

Monetary policy and the well-being of the rich: Evidence from a century of modern economic history

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Abstract

This paper examines the distributional implications of monetary policy from a long-run perspective, with data spanning a century of modern economic history in 12 advanced economies. We employ two complementary empirical methodologies for estimating the dynamic responses of the top 1% income share to a monetary policy shock: vector auto-regressions and local projections. The monetary policy shocks are assessed via different approaches. We notably exploit the implications of the macroeconomic policy trilemma to identify exogenous variations in monetary conditions. This offers a genuine orthogonal perspective to capture fluctuations in short term interest rates unrelated to home economic conditions. The obtained results indicate that expansionary monetary policy strongly increases income inequality and vice-versa. These results have important implications for monetary policy coordination and the design of economic policies that aim at reducing income inequality.

JEL Codes: D63, E62, E64

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1 Introduction

"All public policy is distributional, be it monetary, fiscal, structural or social. The reason I know this is because redistribution is the way public policy works; it is what policy does for a living. Some policies redistribute resources between agents at a point in time. Others redistribute resources between agents over time. If policy is not working through one of these channels, it is not working".

Haldane (2018)

The last decades have been marked by a substantial rise in income and wealth inequality across the developed world. Low-income households in advanced economies have seen their wages stagnating, while wealth has never been so concentrated since the onset of the 20th century. Such patterns motivated a global narrative on inequality, which spilled within circles of policymaking. Especially since uneven income and wealth distribution appear to support excessive household indebtedness (see e.g., Kumhof et al. (2015) and Coibion et al. (2014)), and consequently fuel financial instability (Rajan, 2010).

The potential distributional effects of monetary policy have recently erupted in the inequality debate, as a consequence of the unconventional instruments that central banks implemented following the financial crisis. This was unusual, since it is widely accepted that central banks should not worry about inequality. In fact, they are independent from the political process and therefore, dealing with distributional matters — which fall instead within abilities of fiscal policy — go outside their mandate. Despite that, the ultra low interest rates environment alongside large asset purchase programs, are suspected to wipe out modest household savings and push up asset prices. Meanwhile, central bankers in the Eurozone and elsewhere such as Draghi (2016) or Bernanke (2015), strongly believe that their non-standard monetary policies had modest distributional implications. They argue instead that post-crisis monetary policy toolkit allowed to restore growth and enhance employment levels, which primarily favored low-income households. This debate underlines, in the spirit of Coibion et al. (2017), that the outcome of monetary policy on income inequality would be channeled through : (i) households income composition (some mainly rely on labor incomes while others may receive other forms of revenue as rents or dividends), (ii) earnings heterogeneity to business cycle fluctuations (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 mainly held by rich households and could be the first to benefit from higher asset prices).

In this context, macroeconomic research has been increasingly devoted to analyze the collateral effects of monetary policy on income and wealth distribution. As far as theoretical contributions are concerned, they have mainly built on New Keynesian frameworks. Dolado et al. (2018) for instance put an emphasis on two specific channels : (i) top income households happen to be high-skilled, and push their wages up as a consequence of a monetary expansion, since they benefit from smaller matching frictions in labor markets; and (ii) considering that these individuals present complementary features to capital, an increase in demand for the latter only magnifies income inequality in comparison to poor, low-skilled workers. On the empirical side, numerous country-level studies suggest that conventional monetary tightening increases income inequality (Coibion et al. (2017) for the U.S., Mumtaz and Theophilopoulou (2017) for the U.K., Park (2018) for South Korea, Furceri et al.(2018) for a selection of advanced and emerging economies). In contrast, recent research on the distributional effects of unconventional monetary policy shows that the relationship between monetary expansions and inequality is negative, though small in magnitude (see e.g. Casiraghi et al. (2018), Guerello (2018) and Frost and Saiki (2014),< among others).

The existing empirical literature on the relationship between monetary policy and income distribution mainly used data on inequality over a short period of time. This can be problematic since inequality measures are perceived to be sticky in the short and medium run. Addressing this issue with short-term data also implies giving coverage to fewer macroeconomic events (such as recessions, financial crises or sovereign defaults). Such events result in exceptional monetary policy shocks, which can have an important impact on inequality. In the meantime, analyzing the distributional effects of monetary policy from a historical perspective poses two major challenges. The first one is to mobilize reliable cross-country inequality data and long macroeconomic series. The second has to do with the identification of exogenous monetary policy innovations. Dealing with this point is particularly demanding, since the conduct of monetary policy in advanced economies experienced several changes throughout the 20th century. Such shifts relate for instance to the succession of different exchange rate regimes,

the occurrence of many banking crises and the usage of multiple frameworks in monetary policy decisions (i.e. inflation targeting, the Taylor rule, exchange rates management, etc.). Our paper aims at addressing these challenges using a different setting and a novel approach, providing new evidence on the distributional consequences of monetary policy.

This paper analyzes the relationship between monetary policy and income inequality between 1920 and 2015 using annual data across 12 advanced economies. Such historical analysis was conducted thanks to the combination of two datasets. We mobilize the World Inequality Database (WID), which offers an open access to historical series of income and wealth inequality. The share in national income held by the top 1 percent richest is used as the main inequality indicator. The top 1 percent richest receive a significant share of their total income in form of dividends and capital gains, while being almost untraceable from household income and wealth surveys. As an alternative inequality measure, we use in our analysis the top 10 percent share in national income, which is believed to gather well-off households with heterogeneous income sources. Long series of macroeconomic variables are extracted from the Jordà-Schularick-Taylor Macroeconomic Database, developed by Jordà et al. (2016). Using such historical macroeconomic data is of a great interest, because they offer a rich set of control variables that could enter as potential determinants of inequality. To the best of our knowledge, this is the first time WID and JST Macroeconomic Database are combined.

We employ two complementary empirical methodologies. The first consists in estimating a Panel VAR with 5 variables — including our inequality measure — to obtain the impulse response functions to an unexpected monetary policy shock. The second generates dynamic responses of inequality from local projections *à la* Jordà (2005). As noted by Barnichon and Brownlees (2018), both methodologies complement each other. The Panel VAR model yields consistent results when it is correctly specified, but becomes over-identified when using an important amount of endogenous variables. Local Projections (LP) are admittedly less efficient, yet they remain robust to model misspecification. More importantly, since it does not impose specific dynamics in the equation system, the LP approach allows more flexibility in dealing with the estimations. We exploit these features for our research question in two ways: (i) we construct an instrumental variable, which leans on the well-known macroeconomic policy

trilemma, in order to isolate exogenous fluctuations in the short-term interest rate and, (ii) since local projections accommodate with non-linearities, we test our model in a regime switching setting, where we allow for the response of inequality to depend on the regime of a specific variable (i.e. banking crisis, inflation regime and the output gap).

Our evidence suggest that monetary policy has a significant and considerable economic impact on income inequality. Monetary loosening increases the share in national income held by the top 1 percent richest, while a restrictive monetary policy leads the opposite effect. As far as results from the Panel VAR and OLS local projection are concerned, a 100 b.p. decrease in the short-term interest rate implies a peak increase of the top income index of about 1 and 1.15 percentage points. We demonstrate that the effect of monetary policy on top income households is arguably driven by higher stock prices, which is consistent with the income composition channel. Distributional effects of monetary policy in the instrumental variable setting are however more sizable. A perturbation to the domestic interest rate — via the external instrument — considerably widens inequality, with a peak increase of almost 6 percent percentage points. The results stand up to a battery of robustness checks to the baseline model.

These findings support the theoretical predictions of Dolado et al. (2018), where skill-distribution across households plays an important role in shaping the effect of monetary policy on inequality. Our evidence is also in line with the empirical findings of Romer and Romer (1999). Even though inequality is not the primary concern of central bankers, our results imply that it is a dimension that they should not overlook. This is especially true since income distribution may affect the transmission mechanisms of monetary policy.

The rest of the paper is laid out as follows: Section 2 discusses the estimation methodology and the identification strategy. Section 3 thoroughly describes the data. The fourth section presents the Panel VAR and local projections results, while the fifth and last section concludes.

2 Estimation approach

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

2.1 Panel VAR

Structural VAR are the traditional approach to identify structural monetary policy shocks, and simultaneously trace out the corresponding impulse responses of macroeconomic variables. We begin here with a traditional small monetary VAR amended to take into consideration the panel nature of our data. This model is extended to include an inequality indicator as well as a dummy for exogenous systemic banking crisis. Hence, the VAR contains five endogenous variables: the CPI, real GDP, the nominal short-term interest rate, the stock prices and the top one percent share in national income ($P1$). These series enter in the model, as the log changes for the CPI ($CPI_{i,t}$), real GDP ($GDP_{i,t}$), the stock prices ($S_{i,t}$) and the P1 ($P1_{i,t}$), while they enter in first-difference for the nominal short-term interest rate ($\Delta r_{i,t}$). Let $X_{i,t} = (\pi_{i,t}, y_{i,t}, r_{i,t}, P1_{i,t})$ be a vector of the five endogenous variables in the VAR. The reduced form of the model can be represented as follows:

$$X_{i,t} = \mu_i + \sum_{l=1}^L \beta_l X_{i,t-l} + \kappa D_{i,t} + \nu_{i,t} \quad (1)$$

where the indices t and i relate to years and countries; μ_i corresponds to the country fixed effects; L the lags number of endogenous variables included in the model, set at two according to the Akaike information criterion; B_i are $5 * 5$ matrices of unrestricted coefficients; $D_{i,t}$ is a banking crisis dummy that takes the value 1 when there is a systemic banking crisis and 0 otherwise; and finally $\nu_{i,t}$ a vector of unorthogonalized structural shocks.

