality when shocks are contractionary compared to expansionary. This confirms the hypothesis that unions are especially important when the economy contracts (Dickens et al., 2007). Heathcote et al. (2020) show that declining earnings (through fewer hours worked) is the main driver of higher income inequality in the US since the 1970s, and the decline in earnings is concentrated in recessions. This paper shows that unions mitigate the inequality effect of monetary shocks more so when the policy is contractionary compared to expansionary.

Figure 7: The impact of monetary shocks on Gini for varying union densities - positive and negative shocks separately



Notes: Dashed lines give 95% confidence intervals. The vertical dashed line shows the mean density of union membership, which is equal to 17% in both graphs. The results are based on the coefficients in Table **??**. For the correct interpretation of the marginal effects of the left graph, a minus sign must be added to the marginal impacts, since the shocks are negative (or expansionary). As an example, when union density is equal to 10%, a one percentage point decrease in the monetary shocks, decreases Gini with 0.025 units.

5.2 Potential channels explaining the relationship

To examine potential channels, I estimate model (1) with the log change in average real wages per worker and total employment, both by state, as dependent variables.

Table 4 presents the regression output.[27] When the union density is zero, a positive monetary shock in the previous period significantly increases both wages (by 0.053%) and employment (by 0.051%), within states in the current period. I also find that unions decrease this positive relationship since the interaction variable between the shocks and union density is significant and negative. Lastly, when the monetary shocks are zero, I find that a higher union density has a negative impact on wages and no significant impact on employment.

Table 4: Regression output using wages and employment as dependent variables

	Dependent variables:	
	(1) Average real wages	(2) Employment
Dependent variable$_{i,t-1}$	0.543***	0.562***
	(0.020)	(0.013)
$MPshock_{t-1}$	0.053***	0.051***
	(0.012)	(0.015)
$Union_{i,t-1}$	-0.080***	0.008
	(0.014)	(0.016)
$MPshock_{t-1} * Union_{i,t-1}$	-0.331***	-0.269***
	(0.060)	(0.065)
Taxes$_{i,t-1}$	-0.019***	-0.007
	(0.005)	(0.006)
Expenditures of GDP$_{i,t-1}$	-0.173***	-0.330***
	(0.053)	(0.058)
HPI$_{i,t-1}$	-0.015**	-0.011
	(0.006)	(0.007)
College$_{i,t-1}$	-0.003	-0.005
	(0.007)	(0.007)
Population$_{i,t-1}$	0.053	-0.113***
	(0.050)	(0.054)
GDP$_{t-1}$	-0.115***	0.678***
	(0.042)	(0.045)
Unemployment$_{t-1}$	0.505***	1.516***
	(0.091)	(0.100)
Observations	1900	1900
Number of states	50	50
State FE	YES	YES
Dummy 1987	YES	YES
Adjusted R-square	0.304	0.400
AR(1) p-value	0.000	0.000
AR(2) p-value	0.000	0.006

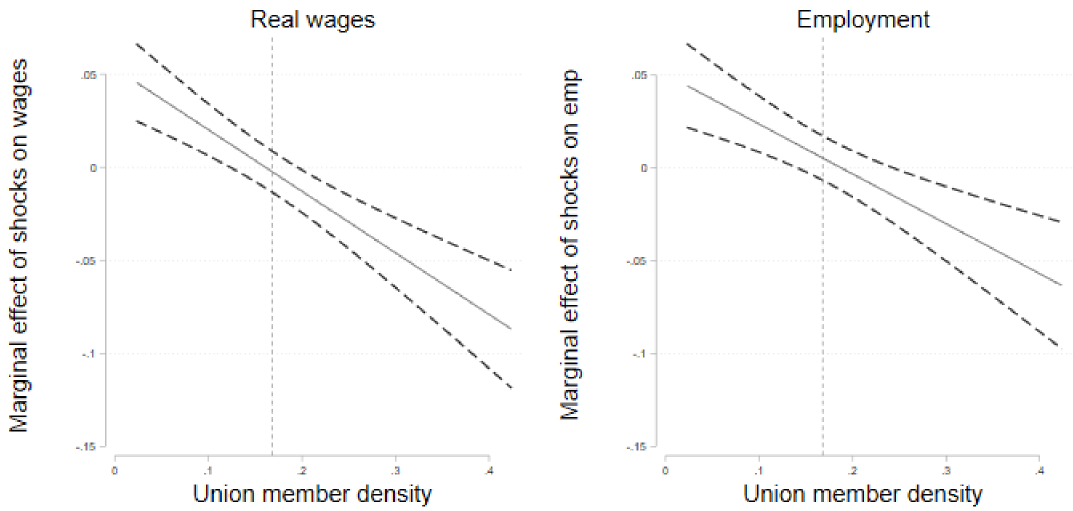
Note: Levels of significance: p<0.01, *** p<0.05, ** p<0.1 *. The dependent variables are the log change in average real wages (1) and total employment (2). The sample consist of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

I plot the effect of the shocks on wages and employment for various union densities in Figure 8. The left graph illustrates that monetary shocks increase real wages when the union density is below mean, and decrease wages when union densities are above the mean. Unions therefore mitigate the positive and significant relationship between monetary shocks and wages below the mean union density, while it amplifies

[27] If I use the log change in total real wages, results are similar to the ones for average wages, but coefficients are somewhat larger.

the negative and significant relationship above the mean. Findings are similar in the right graph for employment responses. A monetary shock in the previous period significantly increases employment in the current period, but this positive relationship is lower when the union density increases and turns negative for union densities above the mean.[28] There are few observations with a union density that high though (see Appendix Figure A4).

Figure 8: The impact of monetary shocks on real wages and employment for varying union densities



Notes: Dashed lines give 95% confidence intervals. The vertical dashed line shows the mean density of union membership, which is equal to 17%. Real wages measure the log change of average real wage and salary income per worker; and employment measures the log change of total employment. The coefficients are from specifications (1)-(2) in Table 4.

That contractionary monetary shocks increase real wages is in line with several previous empirical studies finding that real wages respond pro-cyclically to monetary shocks (see e.g., Christiano et al., 2005; Basu and House, 2016). However, the wage response is typically insignificant in previous papers. My results are inconsistent with the findings in Samarina and Nguyen (2019), which show a negative impact on employment when monetary shocks contract. Neither of these papers include interactions between monetary shocks and labor unions though. The impact of the shocks on employment may be quicker. The Appendix Table A7 shows that the contemporaneous impact of the shocks on wages is positive (but insignificant) and negative (weakly significant) on employment when the union density is zero. Unions significantly weaken the impact of the shocks on wages, but the interaction variable is insignificant for employment. Hence, the impact is weaker compared to the baseline model with one lag.

How do the results on wages and employment relate to the baseline findings on

[28]The results are similar if I add two lags to model (1). The Appendix Table A8 presents the results. The total effect is similar to the short-run findings, but the size of the interaction effect is smaller for the wage regression and larger for the employment regression.

income inequality? One possibility is that when union density is low, the positive impact on wages and employment stems from larger gains in wages and employment for workers belonging to the upper part of the income distribution, compared to those belonging to the lower end of the distribution.[29] This explains the larger positive impact on income inequality when union density is low. When union density is higher the positive impact on wages and employment are smaller (or even negative), explaining the smaller impact on income inequality. One interpretation is that unions adjust shocks more evenly among workers. When the economy is hit by a contractionary monetary shock and aggregate demand falls, firms need to adjust. Intuitively, wage growth should be lower or at least slower and employment should fall for some workers. When union density is low, the net effect on both wages and employment is positive, so the adjustment must fall on groups belonging to the lower part of the income distribution since I find a stronger positive impact on income inequality. When union density is higher it is more difficult for firms to let the adjustment fall disproportionally on certain groups, and the impact of the shock is more evenly spread out. This explains the negative net effect on wages and employment and the lower impact on income inequality.

