Society for the Study of Economic Inequality

Working Paper Series

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ECINEQ WP 2020 - 519

ECINEQ 2020 - 519
January 2020

www.ecineq.org

Monetary policy and the top one percent: Evidence from a century of modern economic history [‡]

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Abstract

While a growing line of research has assessed the distributional consequences of monetary policy, most of these studies rely on survey-based estimates of inequality and feature a shorter time coverage. This paper examines the distributional implications of monetary policy on top income shares in 12 advanced economies between 1920 and 2015. We exploit the implications of the macroeconomic policy trilemma with an external instrument approach to identify exogenous variations in monetary conditions. The obtained results indicate that contractionary monetary policy strongly decreases the share of national income held by the top one percent and vice versa, irrespective of the state of the economy. Our findings also suggest that the effect of monetary tightening on top income shares is likely to be channeled via lower asset prices.

Keywords: Monetary policy, Top income shares, Macroeconomic Policy Trilemma, External Instrument.

JEL Classification: E25, E42, E52.

[‡]We thank Jakob De Haan, JŽr™me HŽricourt, Etienne Farvaque, Paul Hubert, Fabien Tripier, Anna Samarina and Gabriele Galati for helpful comments on various drafts of this paper. We would also like to thank seminar participants at De Nederlandsche Bank, 4eme Journées Internationales de MacroŽconomie et de Finance, 19th Doctoral Meetings in International Trade and International Finance, 36th Symposium in Money Banking and Finance, 25th Dubrovnik Economic Conference, 68th Annual Meeting of the French Economic Association and the 8th Meeting of the Society for the Study of Economic Inequality.

1 Introduction

The potential distributional effects of monetary policy have recently become an active topic in the inequality debate as a consequence of the unconventional measures that central banks implemented following the financial crisis (see Colciago et al. (2019) for a complete survey). This was unusual because it is widely accepted that central banks should not be concerned about inequality: they are independent of the political process, and dealing with distributional matters goes beyond their mandate. Nevertheless, the combination of an ultra-low interest rate environment and large asset purchase programs is suspected to have reduced modest household savings and driven up asset prices. Meanwhile, central bankers such as Draghi (2016) or Bernanke (2015) strongly believe that their non-standard monetary policies had modest distributional implications. They argue instead that the post-crisis monetary policy toolkit allowed for the restoration of growth and increased employment levels, which primarily favored low-income households. In the spirit of Coibion et al. (2017), this debate underlines that the effect of monetary policy on income inequality would be channeled through: (i) households' income composition (some rely primarily on labor income, while others may receive other forms of revenue as rents or dividends), (ii) heterogeneous effects of business cycle fluctuations with respect to earnings (modest and low-skilled workers are generally the most exposed to unexpected shocks), and (iii) the distribution of assets and liabilities between households (financial assets are held primarily by rich households, which could be the first to benefit from higher asset prices).

In this context, macroeconomic research has been increasingly devoted to analyzing the collateral effects of monetary policy on income and wealth distributions. As far as theoretical contributions are concerned, they have mainly built on DSGE models with incomplete markets and heterogeneous agents. Dolado et al. (2018), for instance, emphasized two specific channels: (i) top-income households happen to be high-skilled and experience increasing wages as a consequence of a monetary expansion because they benefit from lower matching frictions in labor markets; (ii) as these individuals present complementary features to capital, an increase in the demand for the latter only magnifies income inequality in comparison to poor, low-skilled workers.

On the empirical side, the literature is still ambiguous on the effects of monetary policy on income distribution. Numerous country-level studies suggest that conventional monetary tightening increases income inequality (see Coibion et al. (2017) for the U.S., Mumtaz and Theophilopoulou (2017) for the U.K. and Furceri et al. (2018) for a selection of advanced and emerging economies). At the same time, some studies argue that expansionary monetary policy also increases income inequality (see Cloyne et al. (2018) for the U.K. and U.S., Inui et al. (2017) for Japan) In contrast, recent research on the distributional effects of unconventional monetary policy mostly shows that the relationship between monetary expansions and inequality is negative, but small in magnitude (see, e.g., Casiraghi et al. (2018), Guerello (2018) and Bivens (2015)).

The existing empirical literature on the relationship between monetary policy and income distribution focuses primarily on survey-based estimates of inequality. This approach can be problematic because, unlike tax data, these estimates produce lower inequality levels and account less for the recent increasing trend in top income shares (see, e.g., Burkhauser et al. (2012), Wolff and Zacharias (2009)). As recalled by Atkinson et al. (2011), this discrepancy stems mainly from the underrepresentation of top-income households and a lower coverage of capital gains. In this paper, we use tax-based estimates of top income shares, which more broadly cover business and capital incomes (Yonzan et al. (2018)).

Another drawback of the literature concerns the use of inequality measures over a short period of time, which implies giving coverage to fewer exceptional macroeconomic events. In this respect, using historical data allows to cover important events experienced in the developed world, such as the Great Depression and the post-war boom, hence giving more variation in the data and in particular, top income shares. However, analyzing the distributional effects of monetary policy from a historical perspective poses the challenge of identifying exogenous variations in monetary conditions. Dealing with this point is particularly demanding because the conduct of monetary policy in advanced economies experienced several changes throughout the 20th century. Such shifts relate to the succession of different exchange rate regimes, the occurrence of many banking crises and the usage of multiple frameworks in monetary policy decisions (e.g., inflation targeting, the Taylor rule, etc.). Our paper aims to address these challenges using a different setting and a novel identification approach.

This paper analyzes the relationship between monetary policy and top income shares between 1920 and 2015 using annual data across 12 advanced economies: Australia, Canada, Italy, Germany, Denmark, France, the U.K, Japan, the Netherlands, Norway, Sweden and the U.S. It was possible to conduct this historical analysis thanks to the combination of two datasets. We mobilize the World Inequality Database (WID), which offers open access to historical series of income and wealth inequality. In particular, we focus on the right tail of the income distribution and use the share of the national income held by the richest one percent as the main top income indicator. As recalled by Roine and Waldenström (2015) and Roine et al. (2009), the richest one percent receive a significant share of their total income in the form of dividends and capital gains while being almost untraceable in household income and wealth surveys. As an alternative top income measure, we use in our analysis the top 10 percent share of the national income, which is believed to capture well-off households with heterogeneous income sources. Long series of macroeconomic variables are extracted from the Jordà-Schularick-Taylor Macrohistory Database, developed by Jordà et al. (2016). Using such data is of great interest because they offer a rich set of control variables that could enter as potential determinants of top incomes.

Our empirical methodology primarily relies on Local Projections (LPs) à la Jordà (2005). The latter generates dynamic responses of top income shares to an exogenous change in the short-term interest rate. The identification of such shocks is based on a quasi-natural experiment approach as recently proposed by Jordà et al. (2019). This approach responds to the fact that the short-term interest rate and top income shares are potentially influenced by common unobserved factors, biasing the empirical effect of an interest rate change on top income shares. Specifically, our approach uses an instrumental variable in the context of local projections (LP-IV) (see, Jordà et al. (2015); Ramey and Zubairy (2018)) to isolate exogenous fluctuations of the short-term interest rate, which are drawn from the well-known macroeconomic policy trilemma. The trilemma states that movements in the base country's short-term interest rate provide exogenous variations in the domestic short-term rate for an open PEG. As a result, policy choices regarding capital mobility, exchange rates and interest rates provide a natural experiment to analyze the effect of monetary policy on top income shares.

In robustness, we follow the work Mertens and Ravn (2014), Gertler and Karadi (2015) and Stock and Watson (2018), by estimating a panel-SVAR-IV with 6 variables, including our top income indicator. As noted by Barnichon and Brownlees (2018), the panel VAR model presents complementary features besides LPs, as it yields consistent results when it is correctly specified; in contrast, LPs are admittedly less efficient, but they remain robust to model misspecification. Finally, because LPs easily accommodate non-linearities, we test our model in a state-dependent setting, where we allow the response of top income shares to depend on the regime of a specific variable (i.e., business cycle, the inflation regime, credit cycles and monetary policy stance).

Our evidence suggests that monetary policy has a significant and persistent impact on income inequality via top income shares. Monetary tightening decreases the share of national income held by the top one percent, while expansionary monetary policy has the opposite effect. A positive shock to the domestic short-term interest rate via the external instrument reduces the top one percent's income share, with an accumulated decrease of 0.35 percentage points. We demonstrate that this effect is arguably driven by lower stock prices, which is consistent with the income composition channel of Coibion et al. (2017) and the indirect income channel of Ampudia et al. (2018). The baseline results are valid regardless of the state of the economy and hold for a battery of robustness checks.

These findings support the theoretical predictions of Dolado et al. (2018) and are also in line with the empirical findings of Romer and Romer (1999). Although income distribution is not the primary concern of central bankers, our results imply that it is a dimension they should not overlook. This is especially true because the income distribution may affect the transmission mechanisms of monetary policy.

The remainder of the paper is organized as follows: Section 2 discusses the estimation methodology and the identification strategy. Section 3 thoroughly describes the data. The fourth section presents LPs results, while the fifth and final section concludes the paper.

2 Estimation approach

The following section presents the two well-established empirical methodologies for estimating impulse responses: local projections and vector auto-regressions.

2.1 Local projections

We follow the general method proposed by Jordà (2005) and its very recent application to our context in Furceri et al. (2018) by estimating impulse response functions (IRF) from local projections (LPs). In its basic form, LP consists of a sequence of regressions of the endogenous variable shifted several steps ahead. As a result, the approach consists of estimating the following equation:

$$\Delta_h y_{i,t+h} = \alpha_i^k + \beta^k \Delta r_{i,t} + \theta^k X_{i,t} + \varepsilon_{i,t}^k \qquad (1)$$

where $\Delta_h y_{i,t+h} = y_{i,t+h} - y_{i,t}$ and corresponds to change in the top income variable from the base year t_0 up to year $t + h$, with $h = 1, ..., H$; $\Delta r_{i,t}$ denotes the change in the short-term interest rate; and $X_{i,t}$ refers to a vector containing a set of control variables. The latter includes the lags of $\Delta_t y_{i,t}$, $\Delta r_{i,t}$, and additional controls that could theoretically explain top income shares and be, at the same time, correlated with monetary conditions.