In order to disentangle the causal chain of events and identify structural shocks of interest, we consider timing assumptions based on contemporaneous restrictions among the exogenous shocks in the VAR. Specifically, our approach departs from the recent literature on the macroeconomic effects of monetary policy. In fact, the latter mostly relied on Romer and Romer (2004) narrative approach to identify innovations to monetary policy. This is notably the approach adopted for instance by Coibion et al. (2017) and Furceri et al. (2018) to study the distributional effects of monetary policy. Nonetheless, such identi-

fication strategy is not tractable in our case since it requires (at least) forecasts of the short-term interest rates, inflation and GDP growth, which are not available over the long-run. On top of that, our interest for the effects of monetary policy on inequality, as well as the use of annual data for this purpose, reduce the problems raised by monetary policy shocks derived from a VAR. Indeed, it is less likely that the short-term interest rates move endogenously with changes in inequality than in output. Further, anticipatory movements in our monetary policy measure are less plausible with annual than quarterly data.

However, our identification strategy is not without problems. As a matter of fact, our simple recursive identification scheme generally delivers puzzling dynamic responses of inflation to monetary policy innovations (the so-called “price puzzle”). This counter-intuitive result questions the validity of the estimates of structural monetary policy shocks and, therefore, the accuracy of the derived inequality impulse response function. As a result, the first step consists in addressing the price puzzle.

Numerous solutions have been proposed in the literature, mostly based on the inclusion of additional variables into the VAR. Yet, these alternatives are not well-fit in our context, characterized by the use of historical data and the limitation of available time-series. In dealing with the “price puzzle, the proposal of Estrella (2015) is however very attractive because it makes the model more simple and less data demanding. Such proposal actually incorporates a theoretically motivated exclusion constraint in the VAR. The constraint is based on the insights of Friedman (1958, 1961) and the empirical observations of Bernanke et al. (1999) and Batini and Nelson (2001), which point to the existence of lags between monetary policy actions and their influence on prices. This leads to impose a single zero restriction on the coefficients matrix for the first lags of the short-term interest rate in the CPI equation. By doing so, besides the traditional short run restrictions — given by constraining the matrix A_0 to be lower triangular — the dynamic responses of key macroeconomic variables to monetary policy are no more puzzle.¹² As a result, the structural form of the estimated model is given by:

¹See Table A3 in the appendix for a comparison of the IRF.

²We test other specification of the model where we add another zero restriction to take account the fact that changes in monetary policy rates are likely to be independent from changes in inequality.

$$\begin{aligned}
& \overbrace{\begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ \alpha_0^{21} & 1 & 0 & 0 & 0 & 0 \\ \alpha_0^{31} & \alpha_0^{32} & 1 & 0 & 0 & 0 \\ \alpha_0^{41} & \alpha_0^{42} & \alpha_0^{43} & 1 & 0 & 0 \\ \alpha_0^{51} & \alpha_0^{52} & \alpha_0^{53} & \alpha_0^{54} & 1 & 0 \end{pmatrix}}^{A_0} \begin{pmatrix} CPI_{i,t} \\ GDP_{i,t} \\ \Delta r_{i,t} \\ S_{i,t} \\ P1_{i,t} \end{pmatrix} \\
& = \mu_i + \sum_{l=1}^L \begin{pmatrix} \alpha_l^{11} & \alpha_l^{12} & 0 & \alpha_l^{14} & \alpha_l^{15} \\ \alpha_l^{21} & \alpha_l^{22} & \alpha_l^{23} & \alpha_l^{24} & \alpha_l^{25} \\ \alpha_l^{31} & \alpha_l^{32} & \alpha_l^{33} & \alpha_l^{34} & \alpha_l^{35} \\ \alpha_l^{41} & \alpha_l^{42} & \alpha_l^{43} & \alpha_l^{44} & \alpha_{l,i}^{45} \\ \alpha_l^{51} & \alpha_l^{52} & \alpha_l^{53} & \alpha_l^{54} & \alpha_l^{55} \end{pmatrix} \begin{pmatrix} CPI_{i,t-l} \\ GDP_{i,t-l} \\ \Delta r_{i,t-l} \\ S_{i,t} \\ P1_{i,t-l} \end{pmatrix} + \kappa D_{i,t} + \varepsilon_{i,t} \quad (2)
\end{aligned}$$

where α_l are unrestricted structural parameters that are allowed to differ for each country and $\varepsilon_{i,t}$ is a vector of uncorrelated iid shocks. As noticed, the impact matrix A_0 is lower triangular. The variable ordering of this recursive identification scheme implies that inequality contemporaneously reacts to innovations in key macroeconomic variables.³

The Panel VAR model is estimated equation-by-equation by a fixed effect estimator. We acknowledge that this method has some flaws as it can yield biased estimates. In fact, demeaning in a dynamic panel model results in correlation between the error terms and regressors. However, as shown by Nickell (1981), the size of the fixed effects bias decreases as the length of the sample increases. Hence, the importance of this bias in our analysis is small because the time dimension is long and much longer than the country dimension⁴.

2.2 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 also estimating impulse

³At best, this ordering choice is questionable. Hence, we also adopt more conservative ordering by setting the inequality variable first in the VAR. The result are in this case very consistent and are available upon request.

⁴We also estimate our model by a mean group type estimator. The method consists in allowing all the coefficients to vary by country and using OLS to estimate the model. By doing so, we capture cross-sectional dynamic heterogeneity. The results of these estimates are reported in the appendix and are very consistent with our baseline results.

response functions (IRF) from local projections. In its basic form, local projection consists of a sequential regressions of the endogenous variable shifted several steps ahead. As a result, the approach consists in estimating the following equation:

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

where $\Delta_h y_{i,t-1} = y_{i,t+h} - y_{i,t-1}$ and corresponds to change in growth rate of the top income share from the base year $t-1$ up to year $t+h$ with $h = 0, 1, \dots, H$; $\Delta r_{i,t}$ denotes the change in short-term interest rate and $X_{i,t}$ refers to a vector containing the control variables. The control variables include the lags of $\Delta_h y_{i,t-1}$ and $\Delta r_{i,t}$ as well as a rich set of additional controls that theoretically explain inequality, and can be correlated with the monetary policy shock.

It is important to note that each step of the accumulated IRF is obtained from a different equation and that it directly corresponds to the estimates of β^k . This means that unlike the VAR approach, the estimated coefficients contained in θ^k are not used to build the IRF. They only serve as controls, and clean up the β^k from the effects of past inequality and monetary policy changes, besides contemporaneous and past changes of different macroeconomic variables (output and CPI, for instance). In other words, the LP approach does not impose any particular dynamics on the variables in the system. As shown by Jordà (2005), this 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 with non-linearities and (iv) can also be estimated with simple regression techniques. However, it has also some drawbacks in terms of efficiency. The VAR approach being more efficient when the model is well specified. Hence, the local projection method has complementarity features with the VAR approach to obtain impulse responses. In what follows, we describe the benefits of local projections with respect to our research question.

First of all, the local projections allow us to control our results for numerous factors that may influence inequality and be, at the same time, correlated with monetary policy actions. The $X_{i,t}$ vector includes numerous supplementary variables compared to the VAR: the stock price growth, the house price growth, the change in financial depth, the change in trade openness, the government expenditure growth and the U.S. patent number growth as a proxy of technological progress.

The second benefit of the local projection is that it offers an original identification strategy to estimate dynamic causal effects. To build shock series, our strategy relies on external instruments. Such variables are correlated with changes in the short-term interest rates, but not with the other macroeconomic shocks hitting the economy. We aim to obtain external sources of variation in the short-term interest rates to provide quasi-random experiments and, thereby, identify more clearly causal effects. This type of strategies, which borrow from microeconometrics, have recently attracted a growing interest in applied macroeconomics (Jordà et al., 2015, 2017; Jordà and Taylor, 2016; Ramey and Zubairy, 2018; Stock and Watson, 2018). To be more specific, it tackles the issue of joint determination of economic policies and macroeconomic outcomes that tends to bias the effects of economic policies. Regarding our research question, we can object that monetary policy is likely to be not driven by inequality and, therefore, that the dynamic causal effect is clear (no simultaneity bias). In such case, using external instruments would be worthless. However, even if inequality 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 variables bias) (Furceri et al., 2018). Accordingly, this calls upon usage of exogenous (to home economic conditions) monetary policy shocks rather than the short-term interest rates. As it is widely agreed in the literature, the major challenge — even for a macro issue like ours — is to find external factors that would make the monetary policy shock as a random treatment.