Combining the baseline findings with the analysis on wages and employment, I draw two conclusions. First, I explain the lower inequality effect of monetary shocks when union density is high by suggesting that unions spread the effects of the shocks more evenly among workers rather than mitigating the aggregate effect of monetary shocks on income inequality. Second, I conclude that labor income is an important channel through which unions impact the monetary policy and income inequality relationship.

5.3 Robustness

5.3.1 Subsample periods

The transmission mechanism of monetary policy may change over time, having implications for the transmission on income inequality. Both Romer and Romer (2004) and Coibion et al. (2017) highlight the Volcker disinflation period in the beginning of the 1980s as an important macroeconomic event.[30] The monetary shocks are particularly large during this period. I estimate model (1) for two subsample periods: 1970-1984 and 1985-2008, where the first sample period includes the larger monetary shocks and the second does not. I also estimate model (1) for the full sample excluding the Volcker disinflation period in 1980-1984 in line with Coibion

[29] The impact on wages and employment among workers at the low end of the income distribution can even be negative, but if the gains in the top of the distribution are larger, the net effect is positive.

[30] The Volcker disinflation period describes disinflationary measures taken when the Fed was headed by Paul Volcker, where inflation was reduced from 9% in 1979 to 4% in 1984 (Goodfriend and King, 2005).

et al. (2017).[31]

Table 5 presents the regression output. For the first subsample period in (1), the coefficients on the main variables are similar when compared to the baseline in (6) in Table 1. For the second subsample in (2), the interaction variable is no longer significant, so I do not find that union density alters the monetary policy and income inequality relationship between 1985-2008. This insignificant result should be explained by the lower variation of the shocks in this time period, as illustrated in Appendix Figure A3.[32]

Table 5: Regression output for subsample periods

	Dependent variable: Gini		
	(1) 1970-1984	(2) 1985-2008	(3) 1970-2008, excluding 1980-1984
$\text{Gini}_{i,t-1}$	0.730***	0.872***	0.895***
	(0.025)	(0.017)	(0.047)
$MPshock_{t-1}$	0.050***	0.022	0.052*
	(0.016)	(0.013)	(0.027)
$Union_{i,t-1}$	-0.115***	-0.079***	-0.103***
	(0.021)	(0.025)	(0.033)
$MPshock_{t-1}$*$Union_{i,t-1}$	-0.216***	0.027	-0.266***
	(0.066)	(0.089)	(0.100)
$\text{Taxes}_{i,t-1}$	0.006	-0.016***	-0.006
	(0.006)	(0.005)	(0.008)
Expenditures of $\text{GDP}_{i,t-1}$	0.010	-0.042	-0.013
	(0.064)	(0.051)	(0.085)
$\text{HPI}_{i,t-1}$	0.007**	0.002	-0.001
	(0.006)	(0.009)	(0.008)
$\text{College}_{i,t-1}$	-0.016*	0.010**	0.008
	(0.008)	(0.005)	(0.014)
$\text{Population}_{i,t-1}$	-0.004	-0.152***	-0.095
	(0.065)	(0.058)	(0.105)
GDP_{t-1}	-0.081*	-0.315***	-0.191***
	(0.047)	(0.043)	(0.056)
$\text{Unemployment}_{t-1}$	-0.199***	-0.673***	-0.503***
	(0.102)	(0.105)	(0.116)
Observations	700	1150	1650
Number of states	50	50	50
State FE	YES	YES	YES
Adjusted R-square	0.707	0.893	0.941
AR(1) p-value	0.000	0.000	0.000
AR(2) p-value	0.298	0.413	0.374

Note: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The sample consists of all US states (except the District of Columbia due to data limitations) and the dependent variable is the Gini index on total income. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

The results are similar to baseline when excluding the Volcker disinflation period in the full sample (3). If anything, uncertainty increases somewhat as the shock coeffi-

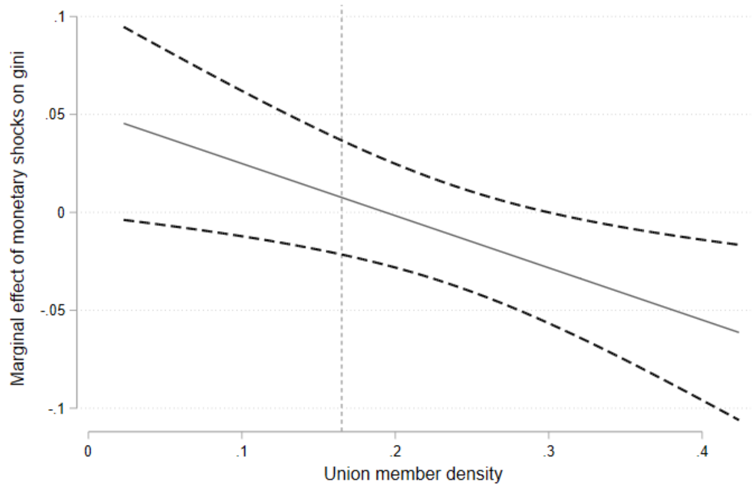
[31] The results are similar if I change the periods to 1970-1985 and 1986-2008, and exclude 1980-1985 in the full sample period.

[32] Boivin et al. (2010) find a smaller impact on output after the Volcker period when estimating a VAR over the pre- and post-Volcker periods.

cient is significant at the 10% level. This aside, the direction, size, and significance level of the interaction variable is similar to baseline. I conclude that my baseline findings are robust to the exclusion of the Volcker disinflation period.

Figure 9 plots the marginal effect of the monetary shocks given the union density, excluding the period 1980-1984. The lower significance level of the shock variable is notable with wider confidence bands. It is still apparent that unions mitigate the inequality effect of monetary shocks though. The more imprecise estimates of the monetary shocks likely relate to the reduced variation in the shocks when excluding the Volcker disinflation period.

Figure 9: The impact of monetary shocks on Gini for varying union densities - excluding the Volcker disinflation period



Notes: Dashed lines give 95% confidence intervals. The vertical dashed line shows the mean density of union membership, which is equal to 17%. The results are based on the coefficients in specification (3) in Table 5.

5.3.2 Potential covariates of labor union intensity

With state-fixed effects, the analysis effectively relies on within-state variation in labor union density. However, some factors covary with labor union intensity and could affect inequality responses to monetary policy. First, labor union intensity reflects industrial composition in each state (e.g., manufacturing vs services). I consider this by including an additional interaction variable between the shocks and the GDP share of either the manufacturing or the agricultural sector. Historically, the manufacturing sector has been more unionized, and the agricultural sector less so. The Appendix Table A9 shows that the sign and size of the interaction coefficient between the shocks and union density do not differ to any large extent in any of the specifications compared to baseline. This suggests that the inequality response of monetary shocks is not solely driven by industrial composition.

Second, labor union intensity reflects political preferences. When a state adopts policies that weaken labor unions, other economic reforms are likely implemented, such as right to work laws, changes in taxation, and minimum wages as some examples. I try to account for this by adding an interaction between the monetary shocks and an economic freedom index. I use the index developed by Stansel et al. (2020) from the Fraser Institute, including government spending (general consumption expenditures by government, transfers, subsidies, as well as insurance and retirement payments), taxation (income and payroll tax revenue, top marginal income tax rate, as well as property and sales taxes), and labor market freedom (full-time minimum wage income, government employment, and union density). The index is at the state-level and is available from 1982. Specification (3) in Appendix Table A9 presents the findings. The interaction between the shocks and the union density is insignificant when adding the economic freedom (EF) index interaction. This likely relates to two issues. The first is the sample starting in 1982. In the previous subsample analysis, the interaction between the shocks and union density was also insignificant when starting the sample in 1985, explained by the lower variance in the shocks from the mid-1980s. Second, the index takes labor market freedom into account, creating a correlation between the index and union density. Hence, it may be difficult to separate the interaction effects.

