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Revisiting the Macroeconomic Effects of Monetary Policy Shocks*

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*This study re-evaluates the macroeconomic impact of monetary policy shocks in the United States using the proxy-SVAR approach of Gertler and Karadi (*American Economic Journal: Macroeconomics*, 2015, 7, 44–76). Despite increased credit costs and the presence of the credit channel, our analysis reveals modest effects on economic activity post-mid-1980s. This robust finding holds across various inference methods and sample considerations. A counterfactual analysis within the framework of Bernanke, Gertler, and Gilchrist (1999,* Handbook of Macroeconomics*, Elsevier) suggests that a stronger response to inflation in monetary policy execution may explain the subdued impact of shocks, offering valuable insights for policy refinement and understanding the dynamics of economic activity.*

I Introduction

Since the groundbreaking contribution of Sims (1980), vector autoregressions (VARs) have become a prominent modelling tool for examining the impacts of structural shocks. Specifically, a considerable body of literature has delved into the analysis of the effects of monetary policy shocks on economic activity using VARs. For an early survey of this literature, refer to Christiano *et al*. (1999). A substantial amount of empirical evidence supports the notion that an unexpected tightening of monetary policy leads to reductions in output, consumption and investment. This conclusion is reinforced by numerous studies, including those by Bernanke and Blinder (1992), Christiano *et al*. (1996), Leeper *et al*. (1996), Leeper and Zha (2003), and Romer and Romer (2004), among others. These empirically established patterns serve

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as fundamental pillars for formulating and evaluating the predictions of structural dynamic stochastic general equilibrium models in the existing literature.

Yet, the post-2000 literature on VAR-based characterisation of monetary policy shocks in the US presents a nuanced picture with mixed evidence. Boivin and Giannoni (2002, 2006) use a VAR with a recursive identification scheme and find that the response of output during the post-Volcker period was only approximately one-quarter of that during the pre-Volcker period. Similarly, Galí and Gambetti (2009) observe a decline in the response of inflation and real activity to demand-type shocks over time, although they do not separately identify a policy shock. In contrast, Primiceri (2005) utilises a time-varying parameter VAR (TVP-VAR) with stochastic volatility to characterise the changing transmission of monetary policy. Unlike Galí and Gambetti (2009), Primiceri (2005) finds little change in the dynamics of monetary policy shocks in the post-war period. Meanwhile, Canova and Gambetti (2009), using the TVP-VAR framework alongside sign restrictions, demonstrate that the transmission of monetary policy shocks has been relatively stable over time. They note somewhat stronger effects on inflation and real activity in the post-1990 period.

Highlighting the potential impact of identification methods, Boivin *et al*. (2010) suggest that discrepancies in findings could arise from the way monetary policy shocks are identified. For example, Canova and Gambetti (2009) leave the impact response of real activity unrestricted, while Boivin and Giannoni (2002, 2006) constrain it to be zero under the recursive identification scheme. Utilising a factor-augmented VAR that accounts for potential omission of important variables, Boivin *et al*. (2010) find that monetary policy innovations have muted effects on real activity and inflation in the post-1980 period. This conclusion contrasts with more recent studies employing proxy SVARs, which show that monetary policy shocks induce persistent declines in real economic activity (Gertler & Karadi, 2015; Caldara & Herbst, 2019; Jarociński & Karadi, 2020; Miranda-Agrippino & Ricco, 2021).

The mixed findings from the literature reveal the sensitivity of VAR-based characterisation of monetary policy shocks to the chosen framework, identification strategy and sample period. Ramey (2016) underscores this by highlighting mixed evidence on the importance of monetary policy shocks during the Great Moderation, asserting that existing estimates lack robustness. In light of these varied outcomes, our paper reassesses the macroeconomic effects of monetary policy shocks using US data. Our identification approach uses a proxy-VAR methodology, using high-frequency surprises in financial markets around tight windows of monetary policy announcements as instruments. This identification strategy was previously employed by Stock and Watson (2012), Mertens and Ravn (2013, 2014), Mertens and Montiel Olea (2018), and Angelini *et al*. (2023), among others. Our choice is motivated by Gertler and Karadi (2015), who advocate the inclusion of financial variables to capture the credit channel of the monetary policy transmission mechanism. Gertler and Karadi (2015) emphasise the need to identify policy surprises that are exogenous to both economic and financial variables in the VAR. They argue against the standard identification strategy based on timing restrictions when financial variables are included, as it does not address the simultaneity issues that may arise. The proxy SVAR framework, using high-frequency measures of policy surprises as external instruments, overcomes simultaneity concerns. Gertler and Karadi (2015) demonstrate that such surprises reveal substantial movements in credit costs, significantly impacting economic activity. These effects hinge on the strong systematic response of monetary policy to financial conditions, a factor often overlooked in studies emphasising modest macroeconomic effects of monetary policy shocks, as highlighted by Caldara and Herbst (2019).

We employ a baseline VAR with monthly data, including six variables: industrial production (IP), consumer price index (CPI), the 1-year government bond rate, excess bond premium, mortgage spread and commercial paper spread, following Gertler and Karadi (2015). The 1-year government bond rate serves as the policy indicator, with innovations capturing surprise changes in the current and expected future path of the funds rate. Our external instrument utilises surprises in futures rates around Federal Open Market Committee (FOMC) policy announcements. In contrast to Gertler and Karadi (2015), our findings suggest that monetary policy shocks have modest real effects post-mid-1980s. We attribute this discrepancy to data from the 'Volcker disinflation' period. While including this episode leads to a statistically significant decline in output following a monetary policy shock (as in Gertler & Karadi, 2015), excluding it results in muted real effects in the post-mid-1980s.

The Volcker disinflation episode, marked by extreme interest rate swings, inflation declines and output losses, likely involves measurement errors in assessing policy innovations. Our findings align with Coibion (2012), who, using Romer and Romer's (2004) approach, found that excluding the non-borrowed reserve targeting period during Volcker's disinflation significantly reduces estimated effects of monetary policy shocks. Despite incorporating the credit channel of monetary policy transmission, a well-established amplifying mechanism (Bernanke *et al.*, 1999; Caldara & Herbst, 2019), we consistently observe modest real effects. Moreover, our proxy-SVAR approach addresses concerns raised by Castelnuovo (2016), who showed that timing restrictions in recursive identification can distort estimated policy shocks. Our results underscore that modest macroeconomic effects of monetary policy shocks are a distinctive empirical feature of the post-1980 US economy, which aligns with Ramey (2016)'s findings, and suggest that monetary policy shocks may not have significantly driven macroeconomic fluctuations during the Great Moderation.[1]

A growing body of literature, including studies like Barth and Ramey (2001), Boivin and Giannoni (2002, 2006), Boivin *et al.* (2010), Castelnuovo (2016), Gertler and Lown (1999), Hanson (2008), Mojon (2008), and Ramey (2016), highlights modest macroeconomic effects of monetary policy shocks during the Great Moderation. Scholars, such as Boivin and Giannoni (2006), Boivin *et al.* (2010), and Castelnuovo (2016), propose that changes in consumer and firm behaviour, possibly due to technological progress or financial innovations, enable better adaptation to interest rate fluctuations. However, the economic response to a monetary policy shock is not solely contingent on the behaviour of households and firms but also on how monetary policy is conducted. As emphasised by Boivin and Giannoni (2006), if monetary policy is effective at influencing output and inflation, and the central bank is able to perfectly offset the effects of exogenous disturbances on these variables, then the estimated impulse response functions with respect to monetary policy shock should display no response of inflation and output. Building on Boivin and Giannoni (2006), we document that modest macroeconomic effects of monetary policy shocks are due to improved policy conduct. Using the financial accelerator model of Bernanke *et al.* (1999), we illustrate that a stronger response to inflation mitigates the impact of monetary policy shocks on economic activity and inflation, despite the operational credit channel of monetary policy. Consequently, changes in the dynamics of monetary policy shocks, as suggested by Boivin and Giannoni (2006), may not necessarily signify reduced effectiveness but rather an improvement in the policy conduct.[2]

The rest of the paper is organised as follows. Section II presents the SVAR-IV framework and explains how external instruments can be used to identify monetary policy shocks, provides a brief overview of the monetary policy transmission mechanism and presents the data used in the

[1] Our finding remains robust across multiple dimensions. We explore alternative measures of economic activity, policy indicators, external instruments and various sample periods, consistently obtaining similar results. In addition, our analysis incorporates alternative inference methods, addressing challenges associated with weak identification.

[2] See Clarida *et al.* (2000), Lubik and Schorfheide (2004), Haque *et al.* (2021), Haque (2022) for evidence of better monetary policy in the post-1980s.

estimation. Section III documents the results using both the proxy SVAR approach of Gertler and Karadi (2015) and recently developed alternative inference methods including weak identification-robust ones of Montiel Olea *et al*. (2021). Section IV provides a theoretical explanation for the empirical findings in the paper. Additional robustness checks are performed in Section V. Section VI concludes. Further empirical results are available in the Appendix S1.

II Econometric Framework

(i) Proxy SVAR Model

We consider the following proxy SVAR setting of Gertler and Karadi (2015), where Y_t is an $n \times 1$ vector of economic and financial variables and follows a stationary p-order structural VAR with underlying reduced-form representation:

$$Y_t = \sum_{j=1}^{p} B_j Y_{t-j} + u_t, \tag{1}$$

where u_t is an $n \times 1$ vector of reduced-form innovations, and B_j's are $n \times n$ matrices of unknown coefficients. The reduced-form innovations u_t are related to the structural shocks ε_t via:

$$u_t = A_0 \varepsilon_t, \tag{2}$$

where A_0 is an $n \times n$ non-singular matrix and the structural shocks ε_t are assumed to be serially and mutually uncorrelated, with

$$E(\varepsilon_t) = 0 \text{ and } E\left(\varepsilon_t \varepsilon_t'\right) = D = \text{diag}\left(\sigma_1^2, \ldots, \sigma_n^2\right). \tag{3}$$

From (3), the variance–covariance matrix of the reduced form innovations is given by

$$E\left(u_t u_t'\right) := \Sigma = A_0 D A_0'. \tag{4}$$

Owing to the stationarity assumption of the underlying reduced-form VAR, Y_t has a structural moving average representation:

$$Y_t = \sum_{k=0}^{\infty} C_k(B) A_0 \varepsilon_{t-k}, \tag{5}$$

where $B = \left(B_1, B_2, \ldots, B_p\right)$, and the notation $C_k(B)$ highlights the dependence of the MA coefficients on the AR coefficients in B, that is

$$C_k(B) = \sum_{m=1}^{k} C_{k-m}(B) B_m, \quad k = 1, 2, \ldots, \tag{6}$$

with $C_0(B) = I_n$ and $B_m = 0$ for $m > p$ (see e.g. Lütkepohl, 1990, 2007).

The structural impulse response coefficient is the response of $Y_{i,t+k}$ to a one-unit change in ε_t^j, which is given by

$$\partial Y_{i,t+k} / \partial \varepsilon_t^j = e_i' C_k(B) A_0 e_j, \tag{7}$$

where e_i and e_j denote the ith and jth columns of the identify matrix I_n, respectively.

Target shock

We are interested in identifying the impulse responses to a *monetary policy shock*, ε_t^{mp}, which we order first without any loss of generality. The impulse responses with respect to this monetary policy shock are determined by $A_0 e_1 = A_{0,1}$ from (7).

Following Gertler and Karadi (2015), we use the 1-year government bond rate as the policy indicator in our baseline analysis, which is associated with exogenous variations due to the structural monetary policy shock ε_t^{mp}. Gertler and Karadi (2015) argue that to include shocks to forward guidance in the measure of the policy innovations, it is important to take as a policy indicator a government bond rate with maturity longer than the current period federal funds rate. The monetary policy shock is identified using an external instrument approach. Let z_t denote an external instrument and $\varepsilon_t^* = \varepsilon_t \setminus \varepsilon_t^{mp}$ be an $(n-1) \times 1$ vector of structural shocks other than the monetary policy shock ε_t^{mp}. The external instrument approach requires z_t to be correlated with ε_t^{mp} but orthogonal to ε_t^*, that is, the following assumption (similar to Gertler & Karadi, 2015) must be satisfied:

Assumption 1. *(a)* $E[z_t \varepsilon_t^{mp}] = cov(z_t, \varepsilon_t^{mp}) = \alpha \neq 0$ *and (b)* $E\left[z_t \varepsilon_t^{*\prime}\right] = 0$.