Here, 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). The purpose here is to use the macroeconomic policy trilemma to find external variations of monetary policy conditions. The macroeconomic trilemma states that a country cannot simultaneously achieve free capital mobility, a fixed exchange rate and an independent monetary policy. By pursuing any two of the goals, it is necessary to give up the third. Building on the trilemma framework (Obstfeld and Taylor, 1998, 2004 ; Shambaugh, 2004), we trace out episodes where external conditions can generate exogenous perturbations of the short-term interest rate. The latter

are considered to be unrelated because base country — as the U.S. in the Bretton Woods era — does not internalize the externalities of its own policy choices on partner countries. The trilemma links home interest rate with base country interest rate. A simple algebraic expression is given by:

$$\Delta r_{i,t} = a + b[PEG_{i,t} * KOPEN_{i,t} * \Delta r_{i,t}^{base}] + \Theta X_{i,t} + \mu_{i,t} \quad (4)$$

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

According to equation 4, variations in $\Delta r_{i,t}$ are related to external condition (the base country) when there is a 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 also an exogenous influence on local monetary policy conditions. It therefore provides a source of variation in short-term interest that is exogenous to the domestic conditions in term of inequality. As a result, $z_{i,t}$ constitutes a theoretically good external instrument. In what follows, like Jordà et al. (2015), we use $z_{i,t}$ as an IV for the change in interest rate to check in equation 3 the consistency of our baseline OLS estimates.

The third interest of the local projection is that it easily accommodates with non-linearities.⁵ This allows to enrich our analysis by checking if the impulse responses of inequality to a monetary policy innovation are state dependent. This is of great interest since we use historical data that cover different monetary policy regimes. This goes also in the direction of many studies, highlighting that the effects of monetary policy vary over the business cycle. In practice, we extend the equation 3 by conditioning the effects of the interest rates on inequality by a state variable:

$$\Delta_h y_{i,t-1} = \alpha_i^k + \beta^k \Delta r_{i,t} + \kappa^k \Delta r_{i,t} * State_{i,t} + \theta^k X_{i,t} + \vartheta^k State_{i,t} + \varepsilon_{i,t}^k \quad (5)$$

where $State_{i,t}$ is a dummy variable indicating the state or regime.

⁵The VAR literature also offer some solutions to deal with non-linearities. However, a richer structure of the VAR model goes with several complications to compute IR, which make often the estimation intractable in practice, if we are outside the baseline framework.

3 Data description

3.1 Inequality

The Gini coefficient has long been used to analyze income inequality, in that it illustrates the degree to which a variable is equally distributed across its population. However, the Gini index gives relatively greater weight for observations in the middle of the distribution, in contrast to those located at the tails. This flaw hampers to take into account aspects of concentration, which are at the very heart of the inequality issue. That is why a sound alternative would be to consider instead measures that exclusively focus on tails of the distribution. Such indicators take the shape of decile ratios or the shares in national income received by the 5, 1 or 0.1 percent people with the highest market incomes. As shown by Davtyan (2017), these measures strongly correlate with the Gini index and therefore do not alter the broad picture of inequality in an important way.

In this paper, top income data are extracted from the World Inequality Database (WID, 2017). Specifically, income inequality is approached by the top 1 percent pre-tax national income share (P1) in 12 advanced economies over the period 1920-2015. Countries include: Australia, Canada, Italy, Germany, Denmark, France, U.K, Japan, Netherlands, Norway, Sweden and finally the U.S. We also conduct our empirical analysis by separately excluding from the sample episode of WWII.⁶ As a robustness check, we test our model on the top 10 percent 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 individuals receiving important shares of capital income, the second contains more high income earners. Hence, P10 would be considered as more heterogeneous compared to P1, in that it gathers "rich" individuals who considerably differ in terms of their income sources. Figure A1 in appendix plots for each country the P1 and P10 over the studied period. The dynamics of top income shares could be characterized in three stages: an increasing top income shares on the eve of the two world wars before they got wiped out; a post war era where top marginal taxation contained the accumulation of private wealth; then the return of a patrimonial society in the beginning of the 1980s, which largely favored top income households.

⁶Table A1 traces out in detail data sources and their availability, for each country during both periods.

3.2 Macroeconomic variables

We exploit the Jordà-Schularick-Taylor Macrohistory Database, which provide us with long series of macroeconomic data⁷. In such database, information on several macroeconomic variables are available from 1870 to 2016 and cover around 17 developed economies⁸.

For the baseline Panel VAR framework, we mobilized the following macroeconomic aggregates: real GDP per capita (index, 2005=100), consumer prices (index, 1990=100) and the short-term interest rate. As aforementioned in section 3.1, to ensure stationarity of the series, real GDP per capita and consumer price index are taken in growth rate (log) while the first difference is taken for the short-term interest rate. In addressing the question of monetary policy and inequality, our paper also departs from the existing literature by building on several macroeconomic controls. The choice of these variables fits in with the literature on determinants of inequality. The set of control variables used for both local projections and the instrumental variable are summarized in Table A2 (see the appendix).

Our battery of control variables can be split into four categories: financial development, openness, government spending and technological progress. The way financial development shapes income inequality remains an open question in the literature. While it was widely believed that financial development would reduce inequality through a better access to credit for low-income households; the recent findings (Bazillier and Hericourt (2017) provide a complete survey on this question) argue on the contrary that more finance primarily 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. That is why to control for this factor, we add a real house price index. For what concerns globalization, trade openness might increase inequality through a "factor" channel. In fact, building on a simple Heckscher-Ohlin framework, we can assert that our country panel features advanced economies that are relatively well endowed in capital. As put by Roine

⁷For sake of clarity, all variables — including those on inequality — are winsorized in order to eliminate outliers.

⁸Our sample is only restricted because of the limited availability of income inequality data.

et al. (2009), inasmuch as "*capital owners coincide with the income rich*", trade openness is likely to benefit top income households and widen inequality.

Regarding government expenditures, they are approached in our analysis by the standard public debt-to-GDP ratio. From the outset, it is not easy to establish from such indicator the relationship between government spending and inequality. However, Dorn and Schinke (2018) show from a panel of 17 OECD countries, that government ideology seems to play a more determinant role. Precisely, between 1970 and 2014, top income shares increased more under right-wing governments than under left-wing ones. Hence, our measure of government spending would affect top income shares depending on the political stance of public finances. Besides, technological change has been repeatedly pointed out in the literature to be playing a potent role in widening wage inequalities. One way to control for this factor consists in mobilizing data on patents. Such data would aim at measuring the amount of inventions, some of which are likely to become marketable. To that end, we rely on a novel dataset, which tracks patent and grant activity in the U.S. since 1790. Specifically, we include data on utility patents applications, given that it primarily concerns "useful" inventions. The choice of focusing on the U.S. stems from the fact that it has usually been the country where most breakthrough innovations have emerged.

3.3 External instruments

As formerly discussed, the instrumental variable $z_{i,t}$ is the product of changes in base country short term interest rate ($\Delta r_{i,t}^{base}$), the exchange rate regime ($PEG_{i,t}$) and the degree of capital control ($KOPEN_{i,t}$). Following Jordà et al. (2015), pegs definition prior to WWII are extracted from Obstfeld et al. (2004, 2005). After WWII, data on exchange rate regimes are completed from Ilzetzki et al. (2017). Table A3 in 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 goes from 0 to 100) initially introduced by Chinn and Ito (2006). Just like Jordà et al. (2015), we use in our paper this index rescaled to the unit interval, with 0 meaning fully closed and 1 fully open. Figure 1 plots for our panel, changes in home interest rate $\Delta r_{i,t}$ against the constructed LP-IV.

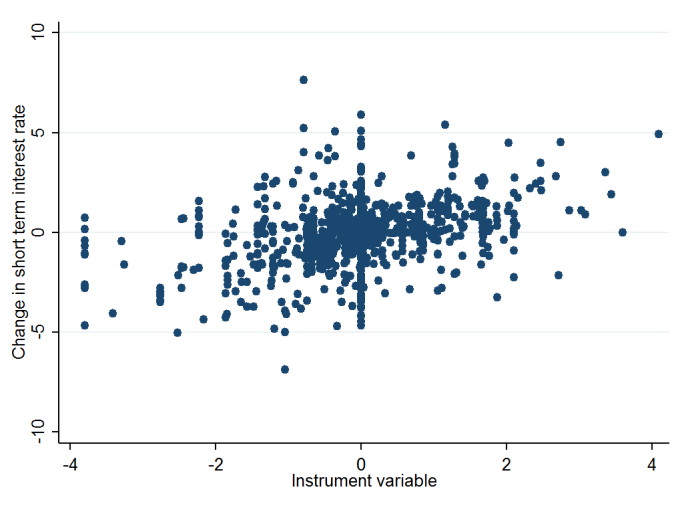


Figure 1: Jorda, Schularick and Taylor based IV: change in short-term interest rate in home and base countries