Lastly, I control for the state minimum wage since it may be correlated with union density. I argue that it is less obvious that the impact of the shocks should be affected by the minimum wage, adding it as a control variable only. The results do not change the baseline findings to any large extent (see specification (4) in Appendix Table A9). The interaction variable is only significant at the 5% and the size of the coefficient drops somewhat, but I still draw the same conclusion that union density decreases the positive impact of monetary shocks on Gini.

5.3.3 Alternative measures of monetary shocks

I use two alternative measures of monetary shocks: the Nakamura and Steinsson (2018) (N&S) shocks available between 2001-2014 and the Gertler and Karadi (2015) (G&K) shocks available between 1991-2013. Both are identified with a high-frequency approach. The identifying assumption is that surprises in interest rates in a 30-minute window around FOMC announcements are dominated by the information about future monetary policy contained in the announcement, and not on other news or movements in economic and financial variables. I choose these shocks since the high-frequency identification is popular in more recent literature (see e.g., Rogers et al., 2018; Swanson, 2018; Jarociński and Karadi, 2020; Miranda-Agrippino and Ricco, 2021).

The Appendix Table A10 presents the regression output. A direct comparison to the baseline findings is not entirely correct since the sample periods differ. This aside, I find that the marginal impact of the shocks and the interaction effect is of a similar sign, but larger compared to baseline. When union density increases by one percentage point, the positive marginal impact of the shocks on Gini significantly decrease by 1.11% and 0.68% for the respective shocks. In baseline, the marginal impact decreased by 0.20% when the union density increased by one percentage point. Appendix Figure A6 shows that the positive and significant relationship between the shocks and Gini decreases up to a union density of 14% for the N&S shocks and roughly 16% for the G&K shocks, which corresponds to the baseline finding in Figure 5. There is also larger uncertainty with the N&S shocks since the interaction variable is significant at the 5% level, which is also visible in Appendix Figure A6 having wider confidence bands.

The larger size of the interaction coefficients relates to the lower variance of the high-frequency shocks compared to the Romer and Romer shocks (see Appendix Figure A3), since both the N&S and G&K shocks are correlated with the Romer and Romer shocks.[33] Another explanation may be the inclusion of the financial crisis in the sample. However, both shocks exclude the apex of the financial crisis (second half of 2008 and first half of 2009). Lastly, it may be the case that the high-frequency shocks capture exogenous movements in monetary policy to a larger extent than the Romer and Romer shocks. When identifying the Romer and Romer shocks, they assume the same reaction function over a long period, and this might not be reasonable. However, since my sample ranges further back in time, I use the Romer and Romer shocks as the main measure of monetary policy.

Previous papers highlight that the increase in income inequality can be attributed to globalization and trade (see e.g., Jaumotte and Osorio Buitron, 2015). I estimate model (1) with an additional control on total exports and the N&S shocks, since the trade data is only available from 1999 (see specification (3) in Appendix Table A10). The results are robust but more uncertain since the interaction variable is significant on the 10% level. Still, unions alter the inequality effect of monetary shocks in the same direction as in baseline. I conclude that my findings are robust to alternative measures of shocks and the inclusion of a trade control.

5.3.4 Top income shares

An alternative measure of income inequality is the income share of the top 10% and top 1%. The increase in income inequality from the 1970s is largely explained by an upsurge in top incomes. This upsurge is driven by a large increase in the share of labor income between the 1970s and 2000s, even at the very top 0.01% (Piketty and Saez, 2003, 2006; Piketty et al., 2018; El Herradi and Leroy, 2021), and by the share of capital income thereafter (Hoffmann et al., 2020). Unions may therefore influence the inequality effect of monetary shocks when inequality is measured with top income shares. The results are robust to baseline (see Appendix Table A11). A one percentage point increase in the union density results in a 0.28% and 0.31% decrease in the positive marginal impact of the shocks on Gini, which is similar to baseline. If anything, the cushioning effect on inequality from unions is stronger compared to baseline. This is also apparent in the Appendix Figure A7, where

[33] The correlation between the Romer and Romer and N&S shocks is 0.85 and between the Romer and Romer and G&K shocks is 0.46.

unions moderate the inequality effect of the shocks up to a union density of approximately 23%, which is higher than baseline. These findings prove that unions mitigate the share of income that goes to the upper part of the income distribution when hit by a contractionary shock. The stronger interaction effect on top income shares compared to Gini supports one mechanism though which labor unions mitigate income inequality, that is to limit incomes of those belonging to the upper part of the income distribution (Lee and Mas, 2012). El Herradi and Leroy (2021) find that a contractionary shock decrease the income share of the top 1% studying 12 OECD countries. Their sample runs back to the 1920s though.

5.3.5 Additional robustness

I carry out several alternative estimations to further check the robustness of the baseline results. In baseline I capture national developments over time with GDP and unemployment. To measure broader developments, I include three principal components based on a large amount of macroeconomic time series at the national level.[34] Contractionary monetary shocks still significantly increase income inequality, but the interaction effect is no longer significant (see column (2) in Appendix Table A12). The standard deviation of the principal components is higher compared to GDP and unemployment, absorbing more of the variation of the monetary shocks. This likely explains the insignificant interaction effect.

Results are robust to the inclusion of controls accounting for dependence on income inequality among states. Figure 2 in Section 2.1 illustrates that neighboring states share similar levels of income inequality. I follow Holly et al. (2010) and include an additional control variable in model (1) capturing dependence between states. First, I measure the dependence as a simple average of all neighboring states' Gini coefficient, $W_i \sum_{n=0}^{N} Gini_{n,t}$, where W for state i is 1/N and N is the total number of neighbors (see (3) in Appendix Table A12). Second, I measure the dependence with a weighted average of all neighboring states' Gini coefficient, $\sum_{n=0}^{N} W_n Gini_{n,t}$, where $W_n = \frac{GDP_n}{\sum_{n=0}^{N} GDP_n}$ is the GDP share for neighbor n of total GDP for all neighbors (see (4) in Appendix Table A12). The results are robust to either way of measuring dependence. I find that the impact of the shocks on Gini decreases by 0.12% and 0.15% for respective specification, when the union density increases by one percentage point. Also, the level of the neighbor's inequality is important, as both controls significantly increase income inequality.

[34] I choose three principal components since they explain 62% of the total variance in the sample. The fourth principal component only explains an additional 5% of the total variance. For a list of all the variables included in the principal component analysis, see Appendix Table A5.

6 Conclusion

This paper confirms that labor unions is an important moderator of the monetary policy and income inequality relationship. Unionization rates have fallen sharply in the United States since the 1970s. Previous literature finds that this erosion of union density partly explains the increase in income inequality during the same period. This paper establishes that labor unions also have implications for how shocks to monetary policy impact income inequality.

The empirical analysis confirms previous findings that contractionary shocks to monetary policy increase income inequality, and suggests that this positive relationship between monetary shocks and income inequality is weaker when union density is higher. Hence, unions mitigate the inequality effect of monetary shocks. This finding is robust to alternative measures of monetary shocks, measures of income inequality and subsample periods. The overall transmission goes via both wages and employment. Results point towards unions making the effect of monetary shocks more evenly spread out among workers, rather than mitigating the aggregate effect of the shocks on income inequality. I also find that the strength of the union impact on the monetary policy transmission on income inequality depends on the sign of the shock, where unions have a stronger influence when shocks are contractionary compared to expansionary.

My findings improve the understanding of how labor market structures may impact the effectiveness of monetary policy. A potential direction for future research would be to investigate the interaction effect of unions and monetary shocks along the income distribution, to understand which groups are driving the result.