Assumption (a) states the relevance of z_t as an instrument for ε_t^{mp}, while Assumption (b) implies that it is a valid (or an exogenous) instrument. Stock and Watson (2018) and Montiel Olea *et al.* (2021) show that if the two assumptions hold, then the parameters of interest in the VAR can be estimated by an IV-regression. Therefore, one can identify the impulse responses to a monetary policy shock under (1)–(7), and conduct statistical inference on the identified impulse response coefficients as discussed below.

Identification of the impulse response coefficients

From (7), the impulse response coefficient of interest, $\lambda_{k,i} \equiv \partial Y_{i,t+k} / \partial \varepsilon_t^{mp}$, depends on the VAR coefficient B and the first column $A_{0,1}$ of A_0. $A_{0,1}$ is identified up to a scale by $cov(z_t, u_t)$ under the aforementioned assumptions, that is

$$\Theta = E[z_t A_0 \varepsilon_t] = \alpha A_{0,1}. \tag{8}$$

Without any loss of generality, assume that $A_{0,11} = 1$. Then we have $\Theta_{11} = E[z_t u_t^{mp}] = \alpha$ and $A_{0,1} = \Theta\Theta_{11}^{-1} = \Theta\left(e_1'\Theta\right)^{-1}$. Therefore, the structural impulse responses with respect to ε_t^{mp} are given by:

$$\lambda_{k,i} = \frac{e_i' C_k(B)\Theta}{e_1'\Theta}. \tag{9}$$

Identification of ε_t^{mp}

The monetary policy shock is identified through the projection of the instrument z_t on the reduced-form innovations u_t:

$$\begin{aligned}\mathrm{Proj}(z_t|u_t) &= \Theta'\Sigma^{-1}u_t = (\alpha A_0 e_1)' \\ &\left(A_0 D A_0'\right)^{-1} A_0 \varepsilon_t = \left(\alpha/\sigma_1^2\right)\varepsilon_t^{mp}. \end{aligned} \tag{10}$$

The projection (10) determines ε_t^{mp} up to the scale factor $\left(\alpha/\sigma_1^2\right)$; dividing by $\left(\Theta'\Sigma^{-1}\Theta\right)^{1/2}$ yields $\varepsilon_t^{mp}/\sigma_1$ up to a sign.

To estimate B, one can simply use least squares estimation of the reduced-form VAR in (1). In particular, letting $S_{ab} = T^{-1} \sum_{t=1}^{T} a_t b_t'$ for any matrices a_t and b_t, we have $\hat{B}_T = S_{YX} S_{XX}^{-1}$, where $X_t = \left(1, Y_{t-1}', Y_{t-2}', \ldots, Y_{t-p}'\right)'$. Then, Θ can be estimated as $\hat{\Theta}_T = S_{z\hat{u}}$, where $\hat{u}_t = Y_t - \hat{B}_T X_t$, and $\hat{\Sigma}_T = S_{\hat{u}\hat{u}}$. The impulse responses can then be constructed using (9).

(ii) Monetary Policy Transmission Channel

This section outlines the monetary policy transmission mechanism within the conventional New Keynesian (NK) framework, which motivates the choice of the variables used in the empirical analysis in the following sections. The framework assumes the monetary authority's control over the short-term nominal interest rate. Changes in this rate, induced by policy, lead to adjustments in longer-term nominal rates as agents exploit differences in risk-adjusted expected returns on nominal debt with varying maturities, following the expectations hypothesis of the term structure of interest rates.

Given the assumption of sticky prices and wages, alterations in nominal interest rates translate into changes in real interest rates. This provides the monetary authority with influence over current and expected future real interest rates. As per the expectations hypothesis of the term structure, the real return of an n-period bond in period t ($r_{n,t}$) can be expressed in terms of the current and expected path of short-term real rates (r_{t+j}) and a constant term premium (tp_n) (to a first-order approximation around the steady state):

$$r_{n,t} = E_t\left[\frac{1}{n}\sum_{j=0}^{n-1} r_{t+j}\right] + tp_n. \tag{11}$$

Changes in real interest rates play a pivotal role in influencing aggregate demand, subsequently impacting economic activity and inflation. Higher real interest rates, for example, elevate the real cost of borrowing, prompting firms to scale back investment expenditures. Similarly, households adjust their consumption patterns, favouring future consumption over current consumption. Consequently, aggregate output and inflation decline, leading to an increase in unemployment. The expectations hypothesis further illustrates how forward guidance can influence longer-term yields by effectively communicating the anticipated path of future short rates.

In the transmission channel outlined above, prevalent in standard NK models, financial markets are assumed to operate without frictions. However, when financial frictions are present, the impact of monetary policy actions can extend through a credit channel. Bernanke and Gertler (1995) and Bernanke *et al*. (1999) connect this credit channel of monetary policy to the balance sheets of firms in the economy. A policy-induced increase in the nominal interest rate raises the payments firms must make to service their debt and diminishes the capitalised value of firms' assets. This deterioration in firms' balance sheets leads to a subsequent increase in firms' cost of credit. Therefore, a contractionary monetary policy not only reduces aggregate demand in the economy through the traditional interest rate channel but also elevates firms' cost of capital via the balance sheet channel, amplifying the real effects of monetary policy shocks.

Monetary policy can also impact the economy through shifts in premiums for the safety and liquidity of assets, termed the *convenience yield* by Krishnamurthy and Vissing-Jorgensen (2012). Safe and liquid assets offer a convenience yield, reflecting their ease of trading, collateral posting and roles akin to money. Krishnamurthy and Vissing-Jorgensen (2012) show that the yield on a money-like asset is below the risk-free cost of capital, capturing the liquidity and collateral value of such assets. In the traditional Keynesian framework, the nominal interest rate on a Treasury bond captures the liquidity premium on money-like assets such as cash and current accounts (see Nagel, 2016, for empirical support). As argued by Van Binsbergen *et al*. (2022), altering the supply of Treasury bonds in the secondary market, through actions like open-market operations or quantitative easing/tightening, can have an impact on the scarcity of safe assets, their convenience yield, and consequently, the safe real rate of return.

(iii) Data and Estimation

Based on the monetary policy transmission mechanism discussed in Section II.(ii), the underlying reduced-form VAR in (1) contains six variables: the logarithm of IP, the logarithm of CPI, the 1-year government bond rate, the excess bond premium, the mortgage spread and the commercial paper spread, as in Gertler and Karadi (2015). Real economic activity is measured by the IP index[3] and inflation is measured by the CPI. Mortgage spread is calculated as the difference between the 30-year mortgage rate and the 10-year government bond rate, and it captures the cost of housing finance. Commercial paper spread is the difference between the 3-month commercial paper rate and the 3-month Treasury bill rate, and it captures the cost of short-term business credit and the cost of financing consumer durables. Excess bond premium is the spread measure of Gilchrist and Zakrajšek (2012) and it captures the cost of long-term credit in the non-farm business sector. The inclusion of the spread variables allows monetary policy actions to influence economic activity via the credit channel. In addition to capturing financial frictions in the economy, the commercial paper spread and the excess bond premium capture safety and liquidity premiums (i.e., the convenience yield) as these spread measures compare Treasury securities with special safety and liquidity characteristics with similar pecuniary payoff but no such special attributes. In fact, Krishnamurthy and Vissing-Jorgensen (2012) and Del Negro *et al*. (2017) also measure the convenience yield using the spread between corporate bond and Treasury bond yields of comparable maturities, which is what the excess bond premium measure of Gilchrist and Zakrajšek (2012) does for the vast majority of dollar-denominated bonds publicly issued in the US corporate cash market. Finally, we use the 1-year government bond rate as the policy indicator, the innovations of which incorporate not only the effects of surprises in the current funds rate but also shifts in expectations about the future path of the funds rate, that is, shocks to forward guidance.

For the external instrument, we use the futures rates surprises on FOMC dates (similar to Gertler & Karadi, 2015), which

[3] In the robustness check section, we also use the unemployment rate as the measure of economic activity.

come from the event study analysis of Gürkaynak *et al.* (2005). For each monetary policy announcement, we measure the surprise component of the change in the federal funds rate target using 3-month ahead federal funds futures (henceforth, FF4). These announcements include not just dates on which the FOMC actually changed the federal funds rate, but also dates on which there was an FOMC meeting followed by no change in policy. In particular, letting f_{t+j} be the settlement price on the FOMC day in month t for fed funds futures expiring in $t+j$ and $f_{t+j,-1}$ be the corresponding settlement price for the period before the FOMC meeting, the surprise in the futures rate can be expressed as[4]:

$$surprise = f_{t+j} - f_{t+j,-1}. \qquad (12)$$

Gürkaynak *et al.* (2005) argue that news about the economy on the FOMC day does not affect the policy choice and only information available on the previous day is relevant. Therefore, surprises in fed funds futures on FOMC days can be considered exogenous with respect to the economic and financial variables in the VAR. To focus on the monetary policy decision itself, the surprises in futures rates are measured within a tight window of 30 min as in Gürkaynak *et al.* (2005). Gertler and Karadi (2015) show that the strength of the 'surprise' instrument is not low (see Gertler & Karadi, 2015, Table 3). They also argue that the use of fed funds futures surprises for contracts that expire in the future, for example, 3 months ahead in this case, captures shocks to forward guidance as these surprises reflect revisions in beliefs on FOMC dates about the future path of short-term rates.

We estimate the VAR with 12 lags using monthly data.[5] We consider two different sub-samples: the original sample period (1979:7–2012:6) of Gertler and Karadi (2015) and a post-1984 period (1984:1–2012:6). Note that the 'surprise' instrument FF4 is only available from 1990:1. Therefore, as in Gertler and Karadi (2015), we use the full sample to estimate the VAR lag coefficients and obtain the reduced-form residuals in (2). We then use these reduced-form residuals and the instrument for the overlapping period to identify the contemporaneous impact of policy shocks.[6]

There are important reasons for examining the effects of monetary policy shock on economic activity during the post-1984 period. As discussed in Goodfriend and King (2005) and Mojon (2008), the first few years of Paul Volcker's Chairmanship correspond to an 'incredible disinflation' associated with deep recessions during which inflation became an order of magnitude smaller, and the dynamics of the economy during the adjustment may have been different to the one that prevailed after the disinflation. Monetary policy during this period is better characterised by non-borrowed reserve targeting, during which the Federal Funds rate experienced extreme swings. Therefore, this period is likely to be particularly noisy for measuring policy innovations, as suggested by Coibion (2012). There is also strong evidence of structural breaks in the mid-1980s with a marked decline in macroeconomic volatility (see e.g. McConnell & Perez-Quiros, 2000; Blanchard & Simon, 2001; Stock & Watson, 2002). In particular, McConnell and Perez-Quiros (2000) show the existence of a significant break in output growth volatility in 1984.