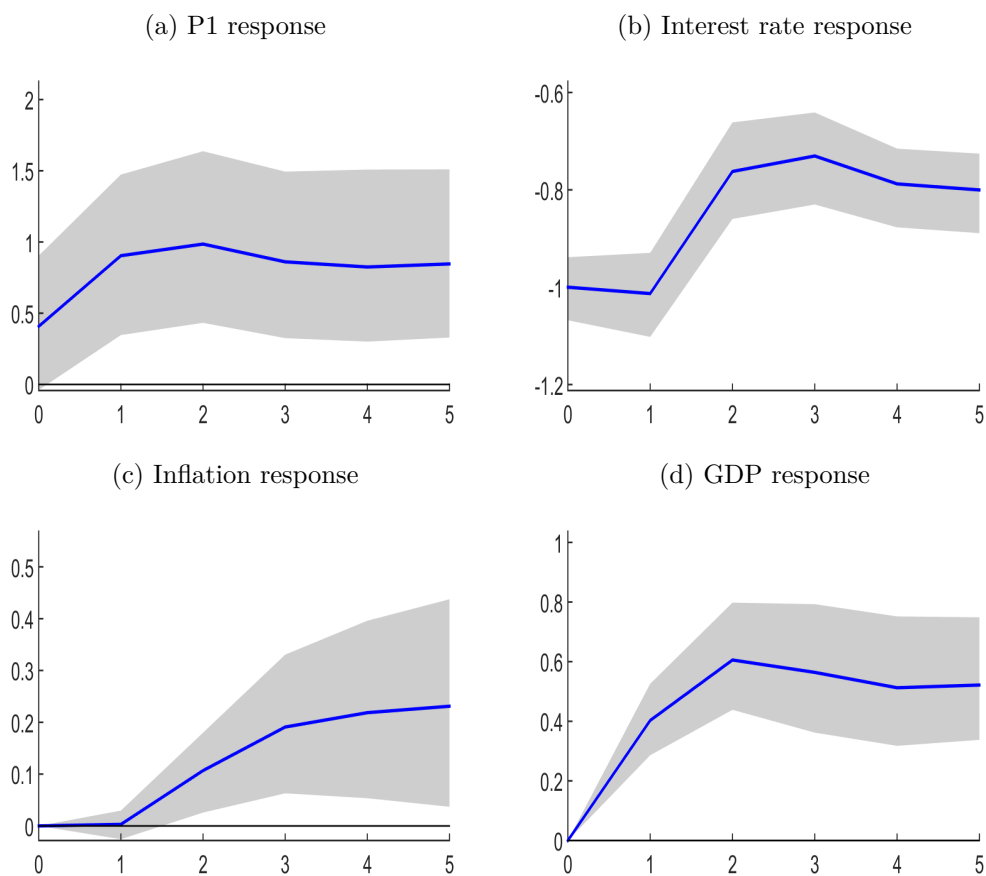
4 Results

This section reports the impulse response functions of inequality to a monetary policy shock both from vector autoregression and local projection.

4.1 Panel VAR results

The panel VAR described in the previous section is estimated by OLS and used to compute impulse response functions. Figure A3 displays impulse responses to orthogonalized 100 b.p. negative interest rate shocks in a 90% confidence band. As we have previously pointed out, one typical concern in the monetary VAR literature is the finding of price puzzle, i.e. the fact that an unexpected monetary tightening leads to an increase of inflation. Hence, a preliminary requirement is to check that the dynamic restriction that we imposed in our VAR following the proposal of Estrella (2015) is well suited to eliminate the price puzzle. The estimated shape of the inflation response to a monetary policy shock does not exhibit counter-intuitive effects. The dynamic effect of an interest shock on real GDP is also fairly standard. All of this comforts the choice of our identification structure and allows us to reasonably analyze the effects of our structural monetary policy shock on inequality.

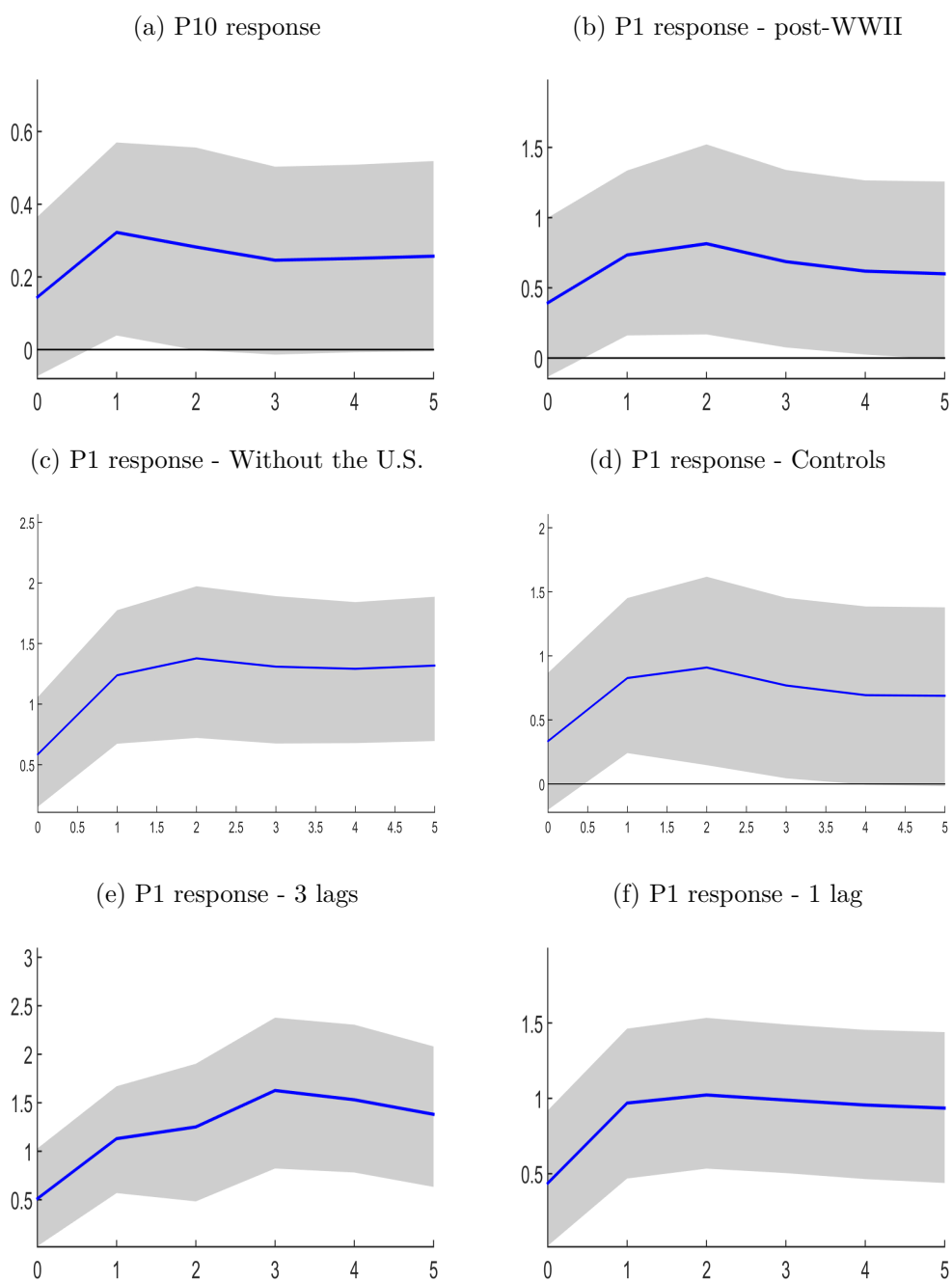
Figure 2: Inequality PVAR responses to a negative short-term interest rate shock: Cumulated effects



Note: The figure shows cumulated impulse responses of inequality to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

The impulse response function of the top 1 percent pre-tax national income share (P1) report in the top-left corner of the Figure A3 indicates that an unanticipated monetary policy shock leads to a rise in income inequality. This means that monetary policy loosening increases inequality. We can see that the monetary policy shock has significant and medium-term effects on inequality. A 100 b.p. decrease of the short-term interest rates increases the top income index by about 1% two year after the shock. Our results are in line with the recent theoretical predictions of Dolado et al. (2018) and the empirical results documented by Davtyan (2017). However, they contradict the empirical findings of Coibion et al. (2017) and Furceri et al. (2018).

Figure 3: Inequality PVAR responses to a negative short-term interest rate shock:
Robustness check



Note: The figure shows cumulated impulse responses of inequality to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

This obviously questions the findings sensitivity to our model and sample choices. Figure 3 displays the impulse responses to a variety of robustness checks. The first impulse response (graph (a)) reports results with our alternative measure of inequality. We find evidence of a positive reaction of the top 10 percent pre-tax national income share (P10) to an unexpected negative monetary policy shock. While the finding is consistent with our baseline result, we note that the reaction is less intense and less significant. This is in line with our expectations. Indeed, monetary policy shocks produce a more potent impact on the tails of income distribution. From this observation, we can infer that the response of the top 0.1 percent would be larger in comparison to our baseline results.