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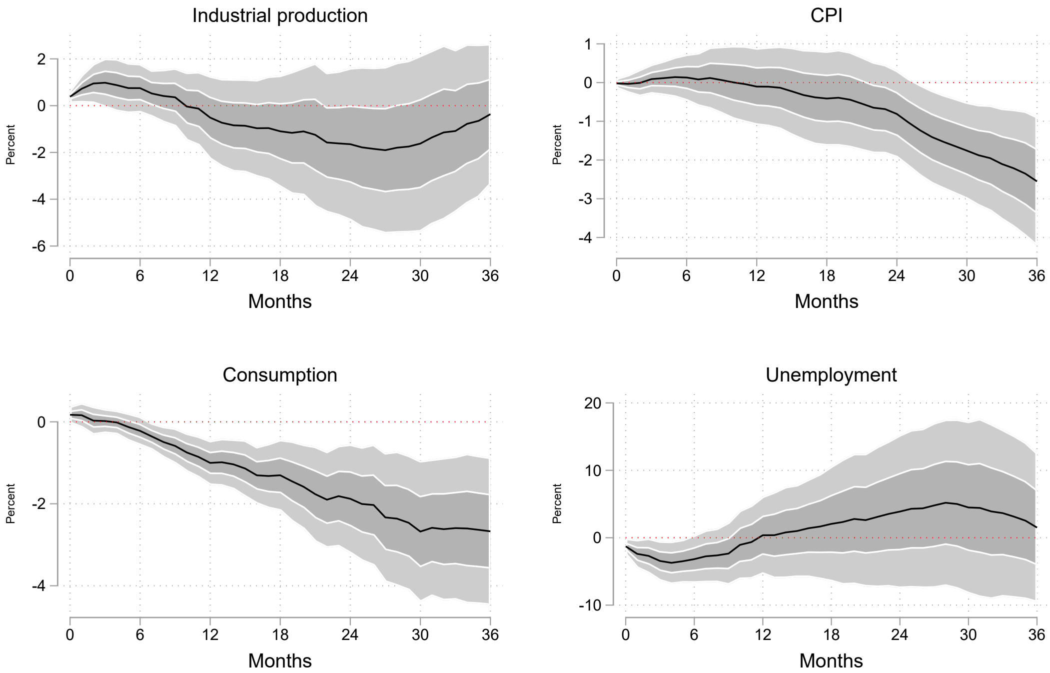
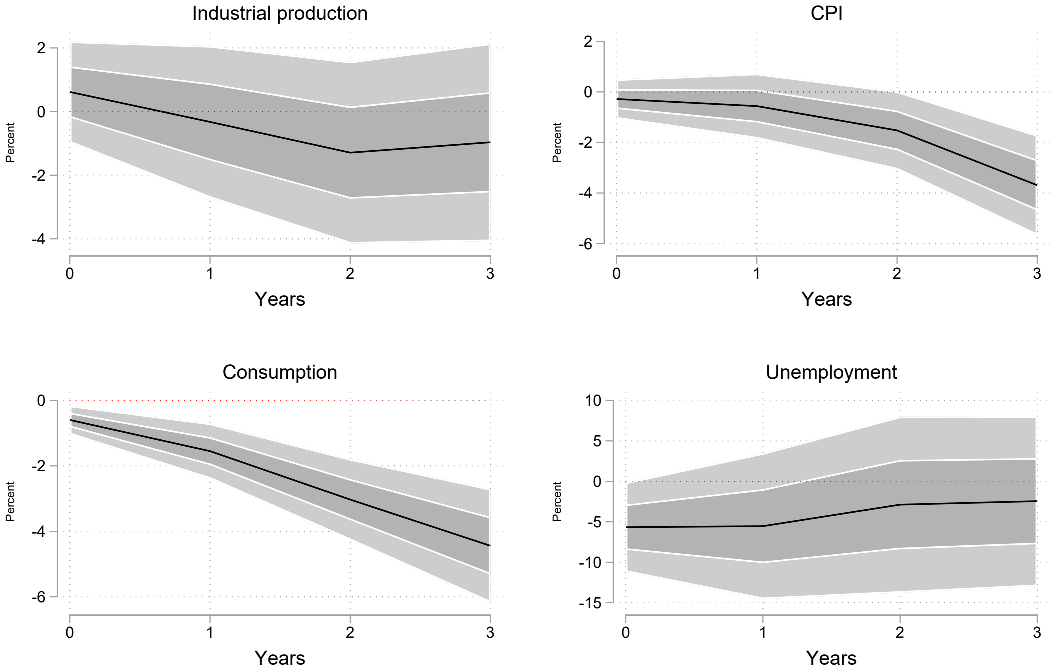
Appendix

Figure A1: Impulse responses of macro variables at a monthly frequency



Notes: The graph plots impulse responses (solid line) as well as 95% and 68% confidence bands (light and dark grey shaded areas) of industrial production (GDP not available at a monthly frequency), CPI, private consumption and unemployment in response to a one percentage point contractionary monetary shock. The sample runs between 1969:M1 and 2008:M12. I cumulate the monetary shocks to be in levels in line with Romer and Romer (2004).

Figure A2: Impulse responses of macro variables at a yearly frequency



Notes: The graph plots impulse responses (solid line) as well as 95% and 68% confidence bands (light and dark grey shaded areas) of industrial production (for comparison with a monthly frequency), CPI, private consumption and unemployment in response to a one percentage point contractionary monetary shock. The sample runs between 1969 and 2008. I cumulate the monetary shocks to be in levels in line with Romer and Romer (2004). The data is from FRED.

Table A1: Regression output for state-level employment at different frequencies

	Frequency of data		
	(1) Monthly	(2) Quarterly	(3) Annual
$MPshock_t$	0.0005***	0.0034***	0.0164***
	(0.0000)	(0.0003)	(0.0050)
Constant	0.0013***	0.0038***	0.0155
	(0.0000)	(0.0000)	(0.0000)
Observations	19750	6550	1600
Number of states	50	50	50
Adjusted R-square	0.0047	0.0117	0.0034
State FE	YES	YES	YES

Note: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The dependent variable is the log change in total employment. The sample consist of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. The data on employment is from the US Bureau of Labor Statistics.

Table A2: Test for structural break

Dependent variable:	Gini	Union density	
Constant	0.534***	0.169***	
	(0.008)	(0.007)	
	Unknown break date	Unknown break date	Known break date
Estimated break date	1987	1984	1987
Wald test	228.362	202.401	138.567
	(0.000)	(0.000)	(0.000)
LR-test	77.891	73.789	61.444
	(0.000)	(0.000)	(0.000)

Note: Levels of significance: p<0.01, ** p<0.05, * p<0.1. Standard errors are reported in parentheses for the regression output. P-values are reported in parentheses for the Wald and LR statistics. The null hypothesis is no structural break.

Table A3: Stationarity tests when including a dummy starting in 1987

Dependent variable:	Gini	Union density
Dummy1987	0.093***	-0.081***
	(0.001)	(0.001)
Constant	0.483***	0.213***
	(0.003)	(0.009)
Observations	2000	2000
Levin-Lin-Chu	0.000	0.000
Harris-Tzavalis	0.000	0.000
Breitung	0.000	0.000
Im-Pesaran-Shin	0.000	0.000
ADF	0.000	0.000
Number of states	50	50

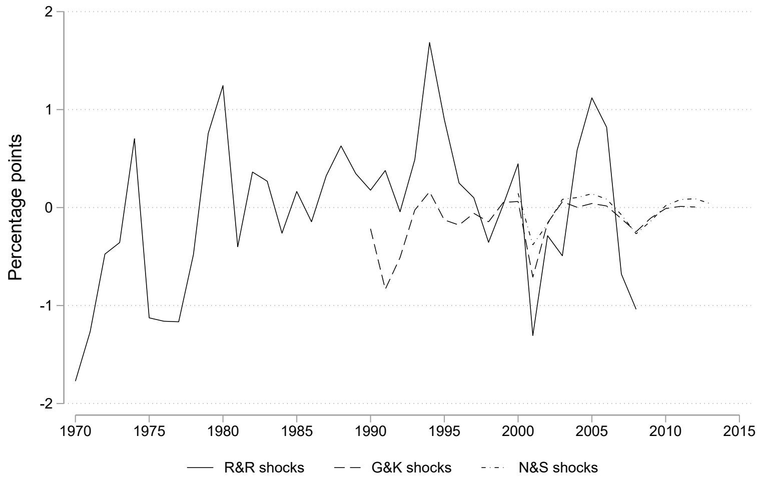
Note: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The null hypothesis in Levin-Lin-Chu is that all panels contain unit roots, while the alternative hypothesis is that all panels are stationary. This is different to the other tests who has an alternative hypothesis of at least one panel is stationary. Therefore, the Levin-Lin-Chu test is the preferred here. P-values are reported for all unit root tests. Standard errors are reported in parentheses for the regression output. The ADF-test includes one lag.