It is important to notice that each step of the local IRF is obtained from a different equation, and directly corresponds to the estimates of β^k. Thus, unlike in a VAR approach, the estimated coefficients contained in θ^k are not used to build the IRF. Instead, they only serve as controls and cleanse the β^k from the effects of past top income and monetary policy changes, in addition to contemporaneous and past changes in other macroeconomic variables (output and CPI, for instance). Moreover, the LP approach is intentionally "model-free", and therefore imposes fewer restrictions – with respect to VARs – for calculating IRFs. As shown by Jordà (2005), such approach confers numerous advantages. This estimation technique is actually (i) more robust to model misspecification, (ii) does not suffer from the curse of dimensionality, (iii) can more easily accommodate non-linearities and (iv) can also be estimated with simple regression techniques. However, it also has some drawbacks in terms of efficiency. The VAR approach is more efficient when the model is well specified. In what follows, we describe the benefits of LPs with respect to our research question.

First, LPs allow for more control variables – before running into dimensionality problems – that may influence top income shares and be, at the same time, correlated with monetary policy actions. The $X_{i,t}$ vector includes: the first difference of the log of the CPI, real GDP, real consumption, government expenditure-to-GDP ratio, house prices, stock prices, the total loans-to-GDP ratio and a trade openness ratio. In addition to these country-time variables, we include three global variables: the growth of U.S. utility patents applications as a proxy for technological progress, along with the world real GDP growth and the West Texas Intermediate (WTI) annual growth rate to parsimoniously remove global business cycle effects.[1]

The second benefit of LP is that it offers an original identification strategy to estimate dynamic causal effects. To build shock series, our strategy relies on external instruments, i.e. variables correlated with changes in short-term interest rates but not with the other macroeconomic shocks affecting the economy. Our aim is to obtain external sources of variation in short-term interest rates to provide quasi-random experiments and thereby more clearly identify causal effects. These types of strategies have recently attracted growing interest in applied macroeconomics (Jordà et al., 2015, 2017; Jordà and Taylor, 2016; Ramey and Zubairy, 2018; Stock and Watson, 2018). Regarding our research question, we notice that monetary policy is not likely to be driven by top incomes; therefore, the dynamic causal effect is clear (no simultaneity bias). However, even if income distribution is not a target of central banks, both inequality and monetary policy decisions depend on economic conditions, which may be improperly measured by the set of control variables in our regressions (omitted variable bias) (Furceri et al., 2018). Accordingly, this situation calls for the use of exogenous shocks to domestic monetary conditions rather than short-term interest rates. As is widely agreed upon in the literature, the challenge is to find external factors that would make the variations in monetary conditions a random treatment.

In this paper, we use the local projection-instrumental variable (LP-IV) method proposed by Jordà et al. (2015), Ramey (2016) and Jordà et al. (2017). We couple this method with the identification strategy of external variations in monetary conditions based on Jordà et al. (2015) and Jordà et al. (2017).

[1]As noticed by Jordà et al. (2019), adding time fixed effects would require almost one hundred additional parameter estimates.

The purpose here is to use the macroeconomic policy trilemma to find external variations in monetary conditions. The latter states that a country cannot simultaneously achieve free capital mobility, a fixed exchange rate and independent monetary policy. By pursuing any two of these goals, it is necessary to abandon the third. Building on the trilemma (Obstfeld et al., 2004, 2005; Shambaugh, 2004), we trace out episodes where external conditions can generate exogenous perturbations to the domestic short-term interest rate. Such perturbations are considered to be unrelated because the base country, for example, the U.S. during the Bretton Woods era, does not internalize the externalities of its own policy choices on partner countries. The trilemma links the domestic interest rate with the base country interest rate. A simple algebraic expression is given by

$$\Delta r_{i,t} = a + b[PEG_{i,t} * KOPEN_{i,t} * \Delta r_{i,t}^{base}] + \Theta X_{i,t} + \mu_{i,t} \tag{2}$$

where $PEG_{i,t}$ defines whether a country has a fixed ($PEG_{i,t} = 1$) or flexible exchange rate ($PEG_{i,t} = 0$); $KOPEN_{i,t}$ indicates whether a country is open ($KOPEN_{i,t} = 1$) or closed ($KOPEN_{i,t} = 0$) to international capital markets, and $X_{i,t}$ is a vector of macroeconomic controls in country i at time t.[2]

According to equation 2, variations in $r_{i,t}$ are related to external conditions (the base country) when there is perfect mobility of capital and a fixed exchange rate regime. Given this natural pseudo-experiment, it appears that the term $z_{i,t} = PEG_{i,t} * KOPEN_{i,t} * \Delta r_{i,t}^{base}$ has an exogenous influence on local monetary policy conditions. It therefore provides a source of variation in short-term interest rates that is exogenous to domestic conditions in terms of income distribution. As a result, $z_{i,t}$ constitutes a theoretically good external instrument. In what follows, as in Jordà et al. (2015), we use $z_{i,t}$ as an instrumental variable for the change in the short-term interest rate.

The third motivation for using LP is that it easily accommodates non-linearities.[3] This feature allows to enrich our analysis by checking whether the IRF of the top income share to a short-term rate shock is state-dependent, which is of great in-

[2]The controls include real per capita GDP growth; CPI inflation growth rate; real consumption growth; government expenditure growth; stock price growth; house price growth; the level of commercial openness; the change in the ratio of loans to the non-financial private sector to GDP; the WTI annual growth rate; U.S. patent activity growth and world GDP growth.

[3]The VAR literature also offers some solutions to deal with non-linearities. However, the richer structure of the VAR model entails several complications in computing IRFs, which often makes the estimation intractable in practice if we are outside the baseline framework.

terest because we use historical data that cover different monetary policy regimes. It also follows many studies, which highlight that the effects of monetary policy vary over the business cycle. In practice, we extend equation 1 and condition the effect of interest rates on the top-income variable by a state variable:

$$\Delta_h y_{i,t} = \alpha_i^k + \beta_1^k \Delta r_{i,t} * State_{i,t} + \beta_2^k \Delta r_{i,t} * (1 - State_{i,t}) + \theta^k X_{i,t} + \varepsilon_{i,t}^k \quad (3)$$

where $State_{i,t}$ is a variable indicating a specific state.

2.2 Panel VAR

In robustness to the LPs setting, we consider a panel-SVAR to generate the IRF of the top one percent's income share following an exogenous shock to the short-term rate.[4] Structural VARs are the traditional approach to identify structural monetary policy shocks and simultaneously trace out the corresponding impulse responses. We assume that our multi-country panel can be described by a small monetary SVAR extended to include the top income variable. Hence, our VAR contains six endogenous variables: the (log) WTI annual price (WTI_t), the top one percent's share of national income ($P1$), the (log) real GDP ($y_{i,t}$), the (log) CPI ($\pi_{i,t}$), the nominal short-term interest rate ($r_{i,t}$) and the (log) of stock prices ($s_{i,t}$).[5] Let $X_{i,t} = (WTI_t, P1_{i,t}, y_{i,t}, \pi_{i,t}, r_{i,t}, s_{i,t})$ be a vector of the six endogenous variables. The reduced form of the model can be represented as follows:

$$X_{i,t} = \mu_i + \Sigma_{l=1}^{L} A_l X_{i,t} + \nu_{i,t} \quad (4)$$

where the indices t and i relate to years and countries; μ_i corresponds to country fixed effects; L represents the number of lags on the endogenous variables included in the model, set at two according to the Akaike information criterion; A_l are $6 * 6$ matrices of unrestricted coefficients; and finally, $\nu_{i,t}$ is a vector of reduced form residuals related to a set of structural shocks as follows:

$$\nu_{i,t} = B\epsilon_{i,t} \quad (5)$$

where B is an invertible matrix of coefficients and $\epsilon_{i,t}$ a vector of structural shocks.

To disentangle the causal chain of events and identify structural shocks of in-

[4]We thank Ambrogio Cesa-Bianchi for providing his MATLAB toolbox for VAR analysis.

[5]The VAR is estimated in log-levels, which allows for possible co-integration relationships between the variables (Sims et al., 1990).

terest, different approaches exist. The standard one is Cholesky decomposition: it considers timing assumptions based on contemporaneous restrictions among the shocks in the VAR, which is typically done by assuming that B is a lower triangular matrix.[6] In particular, monetary policy shocks are identified as innovations to the short-term interest rate, which do not contemporaneously affect macroeconomic conditions. However, when using annual data, this identification scheme is questionable at best.

Therefore, to identify the coefficients of matrix B, we consider a different strategy and adopt an external instrument approach. The latter replicates the work of Mertens and Ravn (2014), Gertler and Karadi (2015) and Stock and Watson (2018). In practice, this identification strategy is implemented in three steps. First, the reduced form of the panel-SVAR model (eq. 4) is estimated equation-by-equation with a fixed effect estimator, which allows us to obtain the reduced-form residuals ($\nu_{i,t}^X$) along with an estimate of unrestricted coefficients (A_l). Then, the reduced-form residual of the short-term interest rate equation ($\nu_{i,t}^r$) is regressed on the instrumental variable ($z_{i,t}$) as defined for local projections (i.e., first-stage regression).[7] Finally, the fitted value of the short-term interest rate ($\hat{\nu^r}_{i,t}$) is regressed by OLS on the reduced-form residual ($\nu_{i,t}$).

Because the instrument $z_{i,t}$ is supposed to be orthogonal to all structural shocks with the exception of innovations to monetary conditions, equation 6 below provides contemporaneous responses of the endogenous variables to a one-unit interest rate shock. The dynamic response is then obtained by A_i and forward iteration.

$$\nu_{i,t}^X = \frac{b^X}{b^r}\hat{\nu^r}_{i,t} + \zeta_{i,t} \tag{6}$$

where b denotes the column of interest in matrix B, which corresponds to the impact of the structural interest-rate shock ($\zeta_{i,t}^r$) on each element of the vector $\nu_{i,t}$.[8]

[6] Another common way to identify exogenous changes in the short-term rate is the narrative approach of Romer and Romer (2004). This is the strategy adopted, for instance, by Coibion et al. (2017) and Furceri et al. (2018) to study the distributional effects of monetary policy. Nonetheless, this identification strategy is not tractable in our case because it requires (at least) forecasts of short-term interest rates, inflation and GDP growth, which are not available over the long run.