III Results

(i) Main Results

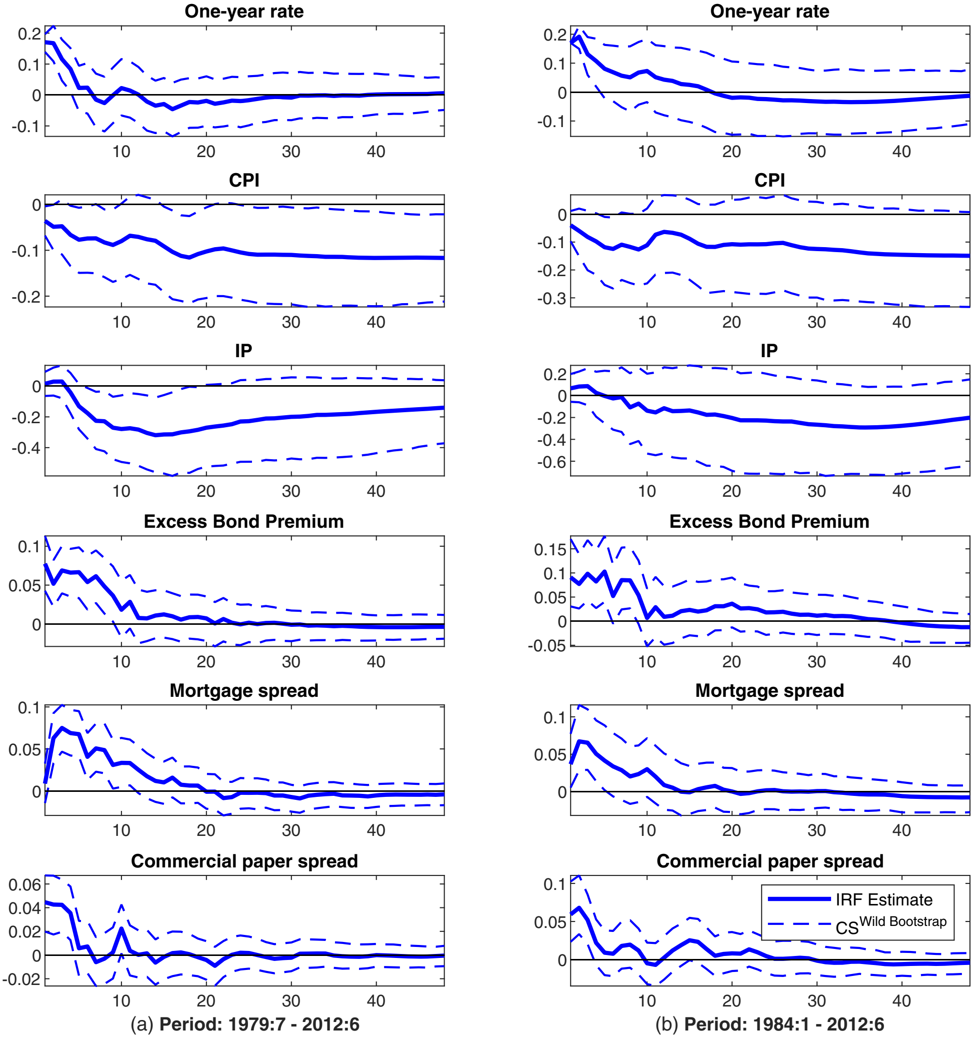
Figure 1 shows the estimated responses of the variables in the baseline model to a one standard deviation monetary tightening shock using the SVAR-IV approach along with the 95 per cent confidence bands. The confidence bands are constructed using

[4] As in Gertler and Karadi (2015), we multiply the surprise in the current month's fed funds futures (f_t) by the factor $\frac{T}{T-t}$, where T is the number of days in the month and t is the number of days elapsed before the FOMC meeting.

[5] We have also used 6, 18 and 24 lags, and our results remain essentially the same.

[6] We have also estimated the VAR over the common sample period (i.e., 1990:1–2012:6) and our results remain the same.

Figure 1
Impulse Responses to a 1-Year Rate Shock



(a) **Period: 1979:7 - 2012:6** (b) **Period: 1984:1 - 2012:6**

the wild bootstrap as in Gertler and Karadi (2015).[7] The left column shows the variables' responses for Gertler and Karadi's (2015) sample (1979:7–2012:6), while the right column illustrates the responses for the post-1984 sample (1984:1–2012:6).[8] Clearly, the left column replicates Figure 2 in Gertler and Karadi (2015).

[7] We use the VAR Toolbox of Ambrogio Cesa-Bianchi (https://sites.google.com/site/ambropo/MatlabCodes?authuser=0) to produce the impulse responses in Figure 1.

[8] In the post-1984 sample we keep the shock size the same as in the full sample for comparability.

Most impulse responses in the post-1984 sample are similar to the ones estimated by Gertler and Karadi (2015). In particular, there is a statistically significant increase in each of the three credit spreads, which is consistent with a credit channel effect on borrowing costs. The CPI declines steadily, although not statistically significantly for the most part. However, there is a notable difference: *the response of output (here IP) is statistically insignificant*. Although the responses are mostly similar and insignificant in longer horizons for both samples, however, in shorter horizons there is a clear distinction: while the response is negative and significant in the full sample, it becomes insignificant when the Volcker disinflation period is left out, despite rises in credit costs. This result points to a notable difference in the effect of monetary policy shock on output than the one documented in Gertler and Karadi (2015).

As mentioned previously, the first years of the Federal Reserve under Volcker correspond to an 'incredible disinflation', during which the inflation rate dropped from 15 per cent in 1980 to 4 per cent in 1983. Monetary policy during this period is better characterised by non-borrowed reserve targeting (rather than interest rate rules) and extreme swings in interest rates. The economy experienced two recessions generally attributed to disinflationary monetary policy. While far less than predicted, the output losses were substantial. Hence, arguably both the operating procedures of the Federal Reserve and the dynamics of the economy were different during the Volcker disinflation.[9] Our finding is in line with Coibion (2012) who, using Romer and Romer (2004)'s approach, shows that dropping the period of non-borrowed reserve targeting by the Federal Reserve during the Volcker disinflation episode significantly lowers the estimated effects of monetary policy shocks.[10] Our finding also aligns with Mojon (2008) who shows that persistent hump-shaped response of inflation to a monetary policy shock disappears during periods without large shifts in the level of inflation such as the post-1984 period. However, notable differences exist. Firstly, we observe modest macroeconomic effects of monetary policy shocks despite the presence of the credit channel in the monetary policy transmission – a mechanism known to theoretically and empirically amplify the impact of monetary policy shocks (Bernanke *et al.*, 1999; Caldara & Herbst, 2019). Secondly, our approach differs in using the proxy VAR methodology. Gertler and Karadi (2015) argue that the standard identification strategy, based on timing restrictions, is unsuitable when identifying monetary policy shocks with financial variables included in the VAR due to simultaneity issues. This is because policy shifts may not only influence financial variables within a period but also be responding to them. The proxy SVAR framework addresses this concern by employing high-frequency identification measures of policy surprises as external instruments, preserving simultaneity between interest rates and financial variables.

One potential issue with the proxy SVAR methodology is the possibility that the policy shock may be confounded with the so-called 'information shock'. Indeed, central bank announcements simultaneously convey information about monetary policy and the central bank's assessment of the economic outlook. Jarociński and Karadi (2020) and Miranda-Agrippino and Ricco (2021) argue that ignoring central bank information shock biases inference on

[9] A contrasting view to this standard narrative suggests that the Volcker disinflation had its roots in 1974 and that monetary policy decisions under then-chair Arthur Burns provided the foundation for economic developments that culminated in the Great Moderation; see, for example, Lubik *et al*. (2016).

[10] Romer and Romer (2004) identify monetary policy innovations by first constructing a historical series of interest rate changes decided upon at the FOMC meetings and then isolating the innovations to these policy changes that are orthogonal to the Federal Reserve's information set. They show that this new measure points to much larger effects of monetary policy shocks than standard VARs.

monetary policy shock. On one hand, a sudden temporary increase in the policy rate should lead to a temporary deterioration of output and inflation, as predicted by theory. On the other hand, an unexpected increase in the policy rate could signal to the markets that the Fed is expecting the economy to heat up, something that could boost the confidence of economic agents and generate expansionary effects. Ignoring this information shock when identifying monetary policy shock could therefore attenuate the macroeconomic effects of the latter due to the presence of the former.

To address this concern, we disentangle the two shocks based on high frequency co-movement of interest rates and stock prices following Jarociński and Karadi (2020), who suggest that a pure monetary policy tightening should lead to a lower stock market valuation while an information shock should lead to a positive co-movement between interest rate and stock price. Therefore, we follow their approach and use the (3-month ahead) fed funds futures surprises in the months when the stock price surprise had the opposite sign of the fed funds futures surprise as the instrument for monetary policy shocks in the proxy SVAR (the instrument is zero otherwise). Jarociński and Karadi (2020) call this methodology 'poor man's sign restrictions'. The implicit assumption is that each month can be classified as hit either by a pure monetary policy shock or by a pure central bank information shock. As Jarociński and Karadi (2020) discuss, this identifying assumption is stronger than the baseline sign restrictions in Jarociński and Karadi (2020), but it is also straightforward to implement, particularly in the context of the proxy SVAR setting.

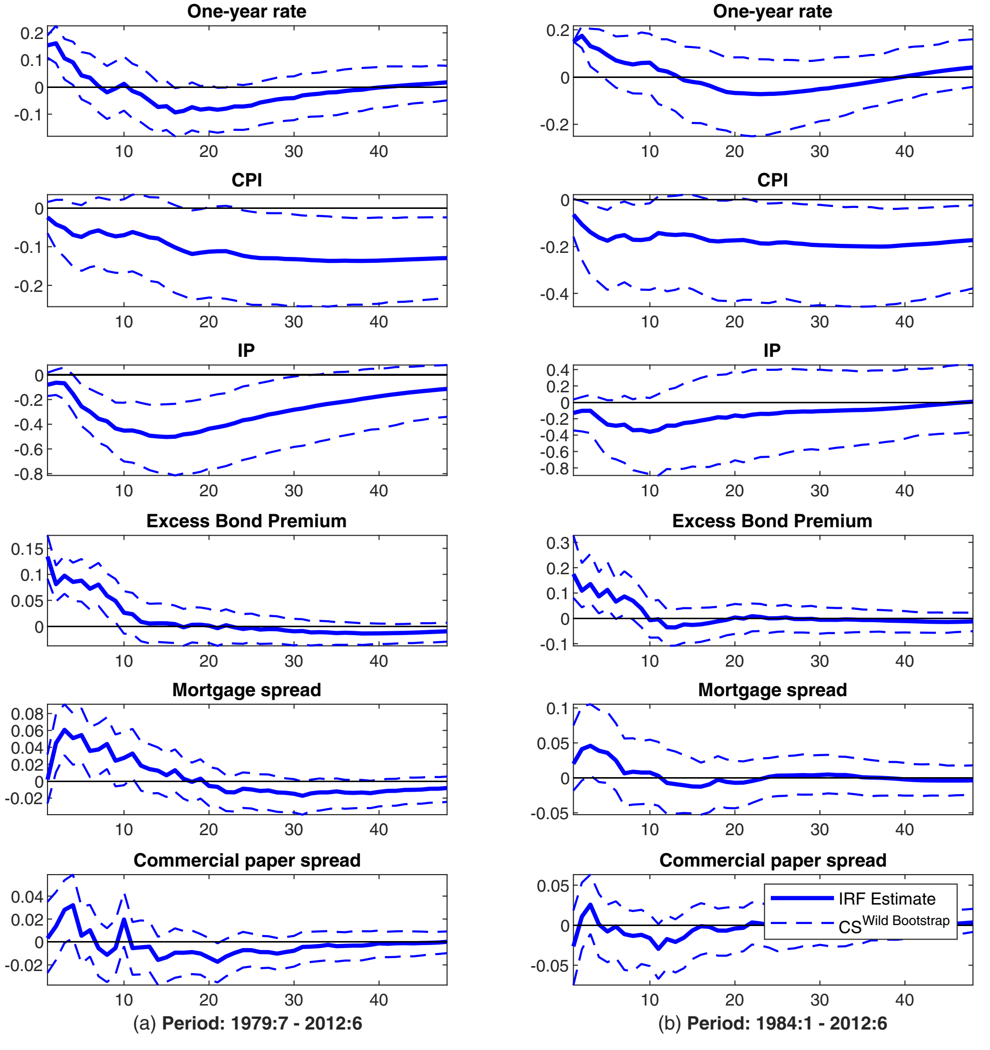
Figure 2 reports the results for both Gertler and Karadi's (2015) sample (left panel) and the post-1984 sample (right panel). As seen in the left panels of this figure, an interest rate increase accompanied by a stock price decline leads to a more significant contraction in output together with a more pronounced price-level decline, which are in line with Jarociński and Karadi's (2020) findings. When looking at the effects in the post-1984 sample (right panels), we find that monetary policy shocks have statistically significant effects on prices even when the Volcker disinflation period is left out. However, monetary policy shocks induce mostly insignificant effects on output, thereby corroborating our previous finding. The Appendix S1 reports the estimated impulse responses when using Miranda-Agrippino and Ricco's (2021) proxy as an instrument and the findings remain similar.[11]

(ii) Alternative Proxy SVAR Inference Methods

In the baseline analysis, the confidence sets for the impulse response functions were obtained using the wild bootstrap (following Gertler & Karadi, 2015). However, Brüggemann *et al*. (2016) show that wild bootstraps are in general not asymptotically valid for inference about estimators that involve the covariance matrix of VAR innovations. Specifically, they argue that the wild bootstrap confidence intervals for recursively identified impulse responses may underestimate the true estimation uncertainty in finite samples. Jentsch and Lunsford (2021) show that the finding of Brüggemann *et al*. (2016) also applies to the wild bootstrap methodology in Mertens and Ravn (2013), and that confidence intervals for impulse response functions can be influenced by the bootstrap algorithm used. Jentsch and Lunsford (2019, 2021) propose a variant of the moving block bootstrap as an alternative inference method in proxy SVARs. Unlike Mertens and Ravn (2013), the block bootstrap method leads to no statistically significant effects of identified tax shocks on the main outcome variables. In their response to Jentsch and Lunsford (2019), Mertens and Ravn (2019) suggest that there is no a priori reason to prefer the block bootstrap type confidence intervals over any of the available asymptotically valid alternatives. In particular, they show that the significance of tax shocks remain robust when using the δ-method or the parametric bootstrap-type confidence intervals as described in Montiel Olea *et al*. (2021). Therefore, to investigate the robustness of our results to alternative asymptotically valid inference methods, we

[11] See Figure S.6 in the Appendix S1.