Table A4: Variable description and descriptive statistics

	Obs	Mean	sd	Min	Max	Description	Source	Comment
Gini	1950	0.535	0.056	0.410	0.709	Gini index on total income	Frank (2009)	
Top 10% income share	1950	0.378	0.056	0.204	0.622	Income share on total income	Frank (2009)	
Top 1% income share	1950	0.129	0.046	0.040	0.361	Income share on total income	Frank (2009)	
Monetary shocks, R&R	1950	-0 .003	0.077	-0.177	0.168	See section 4.2	Romer and Romer (2004)	
Monetary shocks, HF (N&S)	450	-0.006	0.014	-0.024	0.026	See section 4.2	Nakamura and Steinsson (2018)	
Monetary shocks, HF (G&K)	950	-0.015	0.026	-0.083	0.016	See section 4.2	Gertler and Karadi (2015)	
Union member density	1950	0.167	0.080	0.023	0.424	See section 4.2	Hirsch et al. (2001)	
△ Wages, real	1950	0.004	0.019	-0.130	0.199	Average real wage and salary income per worker	BEA	Deflated with national CPI
△ Employment	1950	0.017	0.023	-0.068	0.184	Total employment (number of jobs) for wage and salary workers	BEA	
△ Taxes	1950	0.077	0.077	-0.585	1.082	Total state government tax revenue	BLS	
△ Expenditure of GDP	1950	0.001	0.007	-0.068	0.071	Total state government expenditure	US Cencus Bureau	
△ House price index	1950	0.056	0.064	-0.419	0.442	Price index on single family houses	Fed Housing Agency and Statistical abstract	See notes below
△ College	1950	0.030	0.060	-0.598	0.415	# College graduates/Total state population	Frank (2009)	
△ Population	1950	0.011	0.011	-0.062	0.083	Total state population	BEA	
△ GDP	1950	0.030	0.018	-0.018	0.070		BEA	National level control
△ Unemployment	1950	0	0.009	-0.021	0.029		FRED	National level control
△ Exports	450	0.085	0.142	-0.532	0.919	Total exports by state	US Cencus Bureau	
Weighted Gini, average	1950	0.514	0.117	0	0.674		Frank (2009)	
Weighted Gini, GDP	1950	0.516	0.118	0	0.670		Frank (2009)	
△ Minimum wage	1700	0.035	0.052	0	0.342		Vaghul and Zipperer (2016)	State level minimum wage
△ GDP manufacturing, real	1950	0.018	0.083	-0.402	0.824		BEA	Deflated with GDP deflator
△ GDP agriculture, real	1950	0.004	0.204	-1.192	1.057		BEA	Deflated with GDP deflator
△ EF index	1350	0.007	0.044	-0.252	0.244		Stansel et al. (2020)	
PC1	1950	0	0.052	-0.068	0.154	See Appendix Table A5		
PC2	1950	0	0.039	-0.089	0.070	See Appendix Table A5		
PC3	1950	0	0.029	-0.050	0.071	See Appendix Table A5		

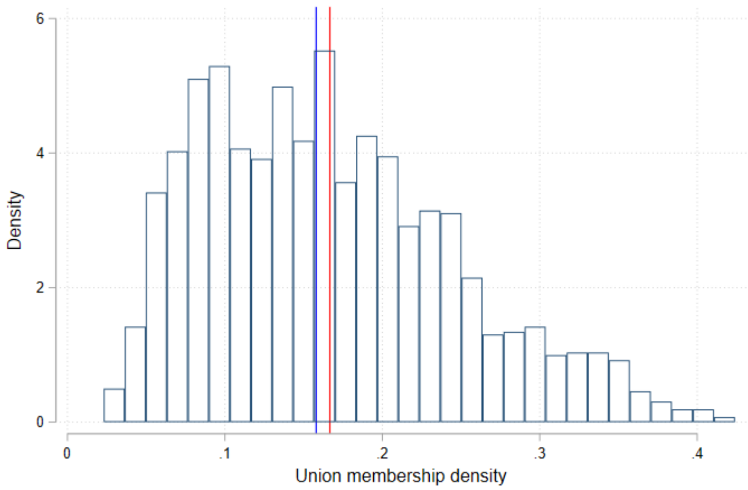
Note: △ means that the variables are log changes, i.e. in growth rates. EF stands for economic freedom. PC stands for principal component. All deflated variables have base year 2010. BEA is the U.S. Bureau of Economic Analysis, BLS is the U.S. Bureau of Labor Statistics, and FRED is the FED St Louis database. About the house price indecx, HPI: Data is only available from 1975 so I therefore inserted data from the Statistical Abstract on housing units and valuation from 1969-1975, and merged the two sets.

Figure A3: Comparison of monetary policy shocks



Notes: Comparison of three sets of shocks: the Romer and Romer (2004) shocks between 1970-2008, the Gertler and Karadi (2015) shocks between 1991-2013, and the Nakamura and Steinsson (2018) shocks between 2001-2014. Monetary policy is contractionary when shocks are positive and expansionary when shocks are negative.

Figure A4: Histogram of union density



Notes: The sample consists of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. The blue line shows the median which is 15.8% and the red line shows the mean which is 16.7%.

Table A5: Variables included in principal components

Description of variables	
Changes in inventories, nominal, % GDP	Labour productivity of the total economy
Commercial bank assets, loans and leases in bank credit	Monetary aggregate M1
Consumer confidence indicator	Monetary aggregate M3
Corporate profits, after tax	Monetary base
CPI, energy	Net lending or net borrowing, % GDP
CPI, all items	Orders for manufacturing goods, real
Core inflation	Owner-occupied residential structures
Commodity prices, crude oil	Passenger car registrations
Credit to corporations from all Sectors, % GDP	PCE, real
Credit to general government from all sectors, % GDP	PCE, price index
Credit to households from all sectors, % GDP	PCE, core, real
Current account balance, % GDP	PCE, durable goods, real
Domestic investment, government, net, % GDP	PCE, durable goods, price index
Domestic investment, private, net, % GDP	PCE, new autos, real
Exchange rate, local currency per US $, Canada	PCE, new autos, price index
Exchange rate, local currency per US $, Germany	PCE, non-durable goods, real
Exchange rate, local currency per US $, Japan	PCE, non-durable goods, price index
Exchange rate, local currency per US $, Switzerland	PCE, services, real
Exchange rate, local currency per US $, UK	PCE, services, price index
Exports of goods and services, real	Personal saving, % of dispposable income
Exports of goods and services, price index	PPI consumer finished goods
Fixed investment, non-residential, gross, real	PPI durable consumer goods
Fixed investment, non-residential, price index	PPI energy
Fixed investment, residential, gross, real	PPI manufacturing goods
Fixed investment, residential, price index	PPI non-durable consumer goods
GDP, real	PPI primary products
GDP, price index	Private inventories, real
Government consumption expeditures and	Production, total construction
gross investment, federal government, real	Production, total industry excluding construction
Government consumption expeditures and	Property prices, commercial
gross investment, federal government, price index	Property prices, residential: new one-family houses
Government consumption expeditures and	Sales of new one family houses
gross investment, state and local government, real	Share prices: NY stock exchange composite
Government consumption expeditures and	Terms of trade: goods and services
gross investment, state and local government, price index	Total factor productivity
Government saving, gross, % GDP	Total retail trade, real
Household financial assets, % of nominal disposable income	Total wholesale trade, real
Household non-financial assets, % of nominal disposable income	Unit labor cost
Housing output	Unit labour cost, manufacturing goods
Housing started	
Imports of goods and services, real	
Imports of goods and services, price index	
Labour - capital substitution	
Labour force (ages 15-64)	