[7] Given that we estimate the VAR in levels, we marginally change the construction of our instrumental variable by considering the level of the interest rate of the base country rather than its first-difference.

[8] Because we are interested only in the effects of exogenous fluctuations in monetary condi-

We compare between this external instrument identification strategy and the standard Cholesky decomposition. In such case, the vector of endogenous variables has the following order: WTI_t, $P1_{i,t}$, $GDP_{i,t}$, $CPI_{i,t}$, $r_{i,t}$ and $s_{i,t}$.

3 Data description

3.1 Inequality

The Gini coefficient has long been used to analyze income inequality, as it illustrates the degree to which a variable is equally distributed across its population. However, the Gini index assigns relatively greater weight to observations in the middle of the distribution than to those located at the tails. This property prevents to account for aspects of concentration; for this reason, a sound alternative would be to consider instead measures that focus on the tails of the distribution. Such indicators take the shape of decile ratios or the shares of national income received by the 10, 1 or 0.01 percent of individuals with the highest market incomes.

In this paper, top income data are extracted from the World Inequality Database (WID, 2017). Specifically, the main top income variable is operationalized by the top one percent's pre-tax national income share (P1) in 12 advanced economies over the 1920-2015 period. The countries considered include: Australia, Canada, Italy, Germany, Denmark, France, the U.K, Japan, the Netherlands, Norway, Sweden and the U.S. We also conduct our empirical analysis by excluding the years of WWII from the sample.[9] As a robustness check, we test our model on the top 10 percent's pre-tax national income share (P10). In fact, as emphasized by Roine et al. (2009), P1 and P10 are quite different: while the first concentrates on individuals receiving important shares of capital income, the second contains more high-income earners. Hence, P10 would be considered more heterogeneous than P1 in that it gathers "rich" individuals who differ substantially in terms of their income sources. We also check how monetary policy affects changes *within the top of the distribution* using the P1/P10 ratio. Figure A1 in the appendix plots for each country P1 and P10 over the studied period. Finally, we extend our analysis by checking how changes in

tions, we do not have to identify all the coefficients of B.

[9] Table A1 traces out in detail the data sources and their availability for each country before and after WWII.

monetary conditions affect the ultra-rich. To do so, we extract data on the share of national income held by the top 0.1% and 0.01% from Atkinson and Piketty (2014).

3.2 Macroeconomic variables

We exploit the Jordà-Schularick-Taylor Macrohistory Database, which provides us with a long series of macroeconomic data. In this database, information on several macroeconomic variables are available from 1870 to 2016 and cover 17 developed economies.[10] In addressing the question of monetary policy and top one percent's income share, our paper also departs from the existing literature by building on several macroeconomic controls.

The set of specific control variables used for both LPs and the instrumental variable are summarized in Table A2 (see the appendix). They cover financial development, globalization, government spending, technological progress and global shocks. The way in which financial development – approached in our paper by the ratio of total loans to GDP – shapes top incomes remains an open question. While it was widely believed that it would reduce inequality through better access to credit for low-income households, recent findings (De Haan and Sturm (2017) provide a complete survey on this question) argue, on the contrary, that more finance mainly favors top income shares. Aside from financial development, real estate has become a strong factor in driving income inequality. As argued by Dustmann et al. (2018), shifts in housing costs in Germany severely exacerbated the rise in income inequality net of housing expenditures. For this reason, we control for this factor by adding a housing price index. Regarding globalization, Jaumotte et al. (2013) demonstrate, for a panel of 51 countries, that its effect on inequality has two offsetting tendencies: while trade globalization is associated with a reduction in inequality, financial globalization is associated with its increase. We control for the first using the ratio of imports and exports to GDP.

The ratio of government expenditure to GDP is also included in our control variables. In fact, based on a political economy model and an empirical analysis using data on OECD countries, Azzimonti et al. (2014) show that governments choose higher levels of public spending when inequality increases. Moreover, technological change has been repeatedly identified in the literature as playing

[10]Our sample is restricted only because of the limited availability of top income data.

a potent role in widening wage inequalities (see Acemoglu (1998), Card and DiNardo (2002), and Jaumotte et al. (2013) among others). One way to control for this factor consists of mobilizing data on patents. The use of such data would make it possible to measure the number of inventions, some of which are likely to become marketable. To that end, we rely on a dataset that tracks patent and grant activity in the U.S. since 1790. Specifically, we include data on utility patent applications, as these concern primarily "useful" inventions. The choice of focusing on the U.S. stems from the fact that it is considered the world technological frontier. Finally, we include the WTI, which is used as a standard benchmark in oil pricing and proxy for supply shocks, together with the world real GDP growth to account for global business cycle effects.

For the panel-SVAR-IV framework, we mobilize the following macroeconomic aggregates: the WTI, real GDP per capita (index, 2005=100), consumer prices (index, 1990=100), the short-term interest rate and stock prices. The top income variable and short-term interest rate are considered in level, while the remaining endogenous variables in the VAR enter in log-levels.

3.3 External instruments

As discussed above, the instrumental variable $z_{i,t}$ is the product of changes in the base country's short-term interest rate ($\Delta r_{i,t}^{base}$), the exchange rate regime ($PEG_{i,t}$) and the degree of capital control ($KOPEN_{i,t}$).[11] Following Jordà et al. (2015), the definitions of pegs prior to WWII are extracted from Obstfeld et al. (2004, 2005). After WWII, information on exchange rate regimes is completed using data provided by Ilzetzki et al. (2017). Table A3 in the appendix lists for each period and country of our sample the applicable exchange rate regime. Similarly, the indicator for capital mobility status builds on the index (which ranges from 0 to 100) initially introduced by Quinn et al. (2011). As in Jordà et al. (2015), we use this index rescaled to the unit interval, with 0 meaning fully closed and 1 fully open.[12] Figure A3 in the appendix plots, for our panel, changes in home interest rate $\Delta r_{i,t}$ against the constructed LP-IV.

[11] We consider a combination of U.S and French rates as the base country between 1920 and 1945. In the post-war era, we follow the selection of the base country short-term interest rate adopted by (Obstfeld et al., 2004, 2005).

[12] We reconstruct the instrumental variable using the indicator of capital mobility developed by Chinn and Ito (2006). The results are consistent with our baseline findings.

4 Results

This section reports the IRFs of the share in national income held by the top one percent to a positive shock to the short-term interest rate, both from local projections and vector auto-regression.

4.1 Local projections results

4.1.1 Baseline results

Our empirical setup builds primarily on LPs estimation, along with a novel identification of exogenous perturbations to monetary conditions. The first step is to assess the strength of our instrument. To do so, we estimate, in the context of equation 3, a first-stage regression of changes in the short-term interest rate on the instrument $z_{i,t}$ and the aforementioned macroeconomic controls, including country fixed effects. The first-stage regression results are reported in Table 1 and underline the importance of the pass-through from base to home rates through several periods. The coefficient estimates of the instrument $z_{i,t}$ are statistically significant at the 1% level and range between 0.52 and 0.53 from year 0 (when the shock is felt) to year 4. Similarly, the F-statistics feature high values across samples. Notice that Stock et al. (2002) recommend a threshold of ten for the first-stage F-statistic. Thus, we can now proceed to analyze the LPs responses of the top one percent share to exogenous fluctuations in the short-term interest rate.

Table 1: Local projection-IV: First-stage results

Δ ***Short-term interest rate***	Year 0	Year 1	Year 2	Year 3	Year 4
IV	0.53***	0.53***	0.53***	0.52***	0.52***
	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)
F-statistic	32.89	31.36	29.78	27.5	27.77
Observations	642	628	616	603	589

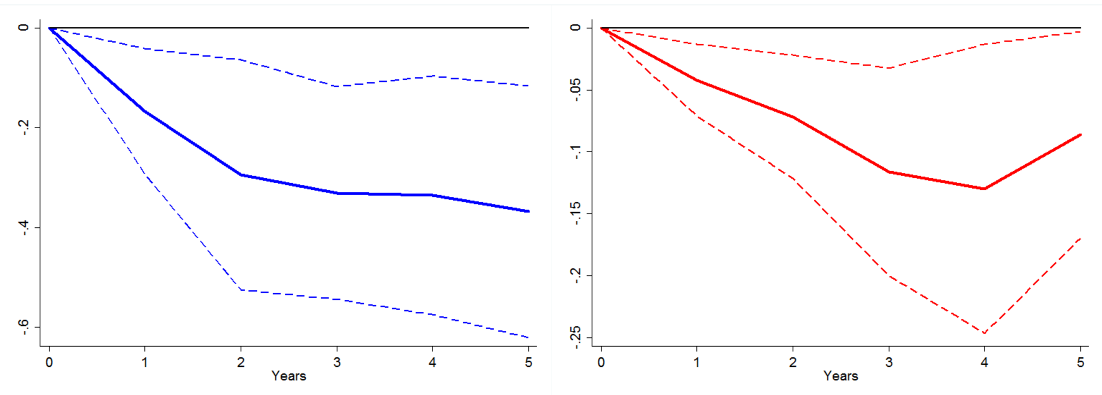
Note: *, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively. Country-based cluster-robust standard errors in parentheses. The short-term interest rate is regressed on the instrumental variable, using country fixed effects and macroeconomic controls (the change in the short-term interest rate; the real per capita GDP growth; the CPI inflation growth rate; the real consumption growth; the government expenditure growth; stock prices growth; house prices growth; the level of commercial openness; the change in the loans to non-financial private sector to GDP ratio; oil prices annual growth rate; the U.S. patent growth and the world GDP growth). We include contemporaneous terms and two lags.