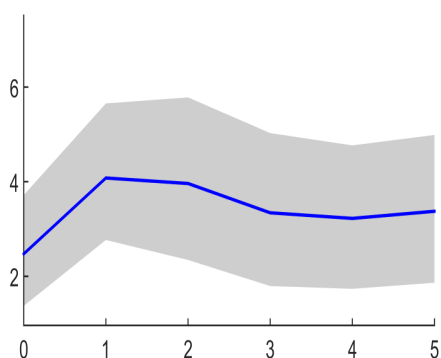
We next examine the effect of our particular sample choice in two ways. First, we exclude the pre-WWII period. By doing so, we check that our results are not driven by what it may be considered as an anachronistic monetary regime. This is also a way to be more closely aligned with the literature, which conducted empirical investigation with sample starting in the early 1980s. Second, we exclude the U.S. from our sample. There are indeed good reasons to believe that the U.S. are characterized by specific behaviors due to their historical international monetary role. This could create a heterogeneity bias because our empirical model controls for unobserved level heterogeneity — by demeaning the data — but not for unobserved dynamic heterogeneity. The results reported in Figure 3 panel (b) and (c) confirm our previous conclusions. On top of that, excluding pre-WWII period and the U.S. does not alter the response of inequality to an unanticipated negative monetary policy shock.

Finally, to assess the robustness of our results, we estimate different specifications of the PVAR. First, we include variables that reflect the dynamics of asset prices. This provides a more complete representation of the macro-level dynamics, in response to the fact that there are linkages between monetary policy, asset prices and inequality. In practice, we estimate a 6-dimensional Panel VAR model that enriches our baseline VAR with a measure of the house price cycle. The graph (d) on Figure 3 shows the consistency of our results to the extension of the PVAR model. Second, we check the sensitivity of our results to the number of lags included in the model. Evidence displayed in panel (e) and (f) show that the inequality response to a negative monetary policy shock remains unchanged with respect to the findings in our baseline model.

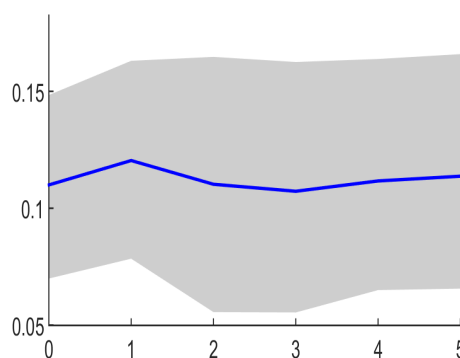
Our PVAR approach has also the advantage to give some insights about the channels through which monetary policy favors the rich. Our previous results could support the income composition channel. To empirically test the latter, we need to formally show that an unexpected negative change in interest rates positively affects asset prices, and in return, that an unexpected rise in asset prices increases the relative income of rich households. By doing so, we link the increase in inequality following a monetary loosening to the heterogeneity of income sources across households. This documents the fact that: the richer the households are, the larger are their income shares from financial markets.

Figure 4: Inequality PVAR responses to a negative short-term interest rate shock: Insights of the financial channel

(a) Monetary policy shock - Stock prices response



(b) Stock prices shock - Inequality response



Note: The figure shows cumulated impulse responses of stock prices and inequality to an unexpected decrease of 100 b.p., of the short-term interest rate and unexpected increase of 100 b.p of the stock prices, respectively. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

The Figure 4 reports the impulse response function of the stock prices to an unexpected monetary policy as well as the response of inequality to an unexpected stock price shock. The former shows that an unexpected decrease of the short-term interest rate increases stock prices, i.e. the expected future income, while the latter indicates that a negative shock on stock prices increases income of the top 1 percent. Therefore, this shows that the effect of monetary policy on the top 1 percent richest is likely to be channeled through higher asset prices.

4.2 Local projection results

Our Panel VAR approach is supplemented by a local projections estimation along with a novel identification of monetary policy shocks. A first step would be to first assess the strength of our instrumental variable. To do so, we estimate in the context of equation 3, a first-stage regression of the short term interest rate on the instrument $z_{i,t}$ and the aforementioned macroeconomic controls, including country-fixed effects. First-stage regression results are reported in Table 3 and underline the soundness of our instrumental variable. The coefficient estimates of the instrument $z_{i,t}$ remain statistically significant and they range from year 0 (when the shock is felt) to year 4 between 0.45 and 0.48, while F statistics feature high values across samples. That being said, we can now proceed to analyze the inequality local projection responses to monetary policy shocks.

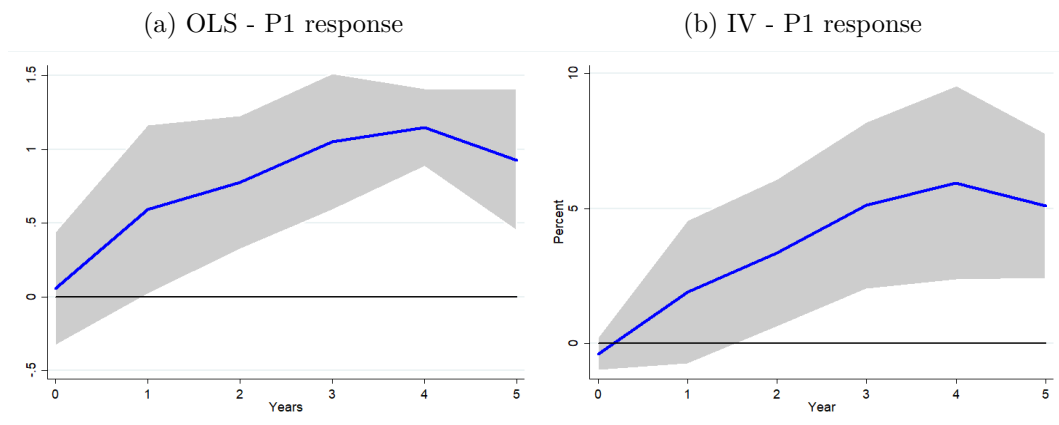
Table 1: Local projection-IV: First-stage results

Δ <i>Short-term interest rate</i>	Year 0	Year 1	Year 2	Year 3	Year 4
IV	0.47*** (0.06)	0.47*** (0.06)	0.48*** (0.06)	0.46*** (0.06)	0.45*** (0.06)
Observations	659	646	633	619	604
F	60	60.84	61.34	55.05	52.46

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

The results obtained from the estimation of equation 3 by local projections are presented in Figure 5. The two graphs illustrate the impulse response functions of inequality to an unexpected negative monetary policy shock with the associated confidence bands, both using OLS and the instrumental variable. A first look on the IRs seems to confirm what has been documented in the previous section, that is, monetary easing significantly and durably increases income inequality. Precisely, an unanticipated decrease of 100 b.p. in the short-term interest rate (graph (a) on the right) increases the top 1% share by about 1.15 percent four years after the shock. This impact on inequality is quite similar to what our Panel VAR has produced. Nonetheless, the effects on inequality are interestingly more pronounced when we use instead the instrumental variable. Indeed, a perturbation to the home interest rate $r_{i,t}$ via the instrument $z_{i,t}$ (graph (b)) increases the P1 by 3.35, 5.11 and 5.95 percent respectively two, three and four years following the shock.

Figure 5: Inequality local projection responses to a negative short-term interest rate shock



Note: The figure shows cumulated impulse responses of inequality to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

Such differences are more perceptible in Table 2, which jointly reports coefficient estimates of OLS and the LP-IV. Comparing the results obtained by both methods stems from the interest of assessing the degree of attenuation bias in the OLS estimation. In doing so, one would notice that impulse responses of both methods feature a relatively common scheme. Yet, coefficient estimates of OLS are moderately lower in comparison to IV, although signs from year 1 to year 5 are correct. For illustration, a monetary policy shock reduces P1 in year 3 after the shock by 1.05% using OLS and by 5.11% with IV. In the meantime, it is fair to emphasize that Jordà et al. (2015) — where they investigate the effect of monetary policy on house prices in the very long-run — document the same observation in a more or less similar magnitude. How to account for such contrast between OLS and the LP-IV coefficient estimates? The answer essentially lies in the way monetary policy shocks are captured by the instrumental variable. To be more specific, the way the instrument $z_{i,t}$ is built makes it more likely to be driven by exogenous perturbations in the world economy, in comparison to fluctuations captured by a simple OLS regression. As a result, home interest rate in the IV estimation is highly sensitive to changes in external financial and monetary conditions.

The reliability of local projection results also requires conducting a variety of alternative estimations. As noticed in Figure 6, we replicate for the OLS and LP-IV estimation methods the same analysis conducted in the Panel VAR model.

The robustness checks begin with our alternative measure of inequality i.e. P10.