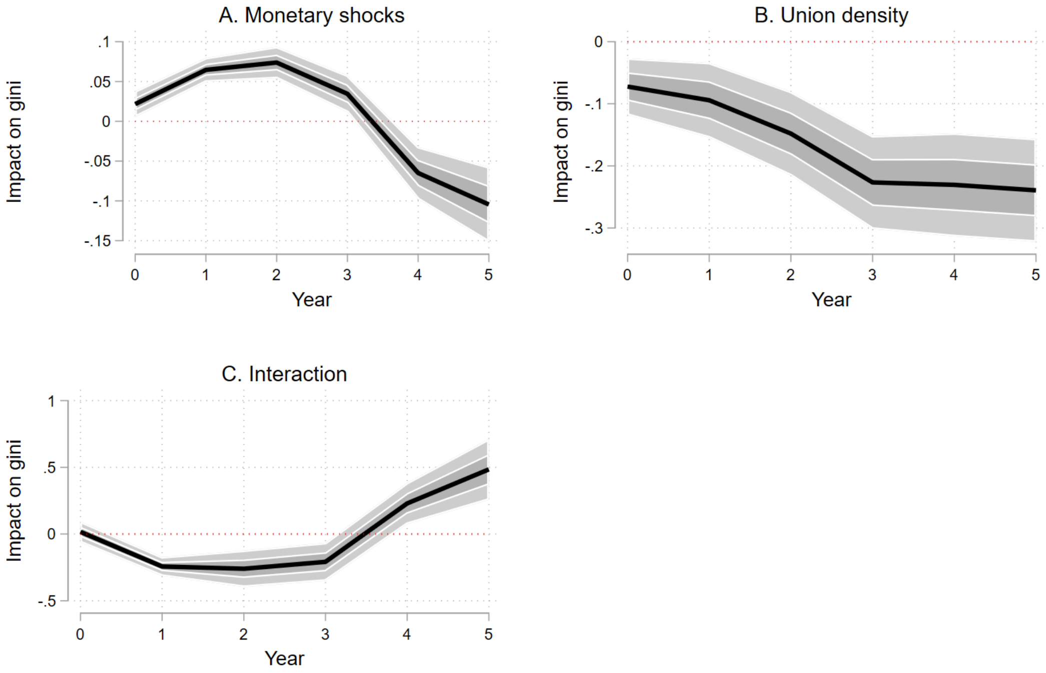
Note: PCE stands for personal consumption expenditures. PPI stands for producer price index. All variables are in growth rates.

Table A6: Baseline regression with contemporaneous impact

Dependent variable: Gini	
$\text{Gini}_{i,t-1}$	0.889***
	(0.011)
$MPshock_t$	0.052***
	(0.009)
$Union_{i,t}$	-0.051***
	(0.012)
$MPshock_t * Union_{i,t}$	-0.044
	(0.047)
$\text{Taxes}_{i,t}$	-0.006
	(0.004)
Expenditures of $\text{GDP}_{i,t}$	-0.093**
	(0.037)
$\text{HPI}_{i,t}$	-0.002
	(0.005)
$\text{College}_{i,t}$	0.007
	(0.004)
$\text{Population}_{i,t}$	-0.126***
	(0.037)
GDP_t	0.141***
	(0.032)
Unemployment_t	0.258***
	(0.067)
Observations	1900
Number of states	50
State FE	YES
Dummy 1987	YES
Adjusted R-square	0.953
# Bootstraps	1000
AR(1) p-value	0.000
AR(2) p-value	0.0450

Notes: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The sample consists of all US states (except the District of Columbia due to data limitations) for the period 1970-2008. The dependent variable is the Gini index on total income. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS (LSDVC) on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

Figure A5: Response of Gini in a local projection analysis



Notes: The graph plots impulse responses (solid line) as well as 95% and 68% confidence bands (light and dark grey shaded areas) of Gini in response to a one percentage point contractionary monetary shock (A), a one percentage point increase in union density (B), and a one percentage point contractionary monetary shock given the average union density, i.e. the coefficient for the interaction variable (C).

Table A7: Regression output for wage and employment regressions, contemporaneous impact

	Dependent variable: Gini	
	(1) Average real wages	(2) Employment
Dep. variable$_{i,t-1}$	0.296***	0.312***
	(0.020)	(0.015)
MP shocks$_t$	0.007	-0.016*
	(0.011)	(0.009)
Union$_{i,t}$	-0.088***	0.048***
	(0.012)	(0.010)
$MPshock_t * Union_{i,t}$	-0.275***	-0.007
	(0.058)	(0.048)
Taxes$_{i,t}$	0.035***	0.033***
	(0.005)	(0.004)
Expenditures of GDP$_{i,t}$	-0.344***	-0.544***
	(0.047)	(0.038)
HPI$_{i,t}$	-0.026***	0.028***
	(0.006)	(0.005)
College$_{i,t}$	-0.005	0.003
	(0.005)	(0.004)
Population$_{i,t}$	0.416***	0.621***
	(0.047)	(0.038)
GDP$_t$	0.654***	0.184***
	(0.040)	(0.033)
Unemployment$_t$	1.033***	-0.856***
	(0.085)	(0.070)
Observations	1900	1900
Number of states	50	50
State FE	YES	YES
Adjusted R-square	0.304	0.708
# Bootstraps	100	100
AR(1) p-value	0.000	0.000
AR(2) p-value	0.000	0.020

Note: Levels of significance: p<0.01, *** p<0.05, ** p<0.1 *. The dependent variables are the log change in average real wages (1) and total employment (2). The sample consist of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

Table A8: Regression output for wage and employment regressions, two lags

	Dependent variables	
	(1) Average real wages	(2) Employment
Dep. variable$_{i,t-1}$	0.596***	0.674***
	(0.020)	(0.018)
Dep. variable$_{i,t-2}$	-0.123***	-0.197***
	(0.027)	(0.034)
$MPshock_{t-1}$	0.052***	0.039***
	(0.013)	(0.014)
$MPshock_{t-2}$	0.017	-0.003
	(0.013)	(0.014)
$Union_{i,t-1}$	-0.082***	-0.054**
	(0.025)	(0.027)
$Union_{i,t-2}$	0.003	0.065**
	(0.026)	(0.027)
$MPshock_{t-1} * Union_{i,t-1}$	-0.342***	-0.191**
	(0.073)	(0.078)
$MPshock_{t-1} * Union_{i,t-2}$	0.091	-0.172**
	(0.065)	(0.068)
Taxes$_{i,t-1}$	-0.024***	-0.001
	(0.006)	(0.006)
Taxes$_{i,t-2}$	-0.014**	0.003
	(0.006)	(0.006)
Expenditures of GDP$_{i,t-1}$	-0.112**	-0.223***
	(0.054)	(0.058)
Expenditures of GDP$_{i,t-2}$	0.037	0.021
	(0.058)	(0.063)
HPI$_{i,t-1}$	-0.004	0.006
	(0.007)	(0.008)
HPI$_{i,t-2}$	-0.008	-0.014**
	(0.007)	(0.007)
College$_{i,t-1}$	0.002	-0.005
	(0.006)	(0.007)
College$_{i,t-2}$	0.017***	0.013*
	(0.006)	(0.007)
Population$_{i,t-1}$	0.454***	0.080
	(0.075)	(0.087)
Population$_{i,t-2}$	-0.395***	-0.121
	(0.074)	(0.078)
GDP$_{t-1}$	-0.042	0.389***
	(0.051)	(0.053)
GDP$_{t-2}$	0.191***	-0.303***
	(0.056)	(0.059)
Unemployment$_{t-1}$	0.573***	0.993***
	(0.114)	(0.121)
Unemployment$_{t-2}$	0.244**	-0.383***
	(0.106)	(0.116)
Observations	1800	1800
Number of states	50	50
Adjusted R-square	0.333	0.447
# Bootstraps	100	100
AR(1) p-value	0.000	0.000
AR(2) p-value	0.022	0.000
AR(3) p-value	0.360	0.398
State FE	YES	YES
Dummy 1987	YES	YES

Note: Levels of significance: p<0.01, *** p<0.05, ** p<0.1 *. The dependent variables are the growth rate of average real wages per worker and total employment. The sample consist of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. The AR(1), AR(2) and AR(3) p-values are Arellano-Bond tests for autocorrelation.