The results obtained from the estimation of equation 3 by LPs are presented in Figure 1. The two graphs illustrate cumulative IRFs of the top income variable to an unexpected negative shock to the short-term interest rate – via the instrument – with the associated confidence bands, using both the instrumental variable and OLS. The initial glance at the IRFs suggests that monetary tightening significantly and durably decreases the share of national income held by the top one percent. Inasmuch as our empirical model is linear, the exact opposite effect holds with respect to expansionary monetary policy (see Figure A4 in appendix). This evidence is in line with the theoretical predictions of Dolado et al. (2018) and the empirical results documented by Romer and Romer (1999). However, they contradict the empirical findings of Coibion et al. (2017) and Furceri et al. (2018).

Precisely, an unanticipated increase of 100 b.p. in the short-term interest rate (graph (b) on the right) reduces the share of the top income variable by approximately 0.12 percentage points three years after the shock. Nonetheless, the effects on P1 are, interestingly, more pronounced under the instrumental variable. Indeed, a perturbation to the domestic interest rate $r_{i,t}$ via the instrument $z_{i,t}$ (graph (a)) reduces P1 by 0.32 and 0.35 percentage points four and five years following the shock, respectively.

Figure 1: Top one percent LPs to a positive short-term interest rate shock

(a) IV - P1 response (b) OLS - P1 response



Note: The figure shows cumulated impulse responses of the top one percent's share in national income to an unexpected 100 b.p. increase in the short-term interest rate. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

These differences are clearer in Table 2, which jointly reports coefficient estimates of OLS and LP-IV. We compare the results obtained by the two methods in order to assess the degree of attenuation bias in the OLS estimation. In doing so, we notice that the impulse responses obtained under both methods exhibit a relatively similar pattern. However, the coefficient estimates obtained via OLS are substantially lower than those produced by the IV. For example, a positive shock to the short-term interest rate reduces P1 in year 4 after the shock by 0.13 percentage points using OLS and by 0.32 percentage points in the IV estimation. Note that Jordà et al. (2015) – who investigate the effect of monetary policy on housing prices in the very long run – document the same observation in a more or less similar magnitude. How should one account for such discrepancy between the OLS and LP-IV coefficient estimates? Some limitations of OLS regression may be at work and explain this contrast. As noted in section 2.1, given that monetary policy is not driven by top income shares, simultaneity bias is not a concern. However, both variables are affected by economic conditions, some of which may be omitted from the set of control variables. Consequently, as put by Jordà et al. (2015): "*the instrument is capturing movements in the short rate that are more likely to reflect exogenous fluctuations than fluctuations we could identify using a simple OLS regression control strategy*".

As shown in Figure 2, we conduct a battery of robustness checks on the baseline LP-IV estimation, which includes additional top income indicators, different model specifications plus a pre-crisis analysis. The impulse response in graph (a) documents the cumulative impact of orthogonal changes in the instrument on P10. The result does not depart from the IRF associated with P1, although the end-of-period effect is larger (0.42 percentage points decrease five years after the shock). Graph (b) in Figure 3 shows the response of the P1/P10 ratio, which takes into account income changes within the top of the distribution. The corresponding IRF outlines that monetary tightening narrows the gap among top income households. The impulse response presented in graph (c) indicates that the baseline effect of restrictive monetary policy on P1 is not affected when one considers only the post-WWII period, except that the response is not statistically significant in the short run. We then assess whether the effects of monetary policy on the top income variable are robust when excluding the U.S. from the sample: the impulse response shown in graph (d) is very similar to that obtained in the baseline estimation.

Table 2: Local projection: OLS and IV estimation results

IV estimates - P1	Year 1	Year 2	Year 3	Year 4	Year 5
Δ *Short-term interest rate*	-0.16**	-0.28*	-0.32**	-0.32**	-0.35**
	(0.08)	(0.15)	(0.14)	(0.15)	(0.16)
R^2	0.054	0.063	0.076	0.097	0.092
Kleibergen-Paap	32.89	31.36	29.77	27.503	27.773
Observations	642	628	616	603	589
OLS estimates - P1					
Δ *Short-term interest rate*	-0.04***	-0.07**	-0.12**	-0.13**	-0.08
	(0.02)	(0.03)	(0.05)	(0.07)	(0.05)
R^2	0.083	0.102	0.105	0.124	0.133
Observations	656	640	627	614	600

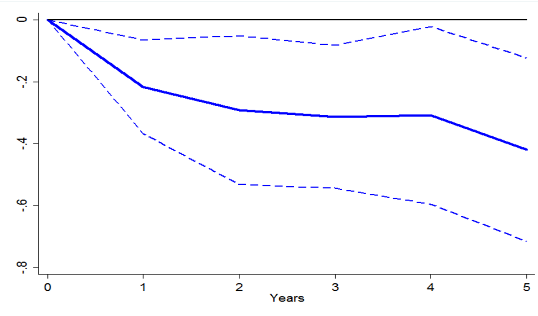
Note: Country-based cluster-robust standard errors are reported in parentheses below the coefficient estimates. The controls include the twice-lagged terms of (i) the change in the short-term interest rate; (ii) the change in top income share; and the contemporaneous and twice-lagged terms of (iii) real per capita GDP growth; (iv) the CPI inflation rate; (v) stock price growth; (vi) real per capita consumption growth; (vii) the level of financial development; (viii) the level of commercial openness; (ix) house price growth; (x) government expenditure; (xi) patent activity; (xii) world GDP growth and (xiii) oil prices growth. We report the Kleibergen and Paap (2006) statistic for weak instruments. *, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

An additional robustness check consists of estimating equation 3 with country fixed effects while omitting the rich set of control variables. This exercise is valuable because it assesses whether or not the IV exclusion restrictions are violated. In fact, a correctly specified instrument would be sufficient to avoid potential endogeneity bias. The evidence depicted in graph (e) does not contradict our main result. The impulse response presented in graph (f) adds more lags to equation 3 and suggests that the LP framework remains robust to different lag numbers.[13] One concern that may arise relates to the fact that our sample period covers episodes where the short-term interest rate reaches the lower bound and becomes inadequate to measure monetary policy stance. To address this issue, we perform two robustness checks: first, we use long-term interest rates as the monetary policy instrument, and second, the Great Recession period is omitted from the sample (the empirical analysis was conducted until 2006). The results are reported in graphs (g) and (h) and show strong consistency with the LP-IV baseline finding. We notice, however, that the maximum impact on P1 is higher when long-term rates are used as the monetary policy instrument.

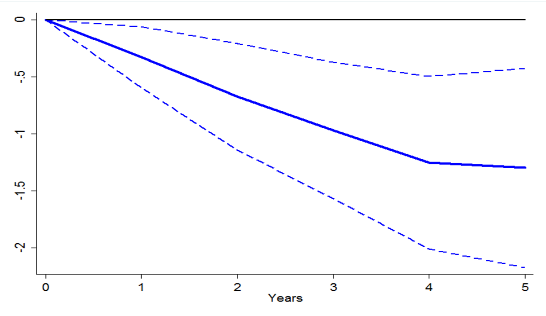
[13]The baseline LP-IV result is also robust to the introduction of only one lag (see Figure A6 in the appendix).

Figure 2: Top one percent income share LPs responses to a positive short-term interest rate shock via the instrument: Robustness check

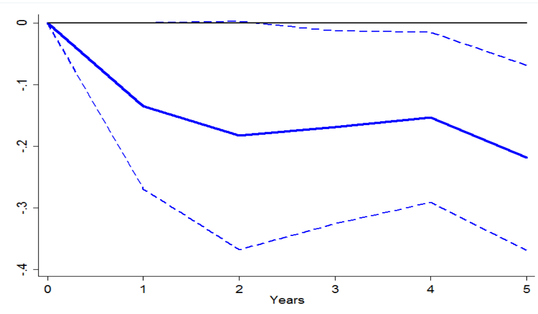
(a) P10 response



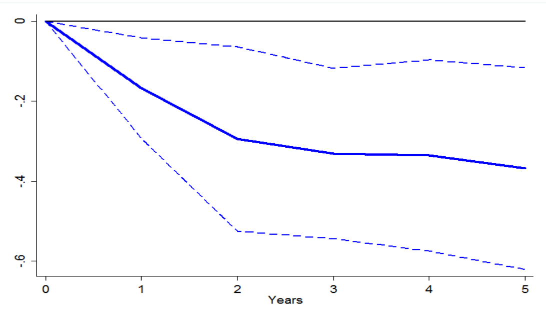
(b) P1/P10 response



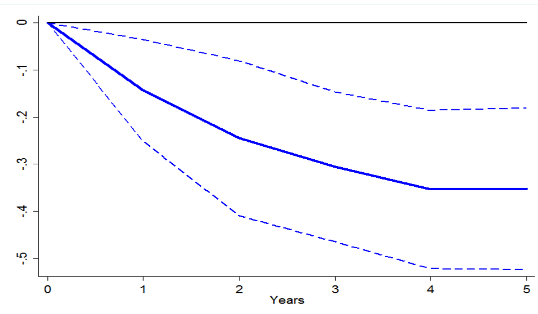
(c) P1 response - post-WWII



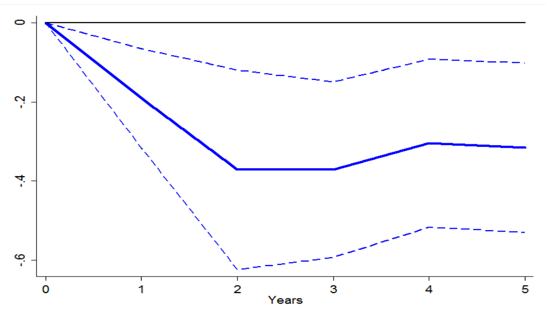
(d) P1 response - without the U.S.