FIGURE 2
Impulse Responses to a 1-Year Rate Shock Using Jarociński and Karadi's (2020) Proxy as Instrument



use the δ-method type confidence sets for impulse response functions based on the recently developed plug-in estimator by Montiel Olea *et al*. (2021).

The plug-in estimator and the associated δ-method confidence set for the impulse response function, $\lambda_{k,i}$, are given by:

$$\hat{\lambda}_{k,i}\left(\hat{B}_T, \hat{\Theta}_T\right) = e_i' C_k\left(\hat{B}_T\right)\hat{\Theta}_T / e_1'\hat{\Theta}_T, \quad (13)$$

$$C_\beta^{Plug-in}(\lambda_{k,i}) = \left(\lambda_{k,i} \middle| \frac{T\left(\hat{\lambda}_{k,i}\left(\hat{B}_T, \hat{\Theta}_T\right) - \lambda_{k,i}\right)^2}{\hat{\sigma}^2_{T,k,i}} \leq \chi^2_{1,1-\beta}\right), \quad (14)$$

respectively, where $\chi^2_{1,1-\beta}$ is the $(1-\beta)$th quantile of a χ^2_1 distributed random variable for some $\beta \in (0,1)$, $\hat{\sigma}^2_{T,k,i}$ is a consistent estimator of $\sigma^2_{k,i}$ that depends on the asymptotic variance of $(\hat{B}_T, \hat{\Theta}_T)$ and the gradient of $\lambda_{k,i}(B,\Theta)$ with respect to (B,Θ) and can be obtained using the δ-method or a suitable bootstrap procedure.[12] Provided that the instrument z_t is strong, $C^{Plug-in}_{\beta}(\lambda_{k,i})$ has level $1-\beta$ asymptotically; see Montiel Olea *et al*. (2021).

The δ-method yields consistent estimates of the VAR parameters and the impulse responses only if the instrument z_t is strong and valid for ε^{mp}_t. Montiel Olea *et al*. (2021) emphasise the possibility that the external instrument may be weakly correlated with the target structural shock, hence biasing the standard SVAR-IV estimates of the impulse response functions. Gertler and Karadi (2015) also raise concerns about this weak identification issue but they show that their baseline instrument (FF4) is not weak. However, their first-stage robust *F* statistics indicate that most of their other instruments, all of which come from the event study analysis of Gürkaynak *et al*. (2005), are weak, particularly when used together with the 2-year government bond rate as the policy indicator.[13] Therefore, given Jentsch and Lunsford's (2019) criticism of the wild bootstrap and recent developments of asymptotically valid weak-identification robust inference methods within the proxy SVAR framework (see e.g. Montiel Olea *et al*., 2021), we also demonstrate the robustness of our baseline results using weak-IV robust inference. We then illustrate the extent to which some of Gürkaynak *et al*.'s (2005) surprise external instruments are uninformative in identifying monetary policy shocks.

Montiel Olea *et al*. (2021) show how small values of $e_1'\Theta$ in the denominator of (9), which arise when $cov(z_t, \varepsilon^{mp}_t) = \alpha$ is small, may lead to poor coverage of the confidence set $C^{Plug-in}_{\beta}(\lambda_{k,i})$ in (14). They propose a weak IV-robust method based on Anderson and Rubin (1949, AR-statistic) to build confidence sets for the impulse response functions. Specifically, using similar notations as in Section II.(i), define

$$H_T = \begin{pmatrix} e_i' C_k(B)\Theta_T \\ e_i'\Theta_T \end{pmatrix} \equiv \begin{pmatrix} H_{T,1} \\ H_{T,2} \end{pmatrix}, \qquad (15)$$

where $\Theta_T = \alpha_T A_{0,1}$, $E[z_t \varepsilon^{mp}_t] = \alpha_T \to \alpha$ as $T \to \infty$, and $\alpha = 0$ is allowed. This framework allows for both strong instrument ($\alpha_T = \alpha \neq 0$) and weak instrument as in Staiger and Stock (1997) ($\alpha_T = \alpha_0/\sqrt{T}$ for some constant α_0, so that $\alpha_T \to 0$). From (9), the impulse response coefficient can be written as $\lambda_{k,i} = H_{T,1}/H_{T,2}$. Montiel Olea *et al*. (2021) show that $\hat{H}_T \overset{a}{\sim} N(H_T, T^{-1}\Omega)$, where $\hat{H}_T$ denotes the plug-in estimator of H_T, $\Omega = G(B,\Theta) W G(B,\Theta)'$ with G denoting the gradient of $\lim_{T\to\infty} H_T$ with respect to (B,Θ), and W is the asymptotic variance of $\sqrt{T}\left[vec(\hat{B}_T - B)', \hat{\Theta}_T - \Theta_T\right]'$. We are interested in testing the null hypothesis $H_0 : \lambda_{k,i} = \lambda_0$. Since $\lambda_{k,i} = H_{T,1}/H_{T,2}$, H_0 can be formulated as a linear restriction on $H_{T,1}$ and $H_{T,2}$, that is, $H_0 : H_{T,1} - \lambda_0 H_{T,2} = 0$. To assess H_0, Montiel Olea *et al*. (2021) propose to use the Wald statistic

$$q_T(\lambda_0) = \frac{T\left(\hat{H}_{T,1} - \lambda_0 \hat{H}_{T,2}\right)^2}{\hat{\omega}_{T,11} - 2\lambda_0 \hat{\omega}_{T,12} + \lambda_0^2 \hat{\omega}_{T,22}}, \qquad (16)$$

where $\hat{\omega}_{T,ij}$ are consistent estimates of the elements of Ω. Under H_0, $q_T(\lambda_0) \xrightarrow{d} \chi^2_1$, irrespective of the strength of the instrument z_t. Therefore, the AR-type confidence set for $\lambda_{k,i}$ with level $1-\beta$ is obtained by inverting $q_T(\lambda_0)$ (see e.g. Dufour & Taamouti, 2005; Doko Tchatoka & Dufour, 2014; Montiel Olea *et al*., 2021):

$$C^{AR}_{\beta}(\lambda_{k,i}) = \left\{\lambda_{k,i} : q_T(\lambda_{k,i}) \le \chi^2_{1,1-\beta}\right\}. \qquad (17)$$

Figure 3 shows the estimated impulse responses and the corresponding 95 per cent confidence interval with the plug-in method (i.e., $C^{Plug-in}_{\beta}(\lambda_{k,i})$ given in (14)) and the weak-IV robust method (i.e., $C^{AR}_{\beta}(\lambda_{k,i})$ given in (17)). To enable comparison with

[12] Montiel Olea *et al*. (2021) show that $\sqrt{T}\left[\hat{\lambda}_{k,i}(\hat{B}_T, \hat{\Theta}_T) - \lambda_{k,i}(B,\Theta)\right]$ converges in distribution to $N(0, \sigma^2_{k,i})$.

[13] Gertler and Karadi (2015) suggest that the 2-year rate is the conceptually preferred policy indicator, based on the argument of Swanson and Williams (2014) and Hanson and Stein (2015) who argue that the Federal Reserve's forward guidance strategy operates with roughly 2-year horizon.

FIGURE 3
Weak-Instrument Robust Impulse Responses to a 1-Year Rate Shock; Period: 1984:1–2012:6

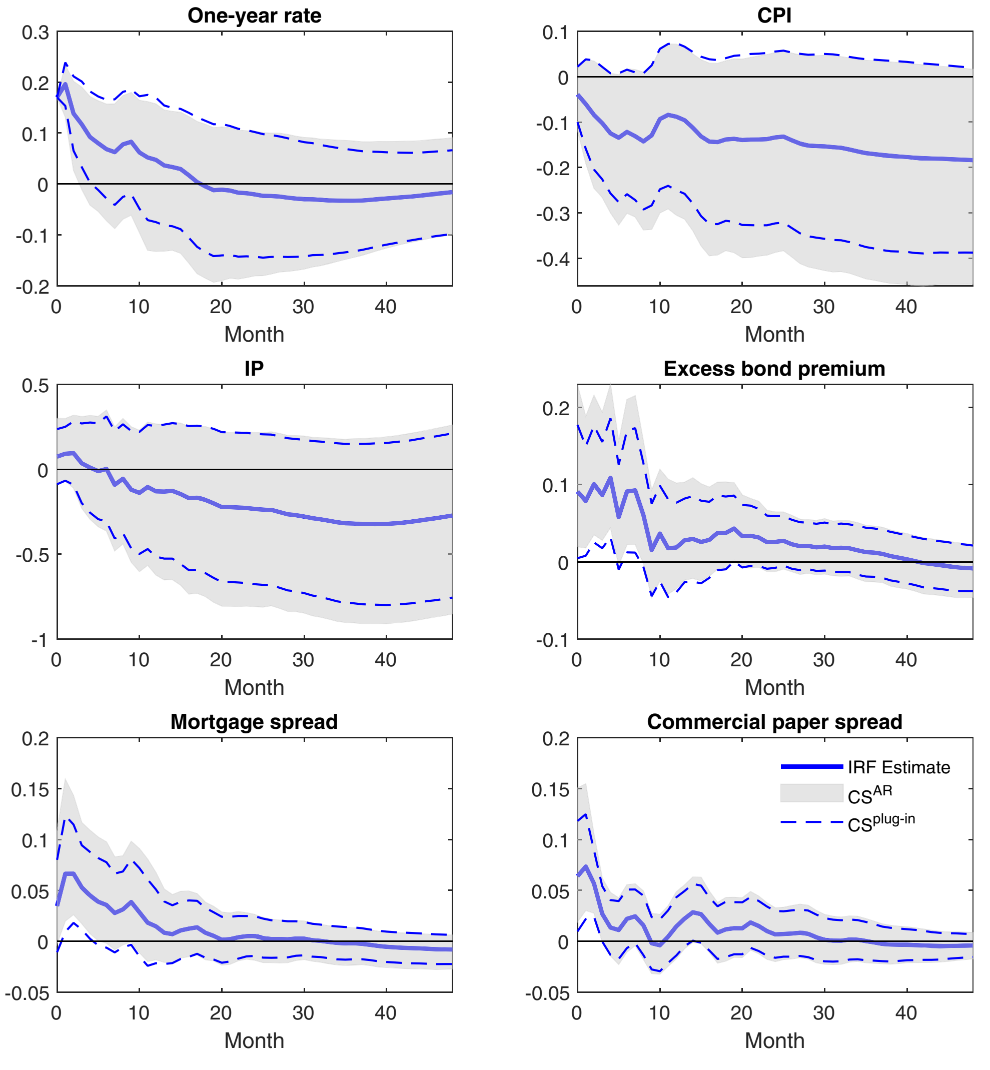


Figure 1, the size of the monetary policy shock is normalised so that the response of the 1-year rate is quantitatively the same in both figures. The VAR is estimated using 12 lags and a constant term as before. The covariance matrix W is estimated using the Eicker–White robust estimator and confidence bounds shown in Figure 3 are based on the δ-method as in Montiel Olea *et al*. (2021).[14] From the plots in Figure 3, two important observations are in order.

[14] A bootstrap method can also be used; see Appendix A.4 in Montiel Olea *et al*. (2021).