Table A9: Regression output accounting for union density covariates

	Dependent variable: Gini			
	(1) Manufacturing interaction	(2) Agricultural interaction	(3) EF index interaction, 1982-2008	(4) Minimum wage control, 1976-2008
$\text{Gini}_{i,t-1}$	0.865***	0.861***	0.881***	0.882***
	(0.011)	(0.011)	(0.014)	(0.013)
MP shocks_{t-1}	0.038***	0.037***	0.019	0.034***
	(0.009)	(0.009)	(0.012)	(0.010)
$\text{Union}_{i,t-1}$	-0.057***	-0.055***	-0.055***	-0.033**
	(0.011)	(0.011)	(0.020)	(0.014)
$MPshock_{t-1} * Union_{i,t-1}$	-0.195***	-0.198***	0.010	-0.112**
	(0.045)	(0.043)	(0.085)	(0.057)
$\text{Manufacturing}_{i,t-1}$	0.005			
	(0.004)			
MP shock_{t-1}*$\text{Manufacturing}_{i,t-1}$	0.009			
	(0.040)			
$\text{Agriculture}_{i,t-1}$		-0.002*		
		(0.001)		
MP shock_{t-1}*$\text{Agriculture}_{i,t-1}$		-0.103***		
		(0.020)		
EF $\text{index}_{i,t-1}$			-0.024***	
			(0.009)	
MP shock_{t-1}*EF $\text{index}_{i,t-1}$			-0.039	
			(0.138)	
$\text{Taxes}_{i,t-1}$	-0.004	-0.004	-0.015***	-0.003
	(0.004)	(0.004)	(0.005)	(0.004)
Expenditures of $\text{GDP}_{i,t-1}$	0.005	-0.019	0.052	0.019
	(0.040)	(0.039)	(0.048)	(0.041)
$\text{HPI}_{i,t-1}$	0.003	0.004	0.006	-0.007
	(0.005)	(0.005)	(0.007)	(0.006)
$\text{College}_{i,t-1}$	0.004	0.004	0.010**	0.008*
	(0.005)	(0.005)	(0.005)	(0.005)
$\text{Population}_{i,t-1}$	-0.096***	-0.091**	-0.114**	-0.092**
	(0.036)	(0.036)	(0.048)	(0.043)
GDP_{t-1}	-0.192***	-0.164***	-0.291***	-0.248***
	(0.031)	(0.030)	(0.038)	(0.038)
$\text{Unemployment}_{t-1}$	-0.494***	-0.467***	0.762***	-0.603***
	(0.066)	(0.066)	(0.086)	(0.086)
Minimum $\text{wage}_{i,t-1}$				-0.002
				(0.006)
Observations	1900	1900	1300	1700
Number of states	50	50	50	50
State FE	YES	YES	YES	YES
Dummy 1987	YES	YES	YES	YES
Adjusted R-square	0.950	0.950	0.929	0.953
AR(1) p-value	0.000	0.000	0.000	0.000
AR(2) p-value	0.430	0.466	0.065	0.747

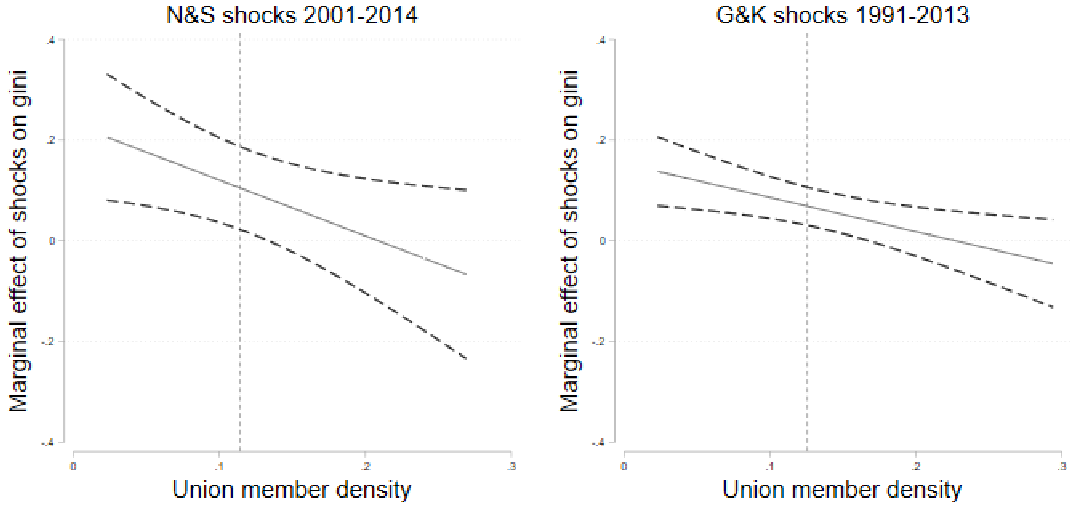
Notes: Levels of significance: *** p<0.01, ** p<0.05, * p<0.1. The sample consists of all US states (except the District of Columbia due to data limitations) for the period 1970-2008. The dependent variable is the Gini index on total income. Manufacturing and agriculture are measured as shares of GDP. EF index stands for economic freedom index. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS (LSDVC) on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

Table A10: Regression output for alternative measures of monetary shocks

	Dependent variable: Gini		
	(1) N&S shocks 2001-2014	(2) G&K shocks 1991-2013	(3) N&S shocks 2001-2014, trade control
$\text{Gini}_{i,t-1}$	0.862***	0.882***	0.908***
	(0.035)	(0.023)	(0.033)
$MPshock_{t-1}$	0.231***	0.153***	0.279***
	(0.073)	(0.040)	(0.087)
$Union_{i,t-1}$	-0.005	-0.066**	0.009
	(0.049)	(0.028)	(0.049)
$MPshock_{t-1} * Union_{i,t-1}$	-1.108**	-0.675***	-1.226*
	(0.498)	(0.256)	(0.645)
$\text{Taxes}_{i,t-1}$	0.001	-0.002	0.001
	(0.006)	(0.005)	(0.007)
Expenditures of $\text{GDP}_{i,t-1}$	-0.078	0.024	-0.055
	(0.065)	(0.056)	(0.066)
$\text{HPI}_{i,t-1}$	-0.032***	-0.003	-0.031***
	(0.010)	(0.009)	(0.010)
$\text{College}_{i,t-1}$	-0.001	-0.001	-0.000
	(0.007)	(0.006)	(0.007)
$\text{Population}_{i,t-1}$	-0.122	-0.139*	-0.094
	(0.099)	(0.081)	(0.102)
GDP_{t-1}	0.077	-0.163***	0.088
	(0.070)	(0.042)	(0.066)
$\text{Unemployment}_{t-1}$	-0.051	-0.371***	-0.056
	(0.094)	(0.074)	(0.094)
$\text{Exports}_{i,t-1}$			0.004
			(0.003)
Observations	650	1100	650
Number of states	50	50	50
State FE	YES	YES	YES
Adjusted R-square	0.870	0.871	0.871
AR(1) p-value	0.000	0.000	0.000
AR(2) p-value	0.002	0.051	0.000