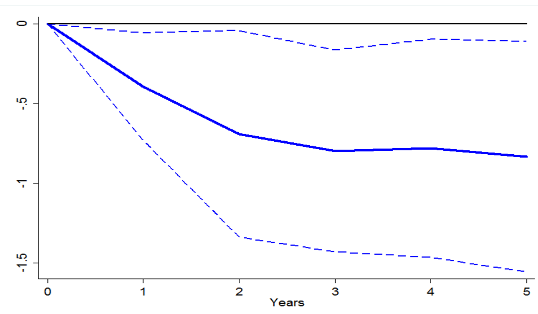
(e) P1 response - No control



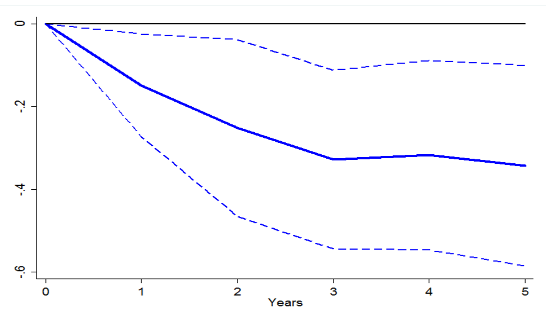
(f) P1 response - More lags



(g) P1 response - long-term rates



(h) P1 response - without the Great Recession



Note: The figure shows cumulated impulse responses of the top one percent income share to an unexpected 100 b.p. increase in the short-term interest rate via the instrument. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

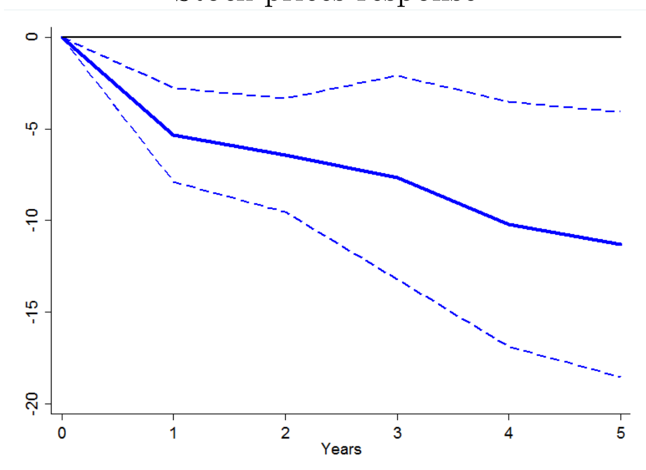
Precisely, the share in national income held by the top one percent decreases by 0.75 percentage points in year 4 after the shock. In addition, IRFs of the top 0.1% and 0.01% reported in figure A5 of the appendix confirm that monetary tightening reduces the share of national income held by the ultra-rich. Overall, all these checks highlight the stability of our estimates and confirm the reliability of our instrument in the context of LPs.

4.1.2 Insights on the transmission channel

Our baseline result suggested that an unexpected monetary tightening lowers the share of national income held by the top one percent. That said, it is relevant to investigate one of the underlying transmission mechanisms of monetary policy towards top income shares. We demonstrate here that our evidence can support the income composition channel. That is, considering the heterogeneity in income sources between households, monetary policy will probably affect the income distribution if it disadvantages some types of income. Figure 3 depicts the cumulative IRF of stock prices to an exogenous shock in monetary conditions. It indicates that the negative effect of restrictive monetary policy on top income shares is likely to be channeled via lower stock prices. Specifically, a monetary tightening significantly reduces stock prices in year 3 and year 4 by 7.64 and 10.20 percentage points, respectively. Such finding is quite expected because financial assets amount to an important share of the very rich's income.

Figure 3: Insights on the income composition channel

Stock prices response



Note: The figure shows cumulated IRFs of stock prices to an unexpected 100 b.p. increase in the short-term interest rate via the instrument. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

4.1.3 State-dependent effects

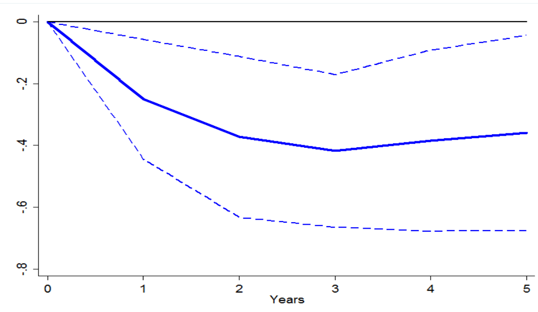
The results we have reported in the previous sections support the accord that monetary policy tightening decreases the share of national income held by the top one percent and vice versa. There is, however, a potential pitfall because our sample encompasses very different economic regimes. Moreover, several studies indicate that some economic variables, such as the short-term interest rate, may, for instance, behave very differently during economic downturns. To overcome this limitation, we take advantage of the fact that the LP-IV method easily accommodates non-linearities. It is convenient to explore whether the effects of changes in monetary conditions on top income shares are state-dependent. Thus, we allow the impact of monetary policy on the top income variable to depend upon the state of another variable (see equation 3). In this way, we can compute conditional impulse responses in a particular regime.

We consider four factors that potentially lead to different impulse responses of monetary policy: the state of the economy over the business cycle, the inflation regime, credit cycles and monetary policy stance. The episodes of business cycle are identified using the Hamilton (2018) filter and take the value of one in the case of an economic expansion, while they are attributed the value of zero during recessions. The same approach is adopted to identify credit booms and slumps. With respect to inflation, a high-inflation episode is defined as a period during which inflation is above its country-specific fourth quartile. Conversely, a country features a low-inflation regime when inflation is below its first quartile. With regard to monetary policy stance, we define a binary variable taking a value of one when there is a positive variation of the short-term interest rate (i.e., monetary tightening) and 0 in the case of a negative variation (i.e., monetary easing). Finally, we check whether the responses in the aforementioned regimes are statistically different from each other by conducting a Wald Chi-Squared Test.

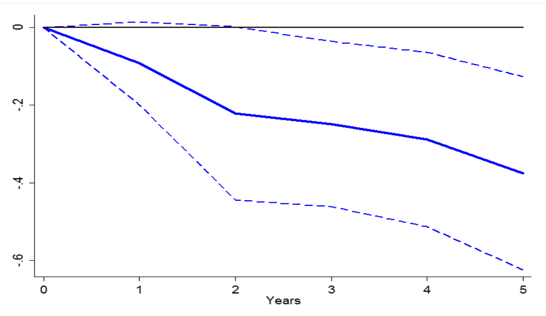
Figure 4 reports the impulse responses estimated with the state-dependent effect model and the instrumental variable (equation 3), for the 4 factors previously described. Overall, the displayed impulse responses do not conflict with the previous results: the effect of monetary policy on P1 continues to hold, irrespective of the state of the economy. As shown in graph (a), monetary policy has more immediate effects on P1 during expansions than during recessions. However, there is no significant difference regarding its effect on the medium run.

Figure 4: Top one percent LPs responses to a positive short-term interest rate shock: state-dependent effects

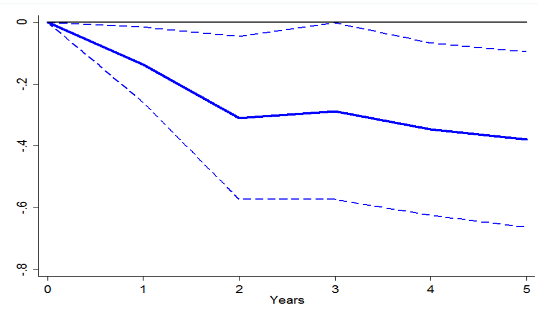
(a) Business cycle boom



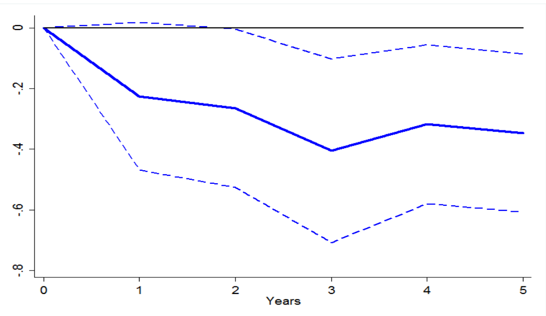
(b) Business cycle slump



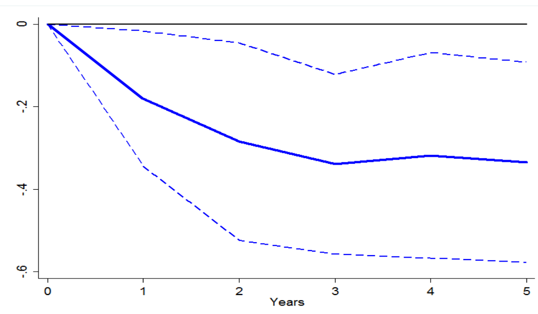
(c) Inflation > to the country's fourth quartile



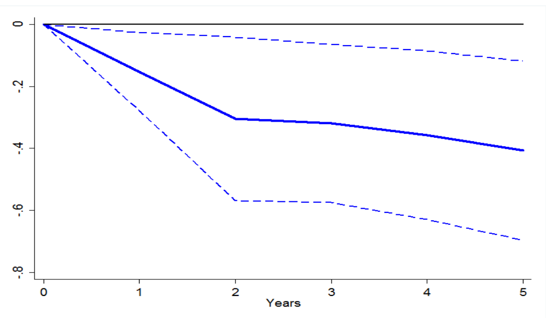
(d) Inflation < to the country's first quartile



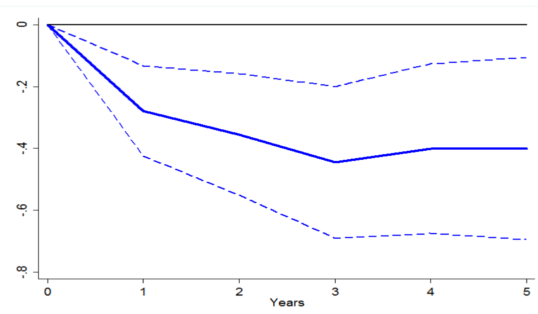
(e) Credit boom



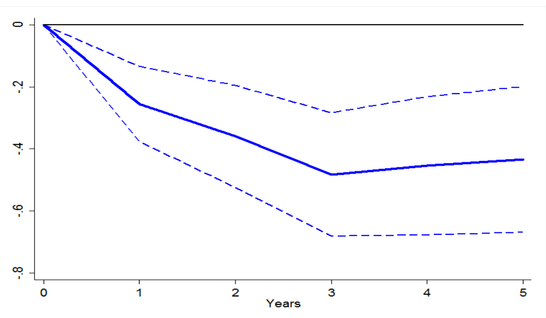
(f) Credit slump



(g) Monetary tightening



(h) Monetary easing



Note: The figure shows, under several regimes, the cumulated impulse responses of the top one percent income share to an unexpected 100 b.p. increase in the short-term interest rate via the instrument. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