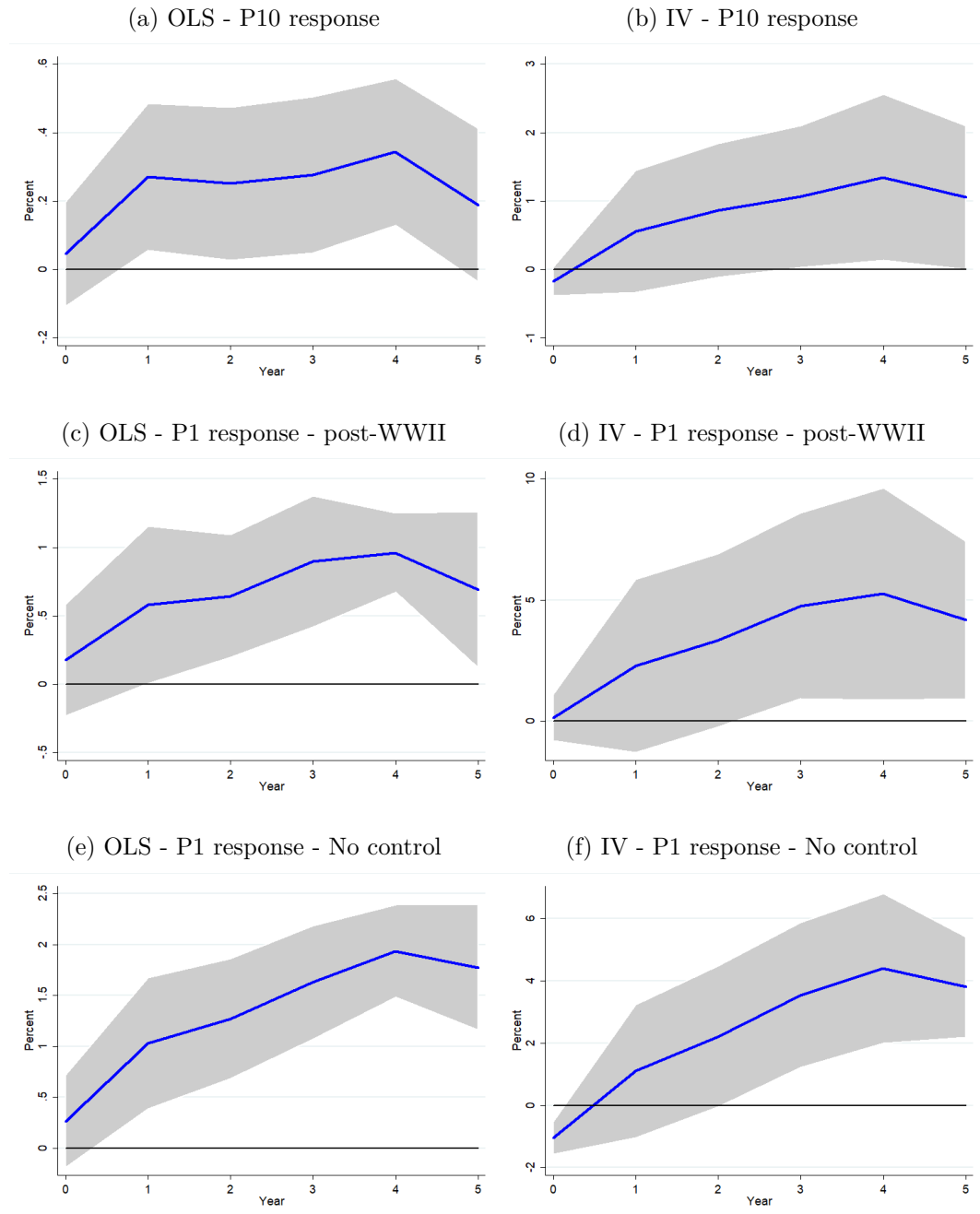
Table 2: Local projection: OLS and IV estimation results

OLS estimates - P1	Year 0	Year 1	Year 2	Year 3	Year 4	Year 5
Δ <i>Short-term interest rate</i>	-0.06 (0.23)	-0.59* (0.35)	-0.78** (0.28)	-1.05*** (0.28)	-1.15*** (0.16)	-0.928*** (0.29)
R^2	0.205	0.195	0.16	0.154	0.15	0.15
Observations	659	646	633	619	604	591
IV estimates - P1						
Δ <i>Short-term interest rate</i>	0.39 (0.38)	-1.91 (1.62)	-3.35** (1.67)	-5.11*** (1.89)	-5.95*** (2.19)	-5.10*** (1.86)
R^2	0.201	0.170	0.095	0.034	0.08	0.06
Kleibergen-Paap	6.66	6.64	6.85	6.71	6.43	48.04
Observations	659	646	633	619	604	591

Note: Country-based cluster-robust standard errors are reported in parentheses below their coefficient estimates. Coefficient estimates of controls and fixed effects are not reported. The controls includes the two lags terms of (i) the change in short-term interest rate; (ii) the change in inequality; and the contemporaneous and two lags terms of (iii) the real per capita GDP growth; (iv) the CPI inflation rate; (v) the stock price growth; (vi) the real per capita U.S. GDP growth; (vii) the level of financial development; (viii) the level of commercial openness ; (ix) house prices growth; (x) government spending; and (xi) patent activity. We report the Kleibergen and Paap (2006) statistic for weak instruments. *, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

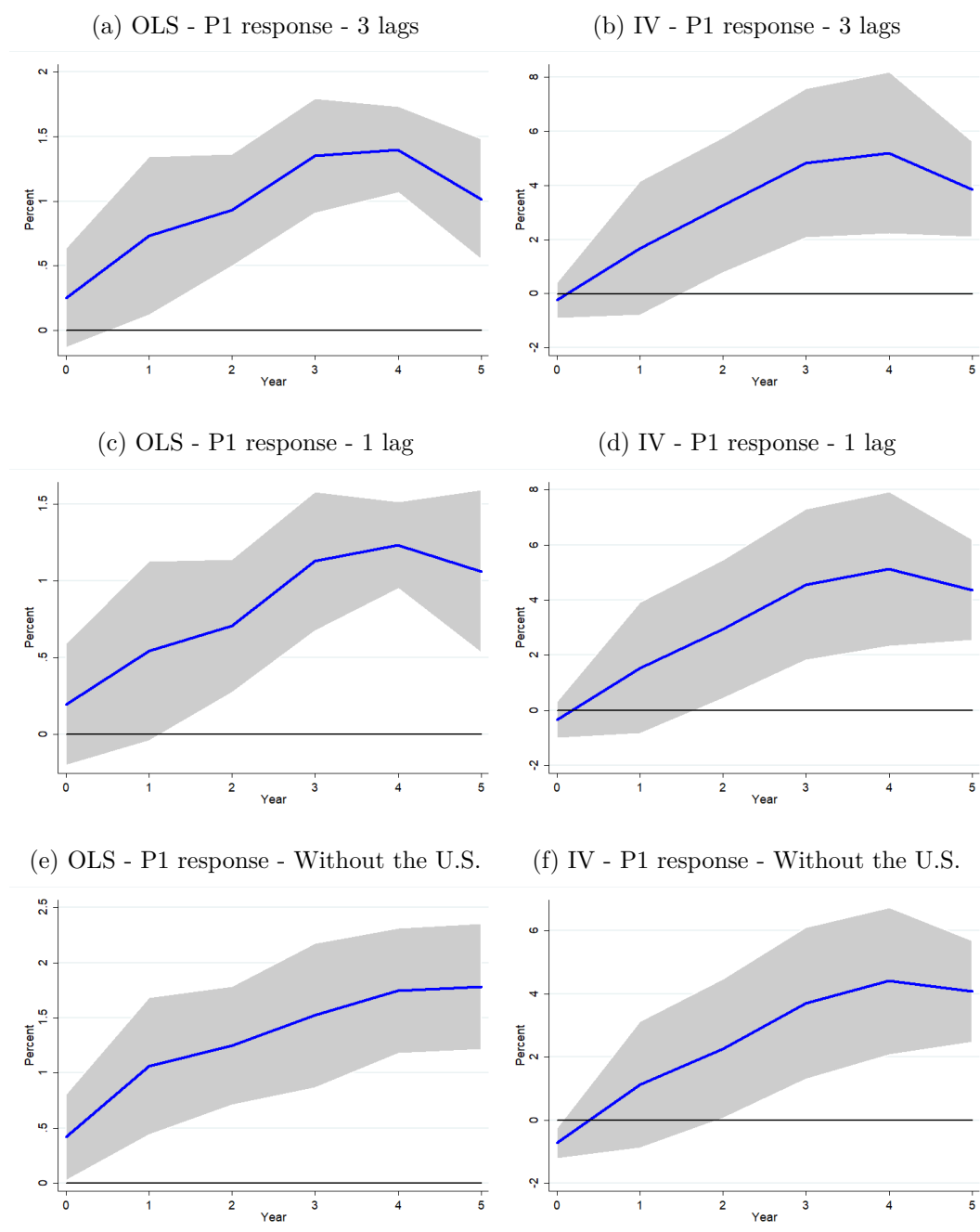
Impulse responses documented in graph (a) and (b), affirm as before that the responses of P10 to an unexpected monetary policy shock are lower by comparison to P1. Two more observations regarding the responses of P10 are worth discussing: (i) coefficient estimates from OLS remain smaller with respect to their IV counterparts; and (ii) OLS estimates turn statistically significant earlier (as of year 1 in contrast to year 3 for IV). The same comment could be put forward with regard to the estimation of local projections starting from the post-WWII period (see graph (b) and (c)). Also, coefficient estimates after the war are very close to those obtained in the benchmark LP estimation, both for OLS and the instrument. Therefore, this does not contradict our main findings and continues to confirm the stability of our estimates. An additional robustness check consists of estimating equation 3 with country-fixed effects while omitting the rich set of control variables. Such exercise is interesting to assess whether the IV exclusion restrictions are not violated. As a matter of fact, a correctly specified instrument would be enough to avoid a potential endogeneity bias. Evidence shown in graph (e) and (f) do not go against our main results. This sheds light on the reliability of the IV and OLS regression in the context of local projections framework.