First, we note that the weak-instrument robust confidence sets (grey areas) mostly coincide with their standard plug-in counterparts (areas delimited by the two blue dashed lines), hence corroborating Gertler and Karadi's (2015) finding that the 3-month ahead federal funds futures surprise instrument ($FF4$) is not weak. Second, the impulse responses in Figure 3 are essentially the same as those in Figure 1. Both the $C_{\beta}^{Plug-in}(\lambda_{k,i})$ and $C_{\beta}^{AR}(\lambda_{k,i})$ confidence sets in Figure 3 show statistically insignificant effects of monetary policy shocks on inflation and output despite substantial increase in credit spreads, thereby confirming our previous results.[15]

Next, we estimate the VAR using the other external instruments from Gertler and Karadi's (2015) analysis, namely the surprises in the current month's fed funds futures (FF1), and the 6-month, 9-month and 1-year ahead futures surprises in 3-month Eurodollar deposits (henceforth ED2, ED3 and ED4, respectively).[16] The impulse responses are mostly similar to those reported in Figure 3 when FF1 and ED2 are used as external instruments, and also the $C_{\beta}^{Plug-in}(\lambda_{k,i})$ and $C_{\beta}^{AR}(\lambda_{k,i})$ confidence sets mostly coincide – therefore suggesting that these instruments are strong. In contrast, the impulse responses when using ED3 and ED4 as external instruments paint a different picture. For example, Figure 4 shows the impulse responses when ED4 is used as the external instrument. As seen in the figure, there are substantial discrepancies between the $C_{\beta}^{Plug-in}(\lambda_{k,i})$ and $C_{\beta}^{AR}(\lambda_{k,i})$ confidence sets, with the latter being much wider for all variables, indicating that this instrument is weak. Therefore, the confidence sets based on the plug-in method in Figure 4 are invalid in the sense that their actual coverage probability (level) can be zero (see Dufour, 1997). Notwithstanding the weak instrument problem, we see that the 95 per cent AR-type confidence sets show statistically insignificant effects of monetary policy shocks on inflation and output.

IV Exploring Why Monetary Policy Shocks Have Modest Macroeconomic Effects

Various explanations have been proposed for the modest macroeconomic effects of monetary policy shocks. For instance, Hanson (2008) highlights the role of decreased output volatility during the Great Moderation. Castelnuovo and Surico (2010) and Castelnuovo (2016), among others, suggest that technological progress or financial innovations contribute to households' enhanced consumption smoothing. Boivin and Giannoni (2006) and Boivin *et al.* (2010) attribute the moderate macroeconomic responses after 1984 to an assertive monetary authority's effective handling of inflation. According to their perspective, the muted effects of monetary policy shocks do not necessarily indicate a decline in the effectiveness of monetary policy but rather an improvement in its conduct.

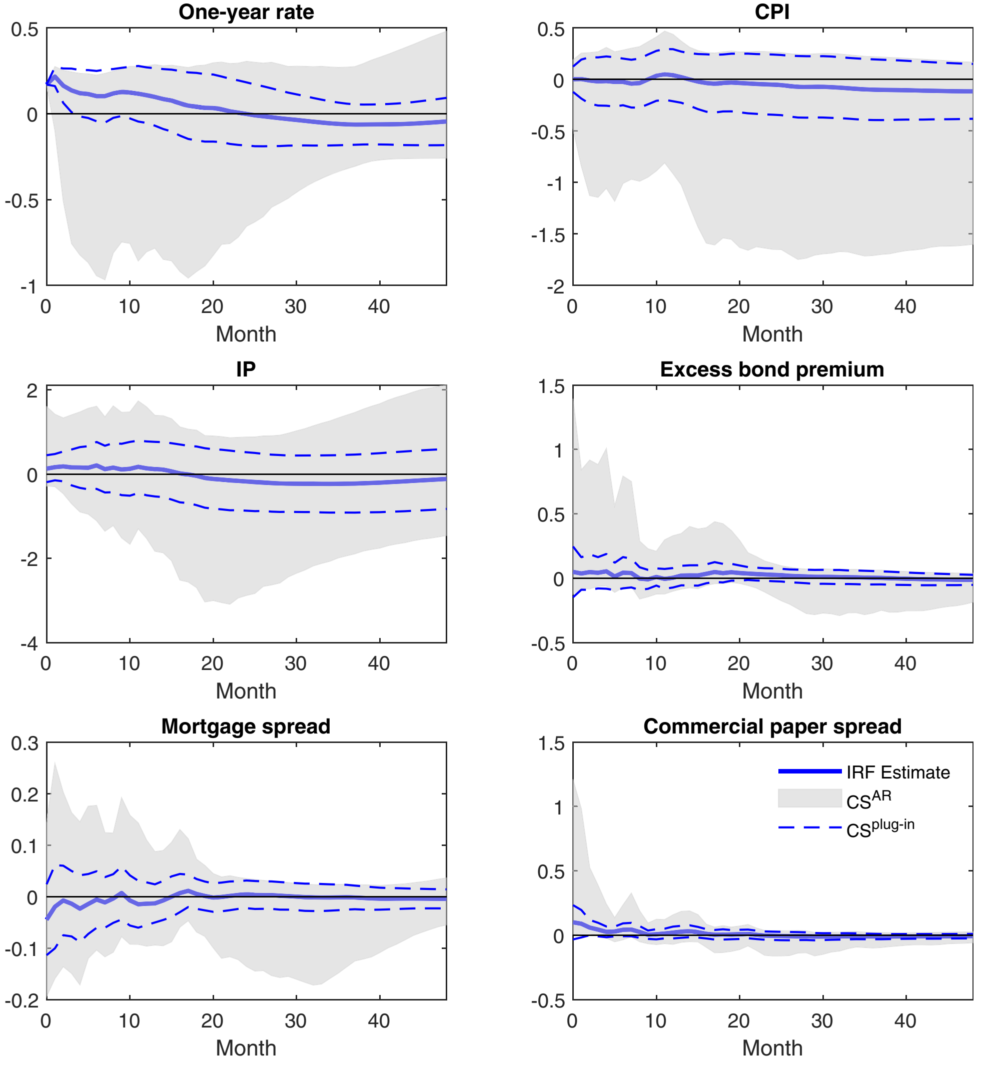
In this section, we explore changes in the implementation of systematic monetary policy as a potential explanation for the limited macroeconomic effects of policy shocks. Following the approach of Boivin *et al.* (2010), we analyse the impulse responses to a monetary policy shock using a structural NK model. However, a crucial distinction from Boivin *et al.*'s (2010) analysis is that we incorporate the credit channel of the monetary policy transmission mechanism, which is known to magnify the impact of policy shocks. Specifically, we examine the effects of a monetary policy shock within the renowned Bernanke *et al.* (1999) (henceforth referred to as BGG) financial accelerator model, which introduces credit market frictions into an otherwise standard NK model.

Section B in Appendix S1. offers the complete set of linearised equilibrium conditions for the BGG model. However, we outline the mechanics of the pivotal equation in the model pertaining to the financial accelerator mechanism below:

[15] We have also estimated the VAR using Jarociński and Karadi's (2020) proxy for monetary policy shock and the results remain robust. However, Jarociński and Karadi (2020) proxy turns out to be relatively weak in the post-1984 period as the identification-robust confidence sets turn out to be wider than their strong-instrument δ-method counterpart. See Figure S.5 in the Appendix S1.

[16] The results are reported in Figures S.2–S.4 in the Appendix S1.

FIGURE 4
Impulse Responses to a 1-Year Rate Shock with ED4 as Instrument; Period: 1984:1–2012:6



$$E_t r^k_{t+1} - r_t = \nu[n_t - (q_t + k_{t+1})], \qquad (18)$$

where r_t is the safe real interest rate, n_t is the net worth of a firm, q_t is the price of capital, k_{t+1} is the capital stock accumulated at time t available for production at time $t+1$ and $E_t r^k_{t+1}$ is the expected return on capital. Therefore, the left-hand side of Equation (18) can be interpreted as the external finance premium and the right-hand side is the (negative of the) leverage ratio, that is, value of assets $(q_t + k_{t+1})$ relative to equity (n_t), and ν is the elasticity of the external finance premium with respect to a change in the leverage position of entrepreneurs.

The financial accelerator mechanism arises as a result of asymmetric information between borrowers (entrepreneurs) and lenders (financial intermediaries). Essentially, entrepreneurs produce intermediate goods using its own net worth (internal fund) and borrowing (external fund) from a financial intermediary. The information asymmetry between entrepreneurs and lenders makes the acquisition of capital using external fund more costly than internal fund (net worth) so that entrepreneurs' demand for capital depends on their financial position. The key insight of the BGG model is that a decrease (increase) in borrower's net worth (n_t) increases (reduces) credit market frictions if $\nu > 0$. This then raises (reduces) the external finance premium and discourages (stimulates) investment and aggregate demand. The essence of the accelerator is that contractionary (expansionary) shocks, such as an unexpected increase (decrease) in interest rates, that reduce (increase) asset prices amplify the negative (positive) effects of such shocks due to a balance sheet deterioration (improvement) of borrowers.

Following BGG, we assume that the monetary authority sets nominal interest rates according to the following interest rate rule:

$$r_t^n = \rho r_{t-1}^n + (1-\rho)\zeta\pi_{t-1} + \epsilon_t^{rn}, \tag{19}$$

where r_t^n is the nominal interest rate, π_{t-1} is lagged inflation, ϵ_t^{rn} is an *i.i.d* policy shock that captures unexpected or surprise changes in the nominal interest rate, ρ is the degree of interest rate smoothing and ζ captures the monetary authority's responsiveness to inflation.[17]

Figure 5 shows the response of output, investment, nominal interest rate and inflation to a 20 basis points surprise increase in the nominal interest rate.[18] Most of the parameters are calibrated following BGG, and details of the calibration are provided in the Appendix S1. The key parameter pertaining to credit market friction is ν, the elasticity of the external finance premium with respect to a change in the leverage position of entrepreneurs, which we set at 0.05 following the estimate of Christensen and Dib (2008).[19] For the interest rate smoothing parameter ρ, we set it at 0.7 following the estimate of Hirose *et al*. (2020). Finally, for the response to inflation in the Taylor rule ζ, we set it at different values starting from $\zeta = 1.5$ (the value used by Taylor, 1993) and gradually increasing the degree of responsiveness to $\zeta = 4$. Haque *et al*. (2021) and Haque (2022) estimate variants of the NK model using Bayesian estimation techniques for the post-mid-1980s period and find that the Federal Reserve has been highly aggressive in its response to inflation during this period. For instance, the posterior mean estimates of the degree of response to inflation are 3.09 and 4.04, respectively, in Haque *et al*. (2021) and Haque (2022).

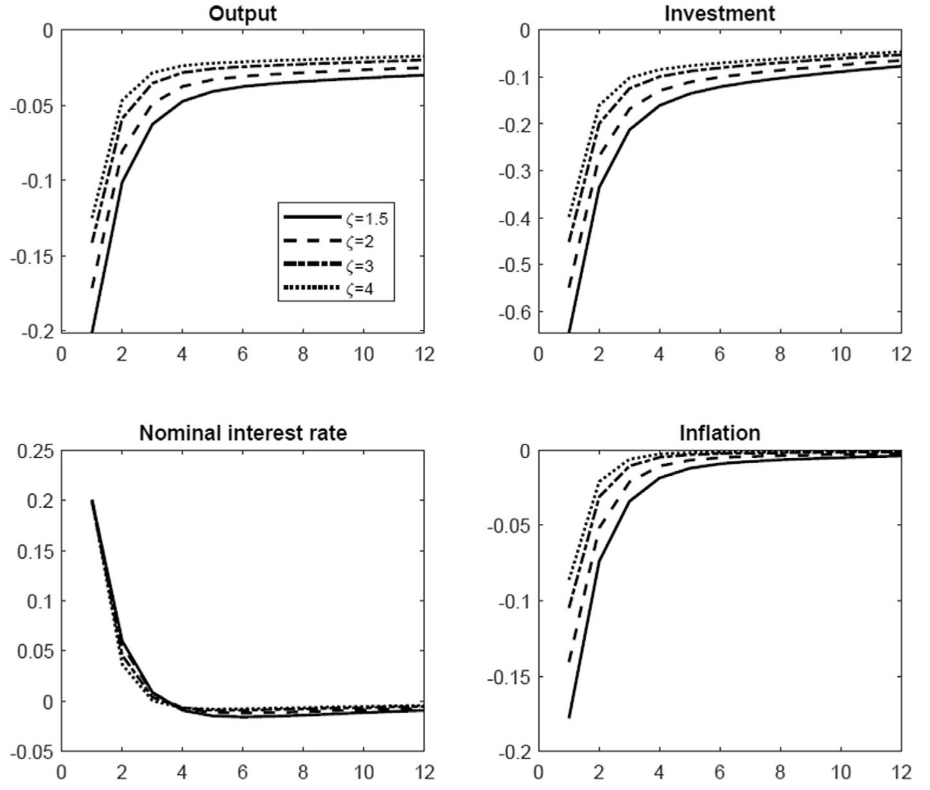
As shown in Figure 5, a stronger response to inflation dampens the effects of the monetary policy shock on output, investment and inflation, and this dampening takes place despite the presence of the financial accelerator channel. As such, our results corroborate Boivin *et al*. (2010)'s finding that modest macroeconomic effects of monetary policy shocks are attributable to a change in the systematic conduct of monetary policy. Therefore, such modest effects do not necessarily reflect diminished effects of monetary policy but rather suggest that by responding more strongly to inflation monetary policy has stabilised the economy more

[17] We have also considered variants of the policy rule that allow for responses to output as well as contemporaneous or expected inflation (as opposed to lagged inflation), and the results discussed below remain largely similar. As Bernanke *et al*. (1999) point out, the greater the extent to which monetary policy is able to stabilise output, the smaller is the role of the financial accelerator to amplify and propagate business cycles.