Interestingly, it also appears from graph (c) that monetary tightening has a strong effect in the short run when considering a high-inflation regime. This makes sense considering that inflation itself is a redistributive tool, which, according to Paarlberg (1993) *"...steals from widows, orphans, bondholders, retirees, annuitants, beneficiaries of life insurance, and those on fixed salaries, decreasing the value of their incomes"*. In addition, the impulse responses presented in graphs (e) and (f) show that the impact of changes in the short-term rates on P1 is not affected by credit cycles. In fact, during episodes of credit booms and slumps, restrictive monetary policy produces very similar impacts on P1. Similarly, graphs (h) and (g) support our baseline result when considering potential asymmetries between expansionary and contractionary monetary policies. Table 3, which reports the Wald test results for each regime along with their respective p-value, indicates that the responses in the respective regimes are not statistically different from each other.[14]

Table 3: Wald Chi-Squared Test of the difference of the cumulated effect of the interest rate shock between the two states

State:	Business cycle boom/slump	Inflation low/high	Credit boom/slump
chi2 Year 5	0.75	0.63	1.05
Prob Year 5	0.39	0.43	0.30

4.2 Panel VAR results

Figure 5 presents impulse responses of the top one percent's income share and standard macroeconomic variables to a shock in the short-term rate. The left-hand panels display impulse responses when structural monetary shocks are identified using external instruments. The right-hand panels show the case where the standard Cholesky decomposition of the residuals is used to identify structural shocks. The presented panels feature the estimated impulse responses with 90% confidence bands, generated by bootstrapping methods.[15]

As previously mentioned, the identification using an external instrument involves two stages. The first stage consists of regressing the residuals computed from the short-term interest rate equation of the panel-SVAR (equation 4) on our external instrument. Table 4 verifies if $z_{i,t}$ has the potential to be a valid

[14] We are unable to conduct such test for monetary policy stance because the effect of each regime (i.e., monetary tightening/easing) is estimated in a separate LPs specification.

[15] In the same way as Mertens and Ravn (2013) and Gertler and Karadi (2015), we use wild bootstrap with 1000 draws.

instrument. The correlation between the instrument and the innovation in the short-term interest rate is perceived to be strong: the coefficient is significant at the 1% level and the F-statistic exceeds 22.

Table 4: SVAR-IV - First-stage results

	Short-term interest rate
IV	0.10***
	(0.02)
R^2	0.04
F-statistic	22.05
Observations	707

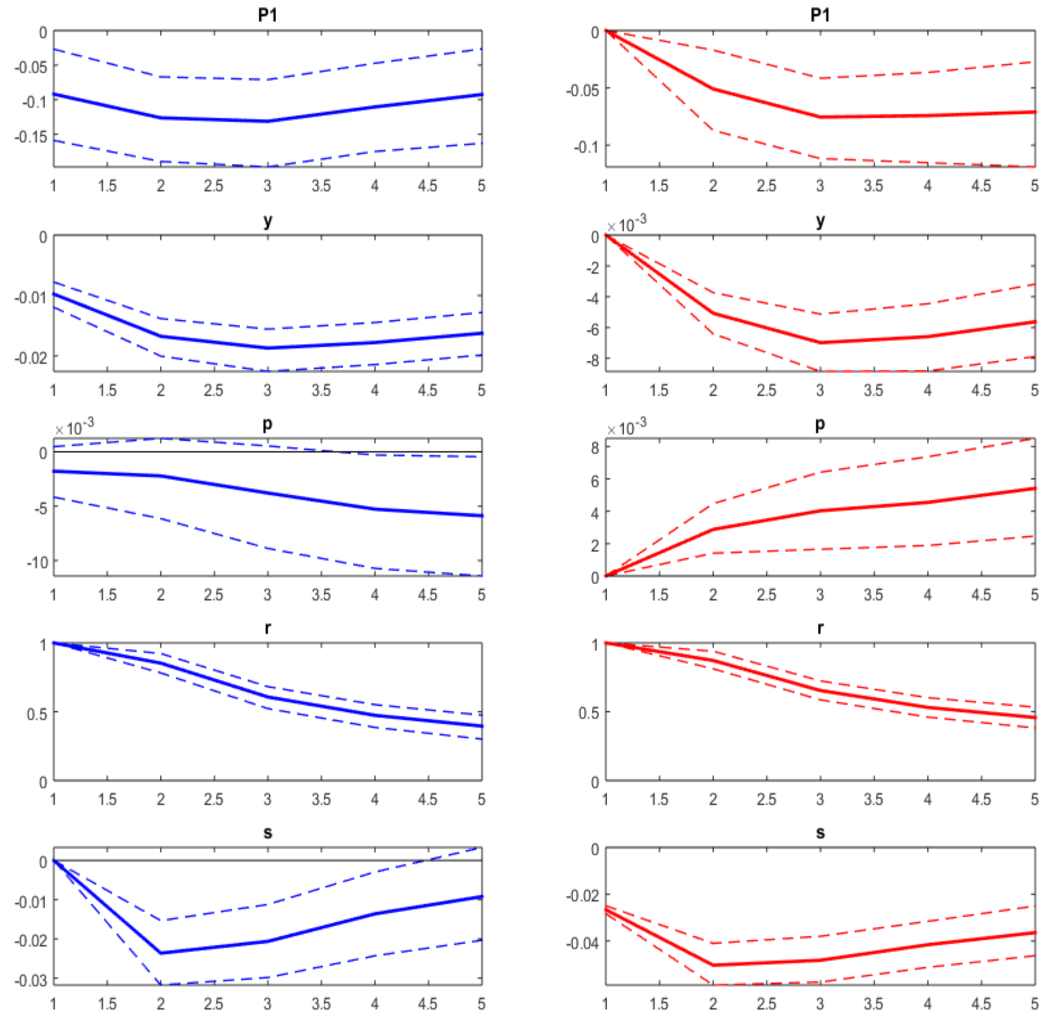
Note: *, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

Confident in our instrument choice, we can now proceed to the second stage and comment on the panels presented in Figure 5. We focus on the results generated from the external instrument. The impulse response of the top one percent pre-tax national income share (P1) reported in the top-left corner of Figure 5 indicates that a 100 b.p. unexpected monetary tightening reduces the level of P1 by 0.1. Five years after the shock, the reduction in the P1 indicator settles at approximately the same level. Because our panel-SVAR-IV is linear in the lagged variables and model parameters, the responses are symmetric under both positive and negative shocks to the short-term interest rate, which implies that monetary policy loosening increases the top one percent's income share.

A typical concern in the monetary VAR literature is the finding of a price puzzle, i.e., the fact that unexpected monetary tightening leads to increased inflation. Hence, a necessary requirement is to check the consistency of our external instrument with standard macroeconomic predictions. First, we observe that the estimated shape of the inflation response to an exogenous perturbation in monetary conditions does not exhibit counter-intuitive effects. Indeed, the estimated impulse response is negative, even though it is not statistically significant at the 90% confidence level (which is in line with the finding of Gertler and Karadi (2015)). Second, the dynamic effects of monetary policy on real GDP and stock prices are fairly standard. Two years after the shock, monetary tightening induces a reduction in GDP of approximately 1.5% and a decline in stock prices of more than 2% (since both variables enter the VAR in log-level).

Figure 5: Top one percent PVAR responses to a short-term interest rate shock

(a) SVAR-IV (b) SVAR-Cholesky



Note: The figure shows impulse responses of the top one percent income share to an unexpected 100 b.p. increase in the short-term interest rate. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

Figure A7 in the appendix displays several robustness checks on the baseline panel-SVAR-IV result. Similar to LPs, we conduct additional estimations, which consider alternative top income indicators (P10 and the P1/P10 ratio) and different model specifications (i.e. post-WWII era estimation, exclusion of the U.S. from the sample and using the long-term interest rates as the monetary policy instrument). We also check the sensitivity of the results before the Great Recession. All the mentioned tests are statistically significant and confirm that monetary tightening reduces the top one percent's income share.

5 Conclusion

This paper sought to investigate the distributional consequences of monetary policy via top income shares between 1920 and 2015 in 12 advanced economies. The central idea that guided this paper's argument is that the existing literature on the distributional effects of monetary policy uses mainly survey-based estimates of income inequality alongside a shorter time coverage. This approach translates into lower inequality estimates – particularly due to the underestimation of business and capital incomes of rich households – and a lower coverage of exceptional macroeconomic occurrences (recessions, sovereign defaults, etc.). We address these shortcomings by studying how changes in the short-term interest rate over a century impacted the share of national income held by the top one percent while controlling for the determinants of inequality. To do so, we combined two large datasets: (i) the World Inequality Database (WID) to extract tax-based data on top income shares and the Jordà-Schularick-Taylor Macrohistory Database, which allows us to access a large series of macroeconomic and financial variables.

Our empirical strategy is based on Local Projections (LPs) to obtain the impulse responses of top income shares to an unexpected positive variation in the short-term interest rate. The motivation for this empirical setup is threefold: (i) LP is a "model free" approach, which allows us to control for several factors that may affect top income shares and be, at the same time, correlated with monetary policy actions; (ii) it offers a quasi-natural experiment as an identification strategy, where exogenous perturbations to the short-term rate are driven by factors unrelated to domestic economic conditions and (iii) it easily accommodates non-linearities so that we could estimate potential state-dependent effects of monetary policy on the top one percent. As a robustness, we have considered the estimation of a panel-SVAR, with a focus on an IV identification approach, with which we drew a parallel relying on a standard Cholesky identification scheme.

The results obtained from both empirical methods indicate that *tight monetary conditions strongly decrease the top one percent's income share and vice* *versa.* In fact, following a positive perturbation to the domestic short-term rate via the external instrument, the share of national income held by the richest one percent decreases by 0.16 to 0.35 percentage points, according to LPs estimates. This effect is persistent and statistically significant in the medium run. We also demonstrate that the reduction in the top one percent's share is arguably the consequence of lower asset prices, which is consistent with the income composition channel. The baseline results hold under a battery of robustness checks, which (i) consider alternative top income measures and (ii) exclude the U.S. economy from the sample, (iii) specifically focus on the post-WWII period, (iv) remove control variables, (v) test different lag numbers, (vi) use long-term interest rates as the monetary policy instrument and (vii) omit the Great Recession from the sample. Furthermore, the state-dependent effects version of our model indicates that our conclusions are robust, regardless of the state of the economy.