Figure 6: Inequality local projection responses to a negative short-term interest rate shock: Robustness check



Note: The figure shows cumulated impulse responses of inequality to to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

Figure 7: Inequality local projection responses to a negative short-term interest rate shock: Robustness check



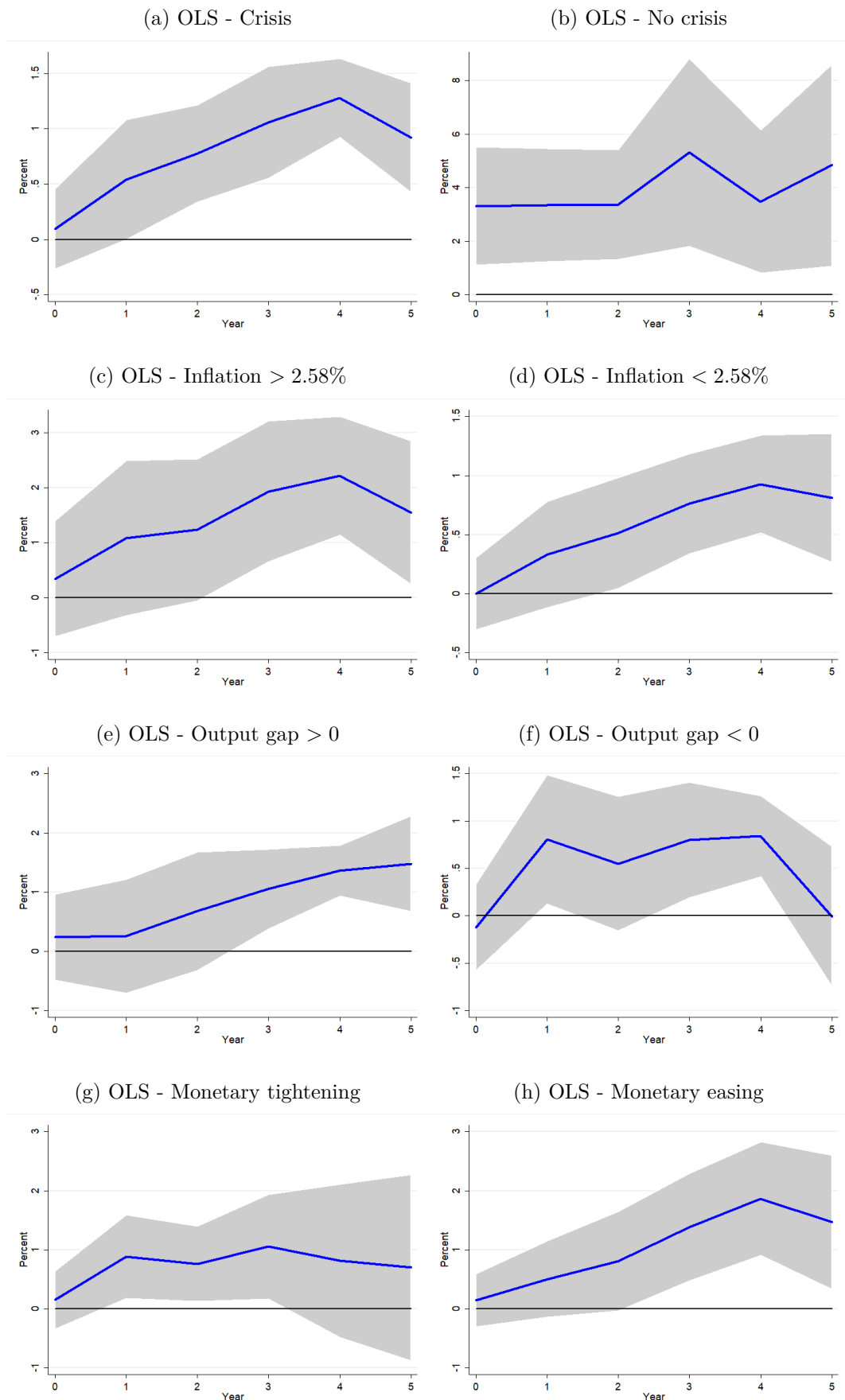
Note: The figure shows cumulated impulse responses of inequality to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

In a second step of our sensitivity analysis, we test the baseline model to different lag numbers, for both estimation methods. Figure 7 exhibits the additional modifications brought to the empirical analysis. To begin with, graphs (a), (b), (c) and (d) point that the LP framework stays robust to different lags and does not depart from results documented by the Panel VAR. Walking through the corresponding impulse responses, we notice that signs do not change, while statistical significance and coefficient estimates remain stable. Finally, we assess whether the effects of monetary policy on inequality are robust when excluding the U.S. from the sample. Once again, results in graph (e) and (f) are very similar to those obtained in the baseline; except that the response of P1 with OLS turns statistically significant immediately in year 0. This further validates that monetary policy shocks are well identified in our empirical analysis.

4.3 Regime switching

The results we have reported in the previous sections suggest that a monetary policy easing increases inequality unconditionally to the state of the economy. There is, however, a potential pitfall because our sample encompasses very different economic regimes. On top of that, several studies point that some economic variables, such as the short-term interest rate, may for instance behave very differently during economic downturns. To overcome this limit, we take advantage of the fact that LP method easily accommodates with non-linearities. Therefore, it is convenient to develop a regime-switching version of the previous model. That means we allow for the impact of monetary policy on inequality to depend upon the regime of another variable (see equation 5). By this way, we can compute conditional impulse responses at a particular regime. In this context, we consider three factors potentially making different the impulse responses of monetary policy: the occurrence of a systemic banking crisis, the inflation regime and the state of the economy over the business cycle. For each of these variables, we define a binary variable taking the value of one when there is a banking crisis or the inflation rate exceeds the median of its sample distribution (2.58%) and 0 otherwise. Also, the same is done when the output gap — computed with a typical HP-filter — is positive.

Figure 8: Inequality local projection responses to a negative short-term interest rate shock: Regime switching



Note: The figure shows cumulated impulse responses of inequality to an unexpected decrease of 100 b.p. of the short-term interest rate.

Figure 8 reports the impulse responses from the estimate of the regime switching model (equation 5) for the 3 factors previously described. Overall, the impulse responses displayed do not conflict with the previous ones. We find that monetary policy loosening has considerably bigger effects outside banking crisis periods and when inflation is high. 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*". The effects of monetary policy also differ according the state of the economy. We find that a monetary policy shock has more immediate effects on inequality during recessions than during expansions. However, there is no a significant difference regarding its effect on the medium-run. Interestingly, the same empirical strategy could be used to check for asymmetries between expansionary (negative variations of interest rates) and contractionary (positive variations of interest rates) monetary policies. The impulse responses reported in graph (g) and (h) of Figure 8 document that both monetary policies do not present striking differences in terms of impact on inequality.

5 Conclusion

This paper sought to investigate the distributional consequences of monetary policy between 1920 and 2015. The central idea that guided this paper's argument is that the existing literature feature the distributional effects of monetary policy using data on inequality over a short period of time. However, inequalities tend to vary more in the medium-long run. We address this shortcoming by studying how changes in monetary policy stance over a century impacted income distribution, while controlling for the determinants of inequality. To do so, we combined two large datasets : (i) the World Inequality Database (WID) to extract data on the share in national income held by the top 1% richest; and the Jordà-Schularick-Taylor Macrohistory Database, which allow to access large series of macroeconomic and financial variables.

Our empirical strategy considers two complementary approaches that conciliate between consistency and flexibility. The first approach deals with a Panel VAR framework, which delivers the orthogonalized impulse response function of inequality to a monetary policy shock. The second generates dynamic responses

of inequality in a Local Projections (LP) setting, wherein a rich set of macroeconomic controls is included. This framework allows to have a natural experiment set up, where exogenous perturbations in monetary policy are driven by factors unrelated to home economic conditions. Such exogenous perturbations enter as an instrumental variable, which traces out impulse responses of inequality.

The results obtained from both empirical methods indicate that *loose monetary conditions strongly increase income inequality and vice-versa*. In fact, following an expansionary monetary policy shock, the share in national income held by the top 1 percent richest increases by around 1% to 6%, according to estimates from the Panel VAR and Local Projections (LP). This effect is statistically significant in the medium-run and economically considerable. We also demonstrate that the increase in top 1 percent share mainly works through higher asset prices. The baseline results stand up to a battery of robustness checks, which: (i) consider an alternative inequality measure, (ii) exclude the U.S. economy from the sample, (iii) specifically focus on the post-WWII period, (iv) remove control variables and (v) test different lag numbers. On top of that, the regime-switching 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 inequality measures, which exclusively focus on "rich" households (i.e. top 5% or top 0.1% with the highest market incomes). Similarly, it would be interesting to confirm if the obtained results hold as well for wealth inequality. This aspect is important as wealth is more unevenly distributed than incomes. Also, given that we use pre-tax data, policymakers may be interested in the effects of monetary policy on inequality, net of the contribution of fiscal policy. Finally, the empirical approach adopted in this paper only considers — in the same spirit of the corresponding literature — the global effects of monetary policy on income distribution. That is, we do not identify all the transmission channels through which distributional effects of monetary policy work. That said, what are the policy implications we can draw from these findings for the ongoing debate on monetary policy and inequality. Central bankers need to be attentive not only to the aggregate consequences of monetary policy, but also to their side effects. Research agenda on optimal monetary policy should also be reconsidered, in order to better take into account distributional issues.