[18] The shock size of the nominal interest rate is set in line with the impact response of the nominal rate in the VAR.

[19] BGG discuss the parameters related to financial frictions in detail in p. 1368 of their paper.

FIGURE 5
Impulse Responses to a Monetary Policy Shock in the Bernanke et al. *(1999) Model for Varying Degree of Responsiveness to Inflation in the Monetary Policy Rule*



effectively in the post-1980s period. There is ample evidence in the empirical macroeconomic literature that supports an improvement in the conduct of monetary policy during the Great Moderation through a stronger response to inflation (Clarida *et al.*, 2000; Lubik & Schorfheide, 2004; Haque *et al.*, 2021; Haque, 2022).

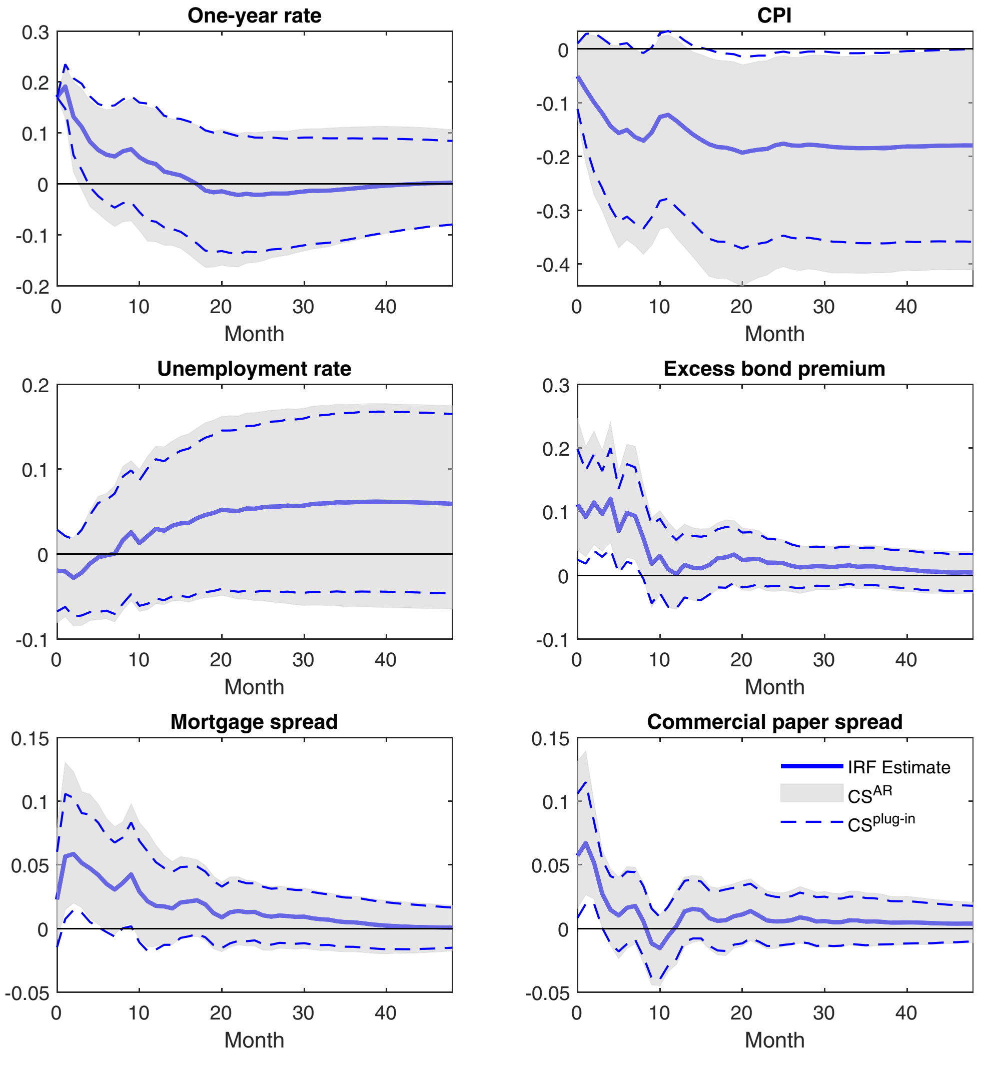
V Robustness Checks

In this section, we investigate the robustness of our results to alternative measure of economic activity and policy indicator, alternative specification and different sample periods.

(i) Unemployment as Measure of Economic Activity

In the empirical VAR literature the unemployment rate is commonly used as an indicator of real economic activity. Therefore, we replace the IP index in the baseline VAR in (1) with the unemployment rate. As in Section III.(i), the change in 3-month ahead federal funds futures (FF4) is used as external instrument to identify the monetary policy shock. Figure 6 shows the impulse response functions to a monetary tightening shock, where the size of the shock is normalised such that the impact response of the 1-year rate is the same as in the

Figure 6
Impulse Responses to a 1-Year Rate Shock with Alternative Economic Activity Indicator; Period: 1984:1–2012:6



baseline estimation. As seen in this figure, the responses of CPI, the 1-year bond rate and the spread variables are similar to those observed in Figure 3. In particular, there is a decline in the CPI in line with economic theory but the response is relatively modest. With regard to the unemployment rate, although it goes up as expected, the response is not statistically significant despite large increases in credit costs.[20]

(ii) Alternative Policy Indicator and Forward Guidance

In this section, we study the role of unconventional monetary policy. In response to the Great Recession, the Federal Reserve, like many other central banks, cut its policy rate close to zero and was unable to use conventional policy measures to further stimulate the economy. However, central banks including the Fed relied on a particular type of unconventional policy, namely forward guidance, to influence market's beliefs about the expected path of future short rates via communication and, in turn, to stimulate the economy. To identify the forward guidance shock, ideally one would follow Gürkaynak *et al*. (2005) to disentangle the component of high-frequency surprises that reflect surprises in future rates (labelled path factor by Gürkaynak *et al.*, 2005) that are orthogonal to surprises in the current short-rate (the target factor). However, as Gertler and Karadi (2015) note, such decomposition between the target and path factors leads to instruments being too weak to credibly identify forward guidance shocks.

Alternatively, we adopt an indirect approach, using the 2-year government bond rate as the policy indicator. Swanson and Williams (2014) argue that the zero lower bound (ZLB) did not impede the Federal Reserve's manipulation of the 2-year rate, and that the Federal Reserve's forward guidance strategy operates on a roughly 2-year horizon. In addition, Gürkaynak *et al*. (2005) provide evidence that forward guidance significantly impacts futures rates relevant for pricing the 2-year bond rate.[21] Therefore, the 2-year government bond rate may serve as a more comprehensive indicator of monetary policy, encompassing forward guidance shocks. By comparing the responses to a 2-year rate shock with our baseline results using the 2-year rate as the policy indicator, we aim to discern whether forward guidance shocks have distinct effects on the economy compared to conventional policy shocks.

We estimate a seven-variate VAR with the 2-year government bond rate added as the policy indicator. Similar to the baseline model, the change in the 3-month ahead federal funds futures (FF4) is used as an external instrument to identify the policy shock. The use of fed funds futures surprises for contracts that expire in the future, that is, FF4 in this case, captures shocks to forward guidance as discussed earlier. Figure 7 presents the results, where the size of the 2-year rate shock is normalised such that the impact response on the 1-year rate is similar to the one in our baseline estimation. As seen in the figure, the 2-year rate shock produces responses similar to the baseline model with responses of inflation and output remaining statistically insignificant as before, suggesting that forward guidance shocks have similar effects as conventional policy shocks.

For completeness, we also estimate the VAR by replacing FF4, one at a time, with FF1, ED2, ED3 and ED4 as the external instrument. These results are reported in Figures S.7–S.10 in the Appendix S1. In line with our previous results, the identification-robust confidence sets for the impulse responses clearly show that FF1 and ED2 are not weak and their use yield similar results as those with FF4. However, ED3 and ED4 remain weak even with the 2-year bond rate as the policy indicator, which is reflected in the wide weak-IV robust confidence sets for the impulse response functions (see Figures S.9 and S.10 in the Appendix S1 for ED3 and ED4, respectively). In fact, the identification-robust confidence sets for the impulse response functions with ED4 used as an external instrument are unbounded, suggesting that ED4 is completely uninformative. Dufour (1997) shows that no valid confidence set which is almost surely bounded exists in such a case.

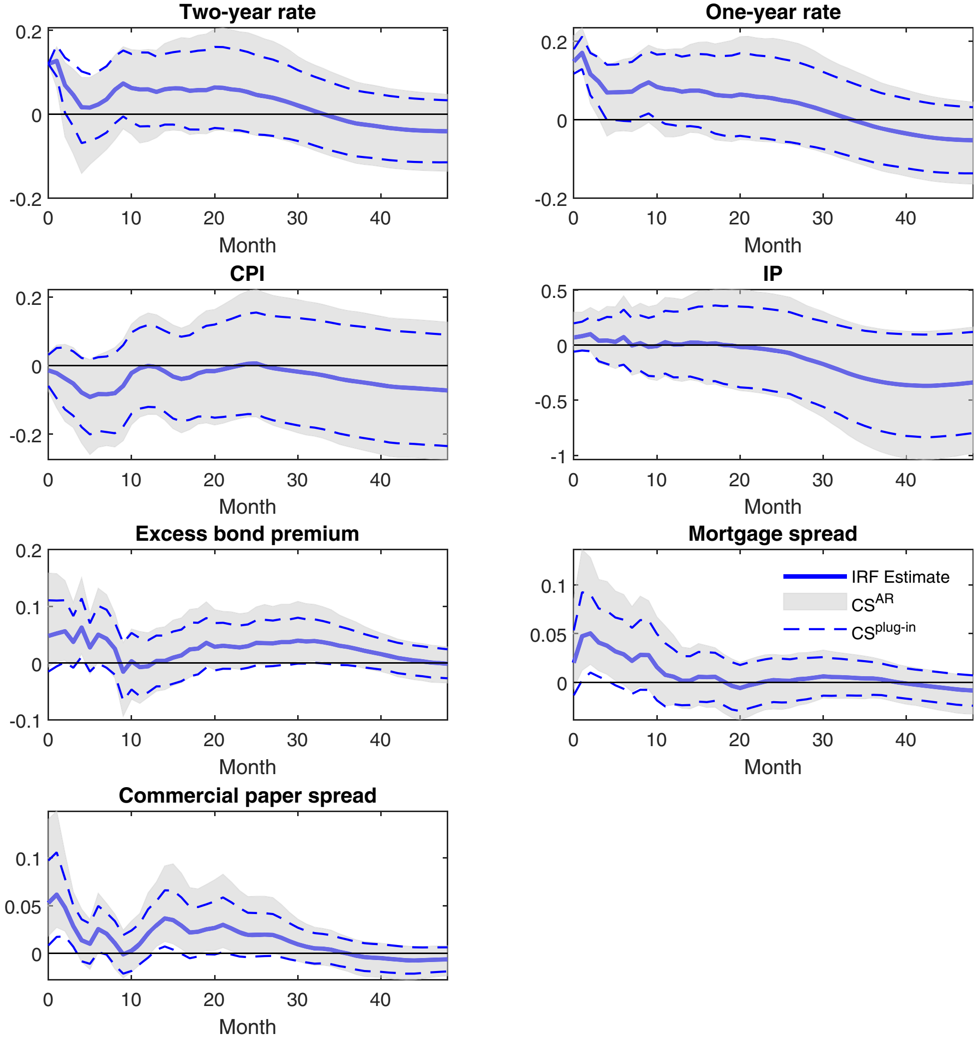
(iii) Alternative Sample Period

We also extend our analysis by estimating the VAR over the period 1984:1–2020:2, that is, also including the ZLB period.