For future research, we would like to test the effects of monetary policy on different income deciles, which focus exclusively on poor and middle-class households (i.e., the bottom 5% or 1% with the lowest market incomes). In the same perspective, are the obtained results also valid for wealth inequality? This aspect is important because wealth is more unevenly distributed than incomes. Moreover, given that we use pre-tax data, policymakers may be interested in the effects of monetary policy on income distribution, net of the contribution of fiscal policy. Finally, in the spirit of the corresponding literature, the empirical approach adopted in this paper considers only the global effects of monetary policy on the income distribution. That is, we do not identify all the transmission channels through which monetary policy affects top incomes. That said, what policy implications can we draw from these findings for the ongoing debate on monetary policy and income distribution? Central bankers need to be attentive not only to the aggregate consequences of monetary policy but also to their side effects.

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Appendix

Table A1: Data sources and periods of top income shares

Country	Period P1	Period P2	Details
Australia	1921-1938	1946-2015	WID (2017)
Canada	1920-1938	1946-2010	WID (2017)
Germany	1925-1938	1950-2013	WID (2017), Dell et al. (2007)
Denmark	1920-1938	1946-2010	WID (2017)
France	1920-1938	1946-2014	WID (2017)
U.K.	-	1951-2014	WID (2017)
Italy	-	1974-2009	WID (2017)
Japan	1920-1938	1947-2010	WID (2017)
Netherlands	1920-1938	1952-2012	WID (2017)
Norway	-	1948-2011	WID (2017)
Sweden	1920-1935	1946-2013	WID (2017)
U.S.	1920-1938	1946-2015	WID (2017), Atkinson et al. (2015)

Note: There are years with missing values in each subperiod.

Table A2: Control variables definition

Variable	Variable definition	Source
Hpnom	House prices growth (real index, 1990=100)	Macrohistory Database JST
Stocks	Stock prices index growth (real index)	Macrohistory Database JST
CPI	Consumer Price Index year-over-year growth	Macrohistory Database JST
Tloans	Ratio of total loans to non-financial private sector to GDP	Macrohistory Database JST, own calculations
Com_open	Ratio of imports and exports to GDP	Macrohistory Database JST, own calculations
gdp_pc	country Real GDP per capita (index, 2005=100)	Macrohistory Database JST
cons_pc	country Real consumption per capita (index, 2006=100)	Macrohistory Database JST
expenditure	government expenditure-to-GDP ratio	Macrohistory Database JST
US_patents	Growth rate of utility patents applications	United States Patent and Trademark Office
World_gdp	World real GDP growth	Macrohistory Database JST
WTI	West Texas Intermediate (WTI) annual price growth	Federal Research Economic Data (FRED), St.Louis Fed

Note: This set of control variables has been used in the context of local projections. To ensure stationarity, real indexes are obtained by dividing the variables by CPI, and growth rates are computed in logs.

Table A3: Exchange rate regimes

Country	Fixed	Floating
Australia	1920-1938, 1946-1983	1939-1945, 1984-2015
Canada	1920-1938, 1946-2015	1939-1945
Germany	1920-1938, 1946-1972, 1999-2014	1939-1945, 1973-1998
Denmark	1920-1938, 1946-2014	1939-1945
France	1920-1938, 1949-2014	1939-1948
U.K.	1920-1938, 1946-2008	1939-1945, 2009-2015
Italy	1920-1938, 1949-2014	1939-1948
Japan	1920-1938, 1948-1977	1939-1947, 1978-2015
Netherlands	1920-1938, 1946-2014	1939-1945
Norway	1920-1938, 1946-2014	1939-1945
Sweden	1920-1938, 1946-2014	1939-1945
U.S.	1920-1938	1939-2015