Note: Levels of significance: p<0.01, ** p<0.05, * p<0.1. The sample consists of all US states (except the District of Columbia due to data limitations) and the dependent variable is the Gini index on total income. I report bootstrap robust standard errors in parentheses. In (1) I use the Nakamura and Steinsson (2018) shocks ranging between 2001-2014. In (2) I use the Gertler and Karadi (2015) shocks ranging between 1991-2013. In (3) I use the N&S shocks but add a control for state-level exports (only available for this time period). I base the initialization for the bias corrected OLS on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

Figure A6: The impact of alternative monetary shocks on Gini for varying union densities



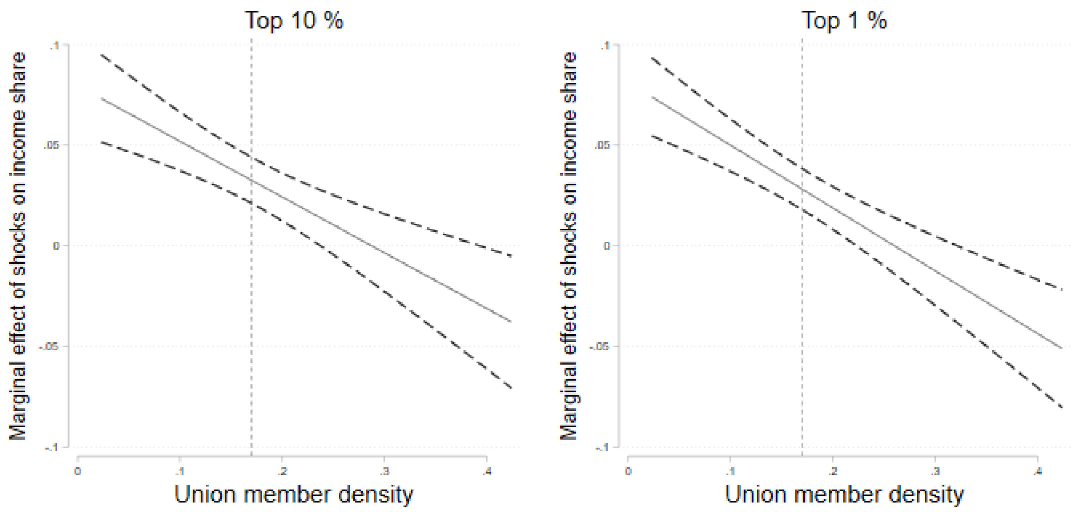
Notes: Dashed lines give 95% confidence intervals. The vertical dashed line shows the mean density of union membership, which is equal to 11% with the N&S shocks and 13% with the G&K shocks. The results are based on the coefficients in columns (1) and (2) in Appendix Table A10.

Table A11: Regression output using top income shares as dependent variables

	Dependent variables	
	(1) Top 10%	(2) Top 1%
Dependent variable$_{i,t-1}$	0.882***	0.889***
	(0.014)	(0.014)
$MPshock_{t-1}$	0.080***	0.081***
	(0.012)	(0.011)
$Union_{t-1}$	-0.110***	-0.123***
	(0.016)	(0.014)
$MPshock_{t-1} * Union_{i,t-1}$	-0.277***	-0.312***
	(0.063)	(0.056)
Taxes$_{i,t-1}$	-0.001	-0.006**
	(0.005)	(0.005)
Expenditures of GDP$_{i,t-1}$	-0.175***	-0.201***
	(0.055)	(0.049)
HPI$_{i,t-1}$	-0.003	-0.010*
	(0.007)	(0.006)
College$_{i,t-1}$	0.005	0.009
	(0.007)	(0.006)
Population$_{i,t-1}$	0.006	0.036
	(0.532)	(0.047)
GDP$_{t-1}$	-0.046	0.071*
	(0.044)	(0.039)
Unemployment$_{t-1}$	-0.052	0.125
	(0.095)	(0.085)
Observations	1900	1900
Number of states	50	50
State FE	YES	YES
Dummy 1987	YES	YES
Adjusted R-square	0.899	0.866
AR(1) p-value	0.000	0.000
AR(2) p-value	0.002	0.000

Note: Levels of significance: p<0.01, *** p<0.05, ** p<0.1 *. The sample consists of all US states (except the District of Columbia due to data limitations) for the years 1970 to 2008. The dependent variables are the top 10% and top 1% income shares. I report bootstrap robust standard errors in parentheses. I base the initialization for the bias corrected OLS on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.

Figure A7: The impact of monetary shocks on top income shares for varying union densities



Notes: Dashed lines give 95% confidence intervals. The horizontal dashed line shows the mean density of union membership, which is equal to 17%. The results are based on the coefficients in Table A11.

Table A12: Additional robustness regression output

	Dependent variable: Gini		
	(1) PC control	(2) Average dependence control	(3) Weighted dependence control
$\text{Gini}_{i,t-1}$	0.865***	0.726***	0.751***
	(0.0011)	(0.014)	(0.014)
$MPshock_{t-1}$	0.044***	0.022***	0.028***
	(0.009)	(0.009)	(0.009)
$Union_{i,t-1}$	-0.029**	-0.015	-0.022*
	(0.012)	(0.011)	(0.012)
$MPshock_{t-1} * Union_{i,t-1}$	-0.070	-0.121***	-0.150***
	(0.044)	(0.043)	(0.043)
$\text{Taxes}_{i,t-1}$	0.001	-0.008**	-0.007*
	(0.004)	(0.004)	(0.004)
Expenditures of $\text{GDP}_{i,t-1}$	-0.039	-0.012	-0.008
	(0.039)	(0.038)	(0.038)
$\text{HPI}_{i,t-1}$	-0.001	0.008*	0.008*
	(0.005)	(0.005)	(0.005)
$\text{College}_{i,t-1}$	0.002	0.005	0.005
	(0.005)	(0.005)	(0.005)
$\text{Population}_{i,t-1}$		-0.147***	-0.157***
		(0.036)	(0.036)
GDP_{t-1}		-0.127***	-0.155***
		(0.030)	(0.0230)
$\text{Unemployment}_{t-1}$		-0.353***	-0.408***
		(0.065)	(0.066)
PC1_{t-1}	-0.035***		
	(0.008)		
PC2_{t-1}	-0.037***		
	(0.012)		
PC3_{t-1}	-0.058***		
	(0.014)		
Weighted Gini, $\text{average}_{i,t-1}$		0.245***	
		(0.017)	
Weighted Gini, $\text{gdp}_{i,t-1}$			0.204***
			(0.017)
Observations	1900	1900	1900
Number of states	50	50	50
State FE	YES	YES	YES
Dummy 1987	YES	YES	YES
Adjusted R-square	0.954	0.761	0.808
AR(1) p-value	0.000	0.000	0.000
AR(2) p-value	0.011	0.044	0.067

Note: Levels of significance: p<0.01, *** p<0.05, ** p<0.1 *. The dependent variable is the Gini index on total income. The sample consist of all US states (except the District of Columbia due to data limitations) from 1970-2008. In (1), I control for three principal components on the national level instead of GDP and unemployment. In (2) I include a control for each state's neighbor's average income inequality. In (3) I include a control for each state's weighted average income inequality, weighted with the share of GDP. I report bootstrap robust standard errors (in parentheses). I base the initialization for the bias corrected OLS on the Arellano-Bond (AB) difference GMM estimator with no constant (which is pre-programmed in the command xtlsdvc). The bootstrap variance-covariance matrix uses 1000 repetitions. The AR(1) and AR(2) p-values are Arellano-Bond tests for autocorrelation.