[20] Liu *et al*. (2019) also find that the response of the unemployment rate to a monetary policy shock has declined in the post-1990s, suggesting that the transmission of monetary policy shocks has significantly changed over time.

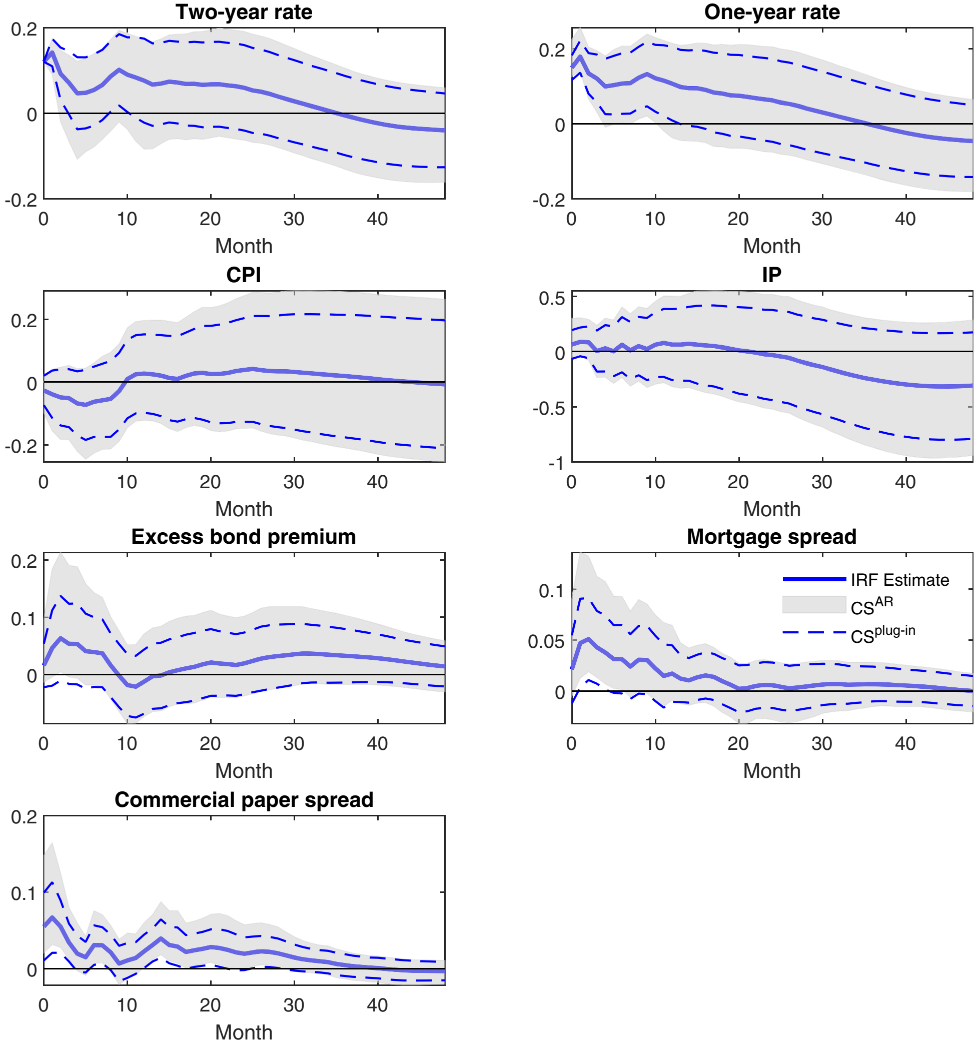
[21] See also Hanson and Stein (2015) and Gertler and Karadi (2015).

FIGURE 7
Impulse Responses to a 2-Year Rate Shock with FF4 as Instrument; Period: 1984:1–2012:6



Monthly data on fed funds futures surprises for this extended sample were constructed as in Gürkaynak *et al*. (2005) and Gertler and Karadi (2015). We continue to use the 2-year government bond rate as the policy indicator to address concerns related to the ZLB while the policy shock is instrumented using the change in 3-month ahead federal funds futures (FF4) as before. We replace the excess bond premium with Moody's Baa spread, as the former is only available through August 2016. Figure 8 shows the responses of the variables to a monetary tightening shock where again the size of the 2-year rate shock is normalised such that the impact response on the 1-year rate is similar to the one in our baseline estimation. As seen in this figure, the effects of the shock on both

FIGURE 8
Impulse Responses to a 2-Year Rate Shock with FF4 as Instrument, Period: 1984:1–2020:2



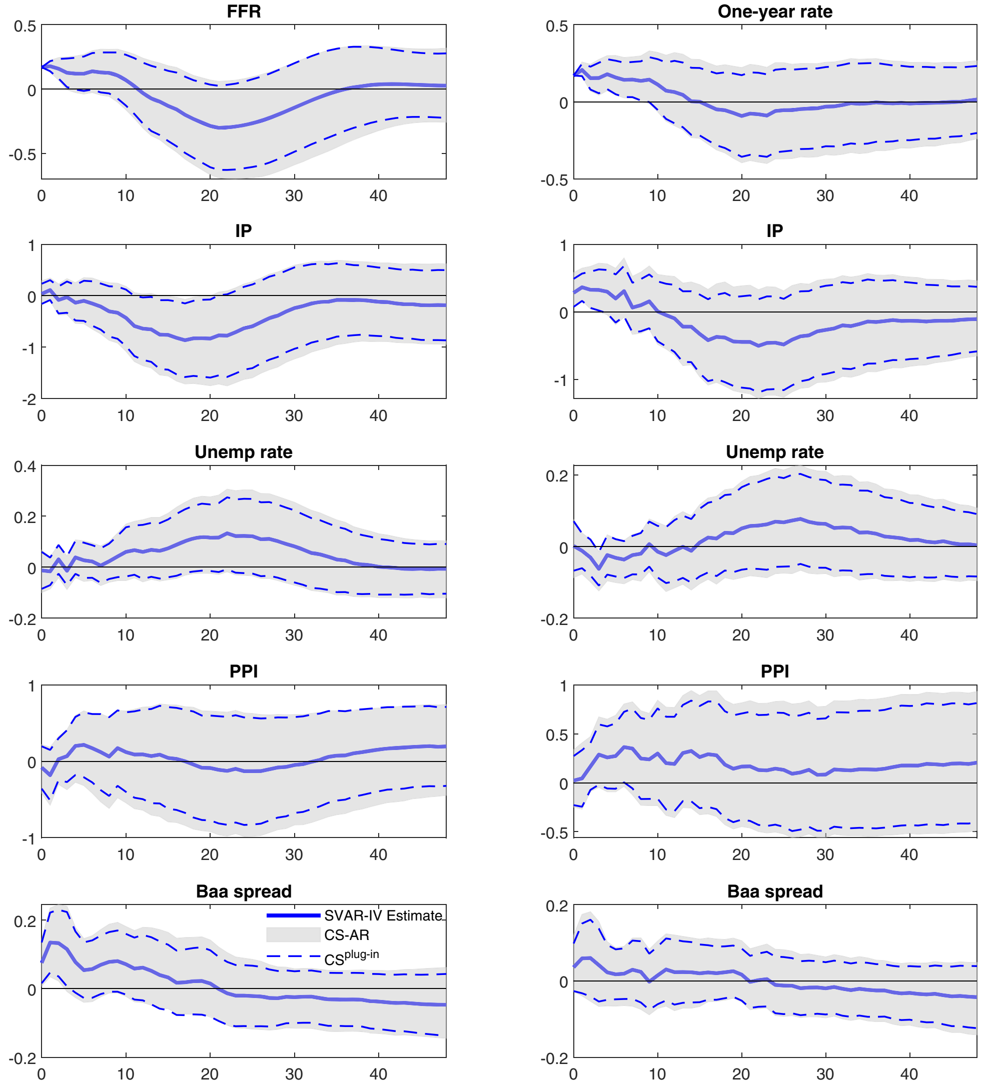
inflation (CPI) and output (IP) are statistically insignificant, hence confirming our previous analyses. The results remain essentially the same when the 1-year government bond rate is used as the policy indicator or when the surprise in the current month's fed funds futures (FF1) is used as instrument.

Ikeda *et al*. (2020) suggest that the dynamics of the economy is different at the ZLB with unconventional monetary policy being less effective on output. Notwithstanding, the results are similar to those reported here when the ZLB periods are excluded, that is, when the sub-sample 1984:1–2008:6 is used in the estimation.[22]

[22] See Figure S.13.

FIGURE 9

Impulse Responses to a Monetary Policy Shock in Caldara and Herbst's (2019) Model, Period: 1994:1–2007:6



(iv) Alternative Specification

Our finding in this paper departs from Caldara and Herbst (2019) who show that monetary policy shock induces persistent declines in real activity owing to a systematic reaction of monetary policy to changes in corporate credit spreads. Caldara and Herbst (2019) estimate a proxy SVAR using similar high frequency external instruments as in Gertler and Karadi (2015) over a sample period starting in 1994 (i.e., well beyond the Volcker disinflation period) and ending in June 2007. To investigate further, we estimate their proxy VAR which consists of the federal funds rate, the log of the manufacturing IP, the unemployment rate, the log of producer price index (PPI) for finished goods and Moody's Baa corporate bond yield relative to the yield on 10-year treasury constant maturity, that is, the Baa corporate credit spread. The left panel of Figure 9 plots the effects of a monetary policy shock. Using the current months fed funds futures surprises as external instrument, we reproduce Caldara and Herbst's (2019) finding that monetary policy shock induces declines in real activity, owing to the inclusion of credit spreads.

To check the sensitivity of this result, we conduct further robustness checks. First, we re-estimate Caldara and Herbst (2019) model but using Gertler and Karadi's (2015) pair of the policy indicator and the instrument, that is, the 1-year government bond rate and the 3-month ahead funds rate surprise (FF4), respectively. The results are reported in the right panel of Figure 9. In contrast to Caldara and Herbst (2019), monetary policy shocks no longer have statistically significant real effects, even when we include credit spread in the model (note that the response of the credit spread variables is also insignificant).[23] We have also estimated this specification using Caldara and Herbst's (2019) Bayesian framework and the results remain essentially the same.[24] Second, we re-estimate the baseline six-variate model of Gertler and Karadi (2015) over Caldara and Herbst's (2019) sample period, that is, 1994:1–2007:6. We also replace the 1-year government bond rate with the federal funds rate and the 3-month ahead futures rate surprise with the contemporaneous surprise in fed fund futures, as in Caldara and Herbst (2019). Figure S.12 in the Appendix S1 shows that monetary policy shocks do not have significant real effects despite significant rise in credit costs.

[23] The finding is similar when we just replace the federal funds rate with the 1-year rate as the policy indicator and continue to use the current month's fed funds futures surprises as the external instrument.

[24] See Figure S.11 in the Appendix S1. Caldara and Herbst (2019) point to several advantages of using the Bayesian framework over the more traditional frequentist inference.

VI Conclusion

This paper revisits the macroeconomic effects of monetary policy shocks using an external instrument identification approach. Using the framework of Gertler and Karadi (2015), we analyse the joint response of financial and economic variables to an identified monetary policy shock. Our results suggest that monetary policy shocks have modest effects on real economic activity in the post-mid-1980s period. This holds despite large movements in credit costs arising due to the policy shock, a finding that deviates from Gertler and Karadi (2015). Using Bernanke *et al.* (1999)'s financial accelerator model, which incorporates the credit channel of monetary policy transmission, we highlight improvements in the conduct of systematic monetary policy during the Great Moderation as a potential explanation for the empirical finding. Accordingly, changes in the dynamics of monetary policy shocks do not necessarily reflect a reduction in policy effectiveness, but rather an improvement in its conduct. Overall, our results are in line with Ramey's (2016) conclusion that 'monetary policy can have big effects, but it is likely that monetary shocks are no longer an important source of macro instability'.

Conflict of Interest

The authors have no competing interests to declare that are relevant to the content of this article.