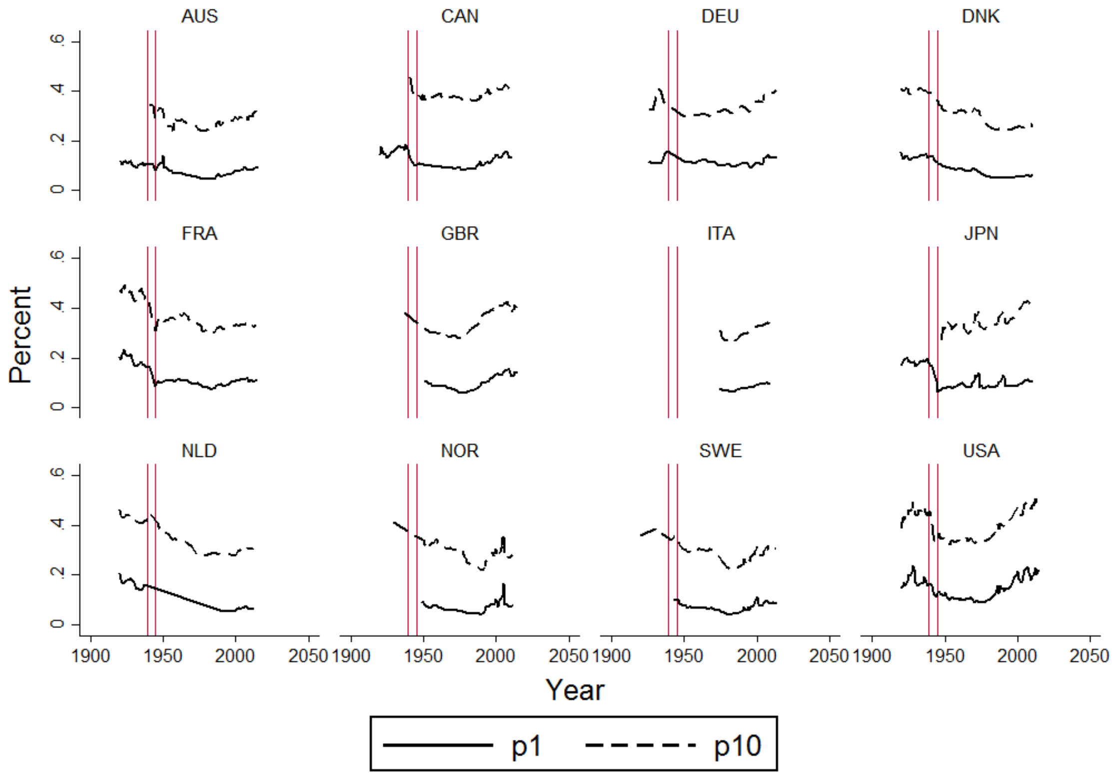


Figure A1: Top income shares over time: 12 countries

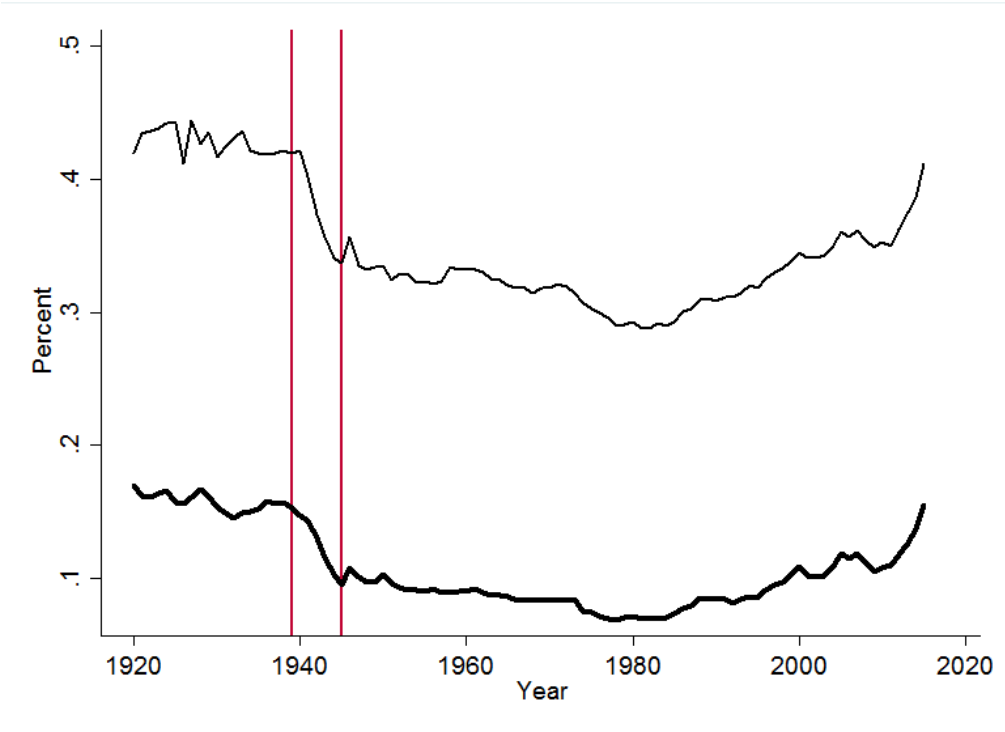


Figure A2: Top income shares over time: 12 countries

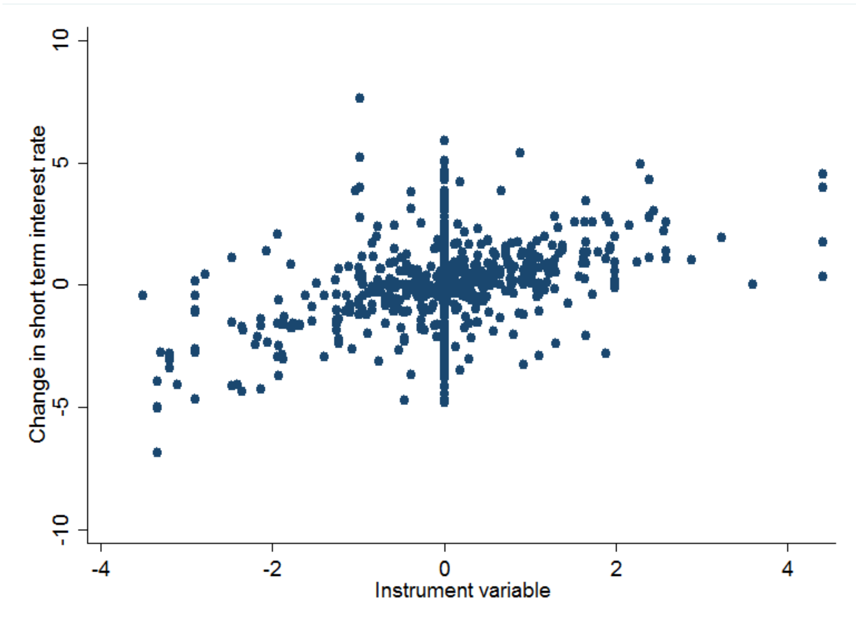
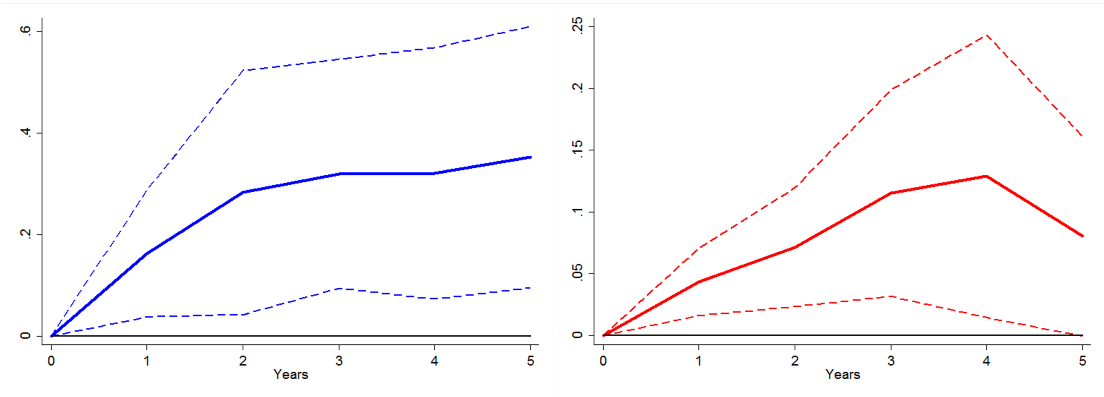


Figure A3: Jorda, Schularick and Taylor based IV: change in short-term interest rate in home and base countries

Figure A4: Top one percent income share LPs to a negative short-term interest rate shock

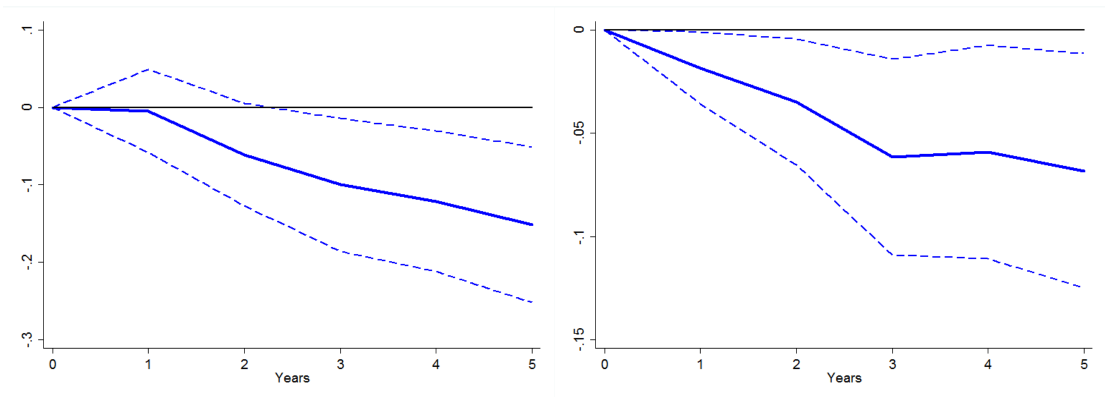
(a) IV - P1 response (b) OLS - P1 response



Note: The figure shows cumulated impulse responses of the top one percent income share to an unexpected 100 b.p. decrease in the short-term interest rate. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

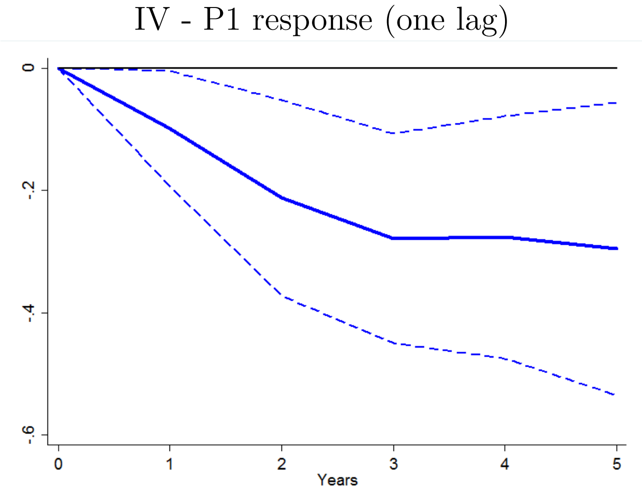
Figure A5: Top 0.1% and 0.01% percent income shares LPs to a positive short-term interest rate shock

(a) IV - Top 0.1% response (b) IV - Top 0.01% response



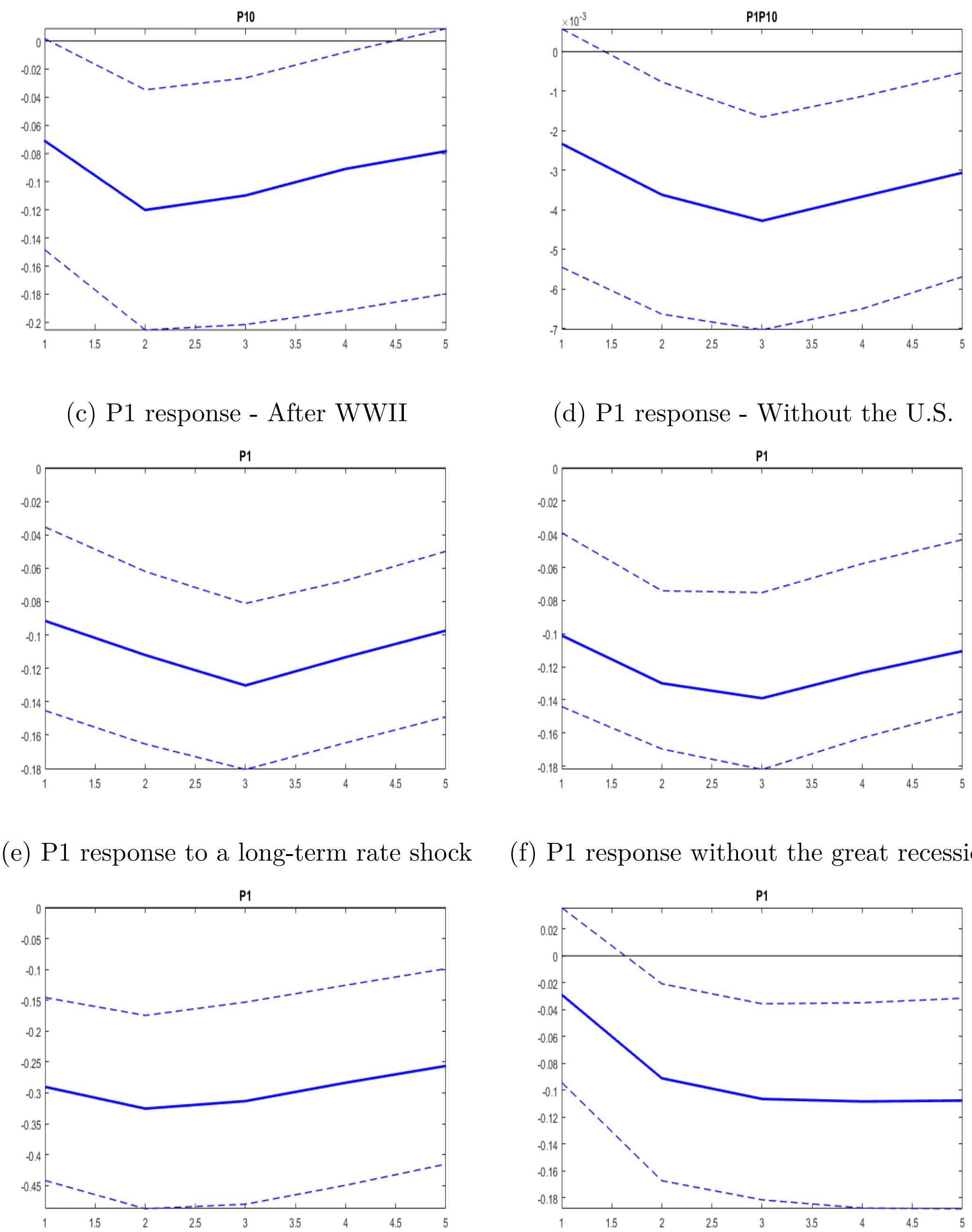
Note: The figure shows cumulated impulse responses of the top 0.1% and 0.01% income shares to an unexpected 100 b.p. increase in the short-term interest rate via the instrument. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

Figure A6: Top one percent income share LPs to a positive short-term interest rate shock: one lag



Note: The figure shows cumulated impulse responses of the top one percent income share to an unexpected 100 b.p. increase in the short-term interest rate via the instrument. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).

Figure A7: Top one percent PVAR responses to to a short-term interest rate shock via the instrument: Robustness check



(a) P10 response

(b) P1-P10 response

(c) P1 response - After WWII

(d) P1 response - Without the U.S.

(e) P1 response to a long-term rate shock

(f) P1 response without the great recession

Note: The figure shows impulse responses of the top one percent income share to an unexpected 100 b.p. increase in the short-term interest rate via the instrument. The dashed lines represent 90% confidence bands generated by bootstrapping (1000 draws).