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Appendix

Table A1: Data sources and periods of inequality measures

Country	Period P1	Period P2	Details
Australia	1921-1938	1946-2015	WID (2017)
Canada	1920-1938	1946-2010	WID (2017)
Germany	1925-1938	2001-2013	WID (2017)
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.

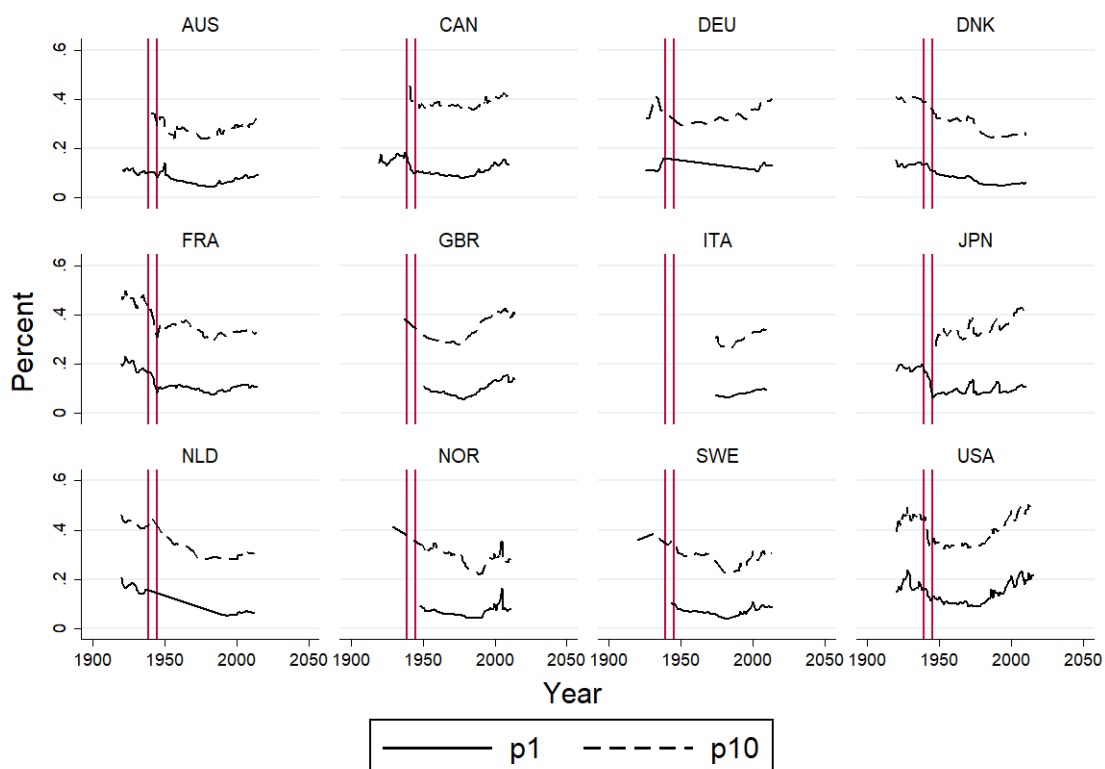


Figure A1: Inequality over time: 12 countries

Table A2: Control variables definition

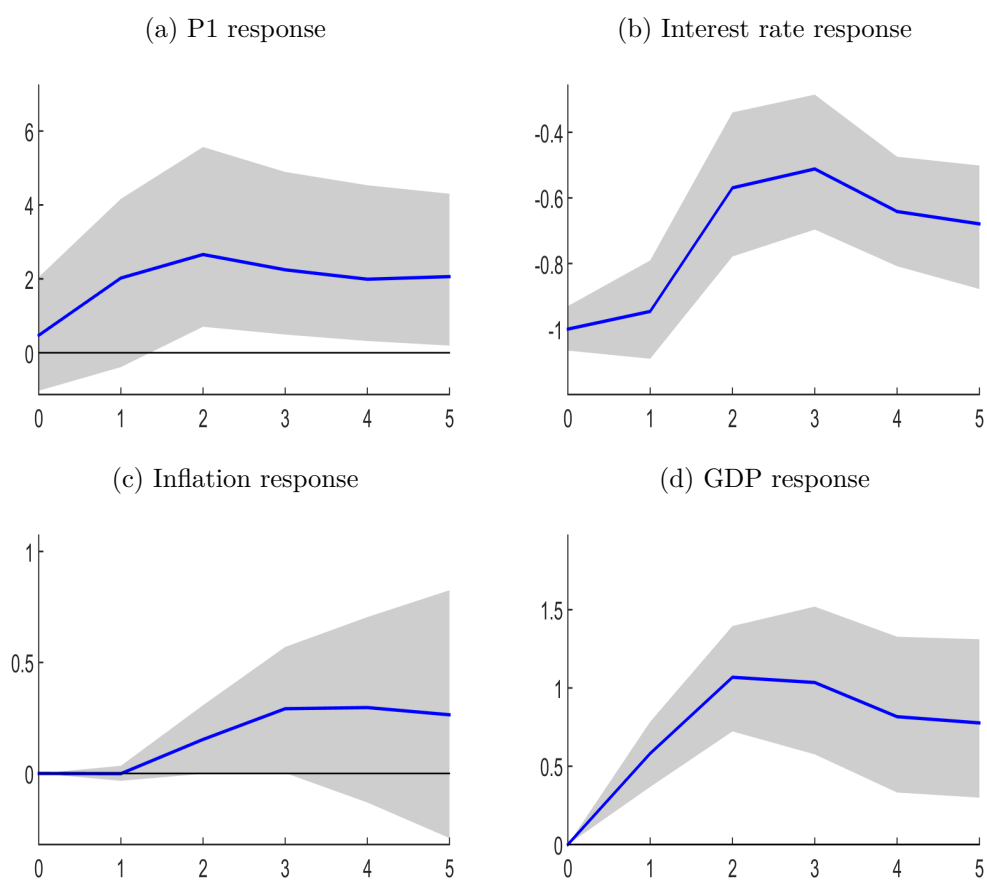
Variable	Variable definition	Source
Hpnom	House prices growth (real index, 1990=100)	Macrohistory Database JST
Stocks	Stock prices index growth (real index)	Macrohistory Database JST
Tloans	Ratio of total loans to non-financial private sector to GDP	Macrohistory Database JST, own calculations
Com_open	Ratio of imports and exports to GDP	Macrohistory Database JST, own calculations
Debt_gdp	Public debt-to-GDP ratio (in log-level)	Macrohistory Database JST
US_gdp	GDP growth of the U.S.	Macrohistory Database JST
US_patents	Growth rate of utility patents applications	United States Patent and Trademark Office

Note: Real indexes are obtained by dividing the variables on CPI and growth rates are computed taking the log.

Table A3: Exchange rate regimes

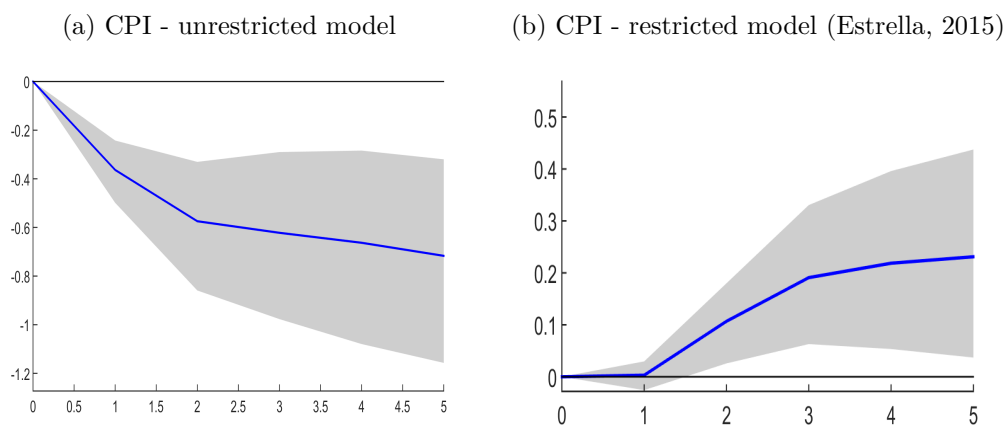
Country	Fixed	Floating
Australia	1920-1938, 1946-1983	1939-1945, 1984-2015
Canada	1920-1938, 1946-2015	1939-1945
Germany	1920-1938, 1946-1972, 1999-2014	1939-1945, 1973-1998
Denmark	1920-1938, 1946-2014	1939-1945
France	1920-1938, 1949-2014	1939-1948
U.K.	1920-1938, 1946-2008	1939-1945, 2009-2015
Italy	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 A3: Inequality PVAR responses to a short-term interest rate shock: Cumulated effects estimated from a mean group type estimator



Note: The figure shows cumulated impulse responses of inequality to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).

Figure A3: Comparison of CPI Impulse response function



Note: The figure shows cumulated impulse responses of CPI to an unexpected decrease of 100 b.p. of the short-term interest rate. The coloured bands represent 90% confidence bands generated by bootstrapping (1000 draws).