Supporting Information

Additional Supporting Information may be found in the online version of this article:

Appendix S1

REFERENCES

Anderson, T.W. and Rubin, H. (1949), 'Estimation of the Parameters of a Single Equation in a Complete System of Stochastic Equations', *Annals of Mathematical Statistics*, **20**, 46–63.

Angelini, G., Caggiano, G., Castelnuovo, E. and Fanelli, L. (2023), 'Are Fiscal Multipliers

Estimated with Proxy-SVARs Robust?', *Oxford Bulletin of Economics and Statistics*, **85**, 95–122.
Barth, M.J. and Ramey, V.A. (2001), 'The Cost Channel of Monetary Transmission', *NBER Macroeconomics Annual*, **16**, 199–240.
Bernanke, B.S. and Blinder, A.S. (1992), 'The Federal Funds Rate and the Channels of Monetary Transmission', *American Economic Review*, **82**, 901–921.
Bernanke, B.S. and Gertler, M. (1995), 'Inside the Black Box: The Credit Channel of Monetary Policy Transmission', *Journal of Economic Perspectives*, **9**, 27–48.
Bernanke, B.S., Gertler, M. and Gilchrist, S. (1999), 'The Financial Accelerator in a Quantitative Business Cycle Framework', in Taylor, J.B. and Woodford, M. (eds), *Handbook of Macroeconomics*, Chap. 21, vol. **1c**. Elsevier; 1341–93. Available from: https://www.sciencedirect.com/science/article/pii/S157400489910034X
Blanchard, O. and Simon, J. (2001), 'The Long and Large Decline in U.S. Output Volatility', *Brookings Papers on Economic Activity*, **2001**, 135–174.
Boivin, J. and Giannoni, M. (2002), 'Assessing Changes in the Monetary Transmission Mechanism: A VAR Approach', *Federal Reserve Bank of New York Economic Policy Review*, **8**, 97–111.
Boivin, J. and Giannoni, M. (2006), 'Has Monetary Policy Become More Effective?', *The Review of Economics and Statistics*, **88**, 445–462.
Boivin, J., Kiley, M.T. and Mishkin, F.S. (2010), 'How Has the Monetary Transmission Mechanism Evolved over Time?' in *Handbook of Monetary Economics*, vol. **3**. Elsevier; 369–422. Available from: https://www.sciencedirect.com/science/article/pii/B9780444532381000089
Brüggemann, R., Jentsch, C. and Trenkler, C. (2016), 'Inference in VARs with Conditional Heteroskedasticity of Unknown Form', *Journal of Econometrics*, **191**, 69–85.
Caldara, D. and Herbst, E. (2019), 'Monetary Policy, Real Activity, and Credit Spreads: Evidence from Bayesian Proxy SVARs', *American Economic Journal: Macroeconomics*, **11**, 157–192.
Canova, F. and Gambetti, L. (2009), 'Structural Changes in the US Economy: Is there a Role for Monetary Policy?', *Journal of Economic Dynamics & Control*, **30**, 477–490.
Castelnuovo, E. (2016), 'Modest Macroeconomic Effects of Monetary Policy Shocks During the Great Moderation: An Alternative Interpretation', *Journal of Macroeconomics*, **47**, 300–314.
Castelnuovo, E. and Surico, P. (2010), 'Monetary Policy, Inflation Expectations and the Price Puzzle', *The Economic Journal*, **120**, 1262–83.
Christensen, I. and Dib, A. (2008), 'The Financial Accelerator in an Estimated New Keynesian Model', *Review of Economic Dynamics*, **11**, 155–178.
Christiano, L.J., Eichenbaum, M. and Evans, C. (1996), 'The Effects of Monetary Policy Shocks: Evidence from the Flow of Funds', *The Review of Economics and Statistics*, **78**, 16–34.
Christiano, L.J., Eichenbaum, M. and Evans, C.L. (1999), 'Monetary Policy Shocks: What Have we Learned and to What End?' in Taylor, J.B. and Woodford, M. (eds), *Handbook of Macroeconomics* Chap. 2, vol. **1A**. Elsevier; 65–148. Available from: https://www.sciencedirect.com/science/article/pii/S1574004899010058
Clarida, R., Gali, J. and Gertler, M. (2000), 'Monetary Policy Rules and Macroeconomic Stability: Evidence and some Theory', *The Quarterly Journal of Economics*, **115**, 147–180.
Coibion, O. (2012), 'Are the Effects of Monetary Policy Shocks Big or Small?', *American Economic Journal: Macroeconomics*, **4**, 1–32.
Del Negro, M., Giannone, D., Giannoni, M.P. and Tambalotti, A. (2017), 'Safety, Liquidity, and the Natural Rate of Interest', *Brookings Papers on Economic Activity*, **2017**, 235–316.
Doko Tchatoka, F. and Dufour, J.-M. (2014), 'Identification-Robust Inference for Endogeneity Parameters in Linear Structural Models', *The Econometrics Journal*, **17**, 165–187.
Dufour, J.-M. (1997), 'Some Impossibility Theorems in Econometrics with Applications to Structural and Dynamic Models', *Econometrica*, **65**, 1365–87.
Dufour, J.-M. and Taamouti, M. (2005), '-Projection-Based Statistical Inference in Linear Structural Models with Possibly Weak Instruments', *Econometrica*, **73**, 1351–65.
Galí, J. and Gambetti, L. (2009), 'On the Sources of the Great Moderation', *American Economic Journal: Macroeconomics*, **1**, 26–57.
Gertler, M. and Karadi, P. (2015), 'Monetary Policy Surprises, Credit Costs, and Economic Activity', *American Economic Journal: Macroeconomics*, **7**, 44–76.
Gertler, M. and Lown, C.S. (1999), 'The Information in the High-Yield Bond Spread for the Business Cycle: Evidence and some Implications', *Oxford Review of Economic Policy*, **15**, 132–150.
Gilchrist, S. and Zakrajšek, E. (2012), 'Credit Spreads and Business Cycle Uctuations', *American Economic Review*, **102**, 1692–1720.

Goodfriend, M. and King, R.G. (2005), 'The Incredible Volcker Disinflation', *Journal of Monetary Economics*, **52**, 981–1015.

Gürkaynak, R.S., Sack, B. and Swanson, E.T. (2005), 'Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements', *International Journal of Central Banking*, **1**, 55–93.

Hanson, M. (2008), 'The "Price Puzzle" Reconsidered', *Journal of Monetary Economics*, **51**, 1385–1413.

Hanson, S.G. and Stein, J.C. (2015), 'Monetary Policy and Long-Term Real Rates', *Journal of Financial Economics*, **115**, 429–448.

Haque, Q. (2022), 'Monetary Policy, Inflation Target, and the Great Moderation: An Empirical Investigation', *International Journal of Central Banking*, **18**, 217–258.

Haque, Q., Groshenny, N. and Weder, M. (2021), 'Do we Really Know that US Monetary Policy Was Destabilizing in the 1970s?', *European Economic Review*, **131**, 103615.

Hirose, Y., Kurozumi, T. and Van Zandweghe, W. (2020), 'Monetary Policy and Macroeconomic Stability Revisited', *Review of Economic Dynamics*, **37**, 255–274.

Ikeda, D., Li, S., Mavroeidis, S. and Zanetti, F. (2020). 'Testing the Effectiveness of Unconventional Monetary Policy in Japan and the United States', *arXiv preprint* arXiv:2012.15158.

Jarociński, M. and Karadi, P. (2020), 'Deconstructing Monetary Policy Surprises – The Role of Information Shocks', *American Economic Journal: Macroeconomics*, **12**, 1–43.

Jentsch, C. and Lunsford, K.G. (2019), 'The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States: Comment', *American Economic Review*, **109**, 2655–78.

Jentsch, C. and Lunsford, K.G. (2021), 'Asymptotically Valid Bootstrap Inference for Proxy SVARs', *Journal of Business & Economic Statistics*, **40**, 1–16.

Krishnamurthy, A. and Vissing-Jorgensen, A. (2012), 'The Aggregate Demand for Treasury Debt', *Journal of Political Economy*, **120**, 233–267.

Leeper, E.M., Sims, C.A., Zha, T., Hall, R.E. and Bernanke, B.S. (1996), 'What Does Monetary Policy Do?', *Brookings Papers on Economic Activity*, **1996**, 1–78.

Leeper, E.M. and Zha, T. (2003), 'Modest Policy Interventions', *Journal of Monetary Economics*, **50**, 1673–1700.

Liu, P., Theodoridis, K., Mumtaz, H. and Zanetti, F. (2019), 'Changing Macroeconomic Dynamics at the Zero Lower Bound', *Journal of Business & Economic Statistics*, **37**, 391–404.

Lubik, T.A., Matthes, C. and Sablik, T. (2016). 'The Burns Disinflation of 1974', Richmond Fed Economic Brief.

Lubik, T.A. and Schorfheide, F. (2004), 'Testing for Indeterminacy: An Application to US Monetary Policy', *American Economic Review*, **94**, 190–217.

Lütkepohl, H. (1990), 'Asymptotic Distributions of Impulse Response Functions and Forecast Error Variance Decompositions of Vector Autoregressive Models', *Review of Economics and Statistics*, **72**, 116–125.

Lütkepohl, H. (2007), 'General-to-Specific or Specific-to-General Modelling? An Opinion on Current Econometric Terminology', *Journal of Econometrics*, **136**, 319–324.

McConnell, M.M. and Perez-Quiros, G. (2000), 'Output Fluctuations in the United States: What Has Changed since the Early 1980's?', *American Economic Review*, **90**, 1464–76.

Mertens, K. and Montiel Olea, J.L. (2018), 'Marginal Tax Rates and Income: New Time Series Evidence', *The Quarterly Journal of Economics*, **133**, 1803–84.

Mertens, K. and Ravn, M.O. (2013), 'The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States', *American Economic Review*, **103**, 1212–47.

Mertens, K. and Ravn, M.O. (2014), 'Fiscal Policy in an Expectations-Driven Liquidity Trap', *The Review of Economic Studies*, **81**, 1637–67.

Mertens, K. and Ravn, M.O. (2019), 'The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States: Reply', *American Economic Review*, **109**, 2679–91.

Miranda-Agrippino, S. and Ricco, G. (2021), 'The Transmission of Monetary Policy Shocks', *American Economic Journal: Macroeconomics*, **13**, 74–107.

Mojon, B. (2008), 'When Did Unsystematic Monetary Policy Have an Effect on Inflation?', *European Economic Review*, **52**, 487–497.

Montiel Olea, J., Stock, J. and Watson, M. (2021), 'Inference in Structural Vector Autoregressions Identified with an External Instrument', *Journal of Econometrics*, **225**, 74–87.

Nagel, S. (2016), 'The Liquidity Premium of near-Money Assets', *The Quarterly Journal of Economics*, **131**, 1927–71.

Primiceri, G.E. (2005), 'Time Varying Structural Vector Autoregressions and Monetary Policy', *Review of Economic Studies*, **72**, 821–852.

Ramey, V.A. (2016), 'Macroeconomic Shocks and Their Propagation', *Handbook of Macroeconomics*, **2**, 71–162.

Romer, C. and Romer, D. (2004), 'A New Measure of Monetary Shocks: Derivation and Implications', *American Economic Review*, **94**, 1055–84.

Sims, C.A. (1980), 'Macroeconomics and Reality', *Econometrica*, **48**, 1–48.

Staiger, D. and Stock, J. (1997), 'Instrumental Variables Regression with Weak Instruments', *Econometrica*, **65**, 557–586.

Stock, J. and Watson, M. (2002), 'Has the Business Cycle Changed and Why?', *NBER Macroeconomics Annual*, **17**, 159–218.

Stock, J. and Watson, M. (2012), 'Disentangling the Channels of the 2007–09 Recession', *Brookings Papers on Economic Activity*, **43**, 81–156.

Stock, J. and Watson, M. (2018), 'Identification and Estimation of Dynamic Causal Effects in Macroeconomics Using External Instruments', *The Economic Journal*, **128**, 917–948.

Swanson, E.T. and Williams, J.C. (2014), 'Measuring the Effect of the Zero Lower Bound on Mediumand Longer-Term Interest Rates', *American Economic Review*, **104**, 3154–85.

Taylor, J.B. (1993), 'Discretion Versus Policy Rules in Practice', *Carnegie-Rochester Conference Series on Public Policy*, **39**, 195–214.

Van Binsbergen, J.H., Diamond, W.F. and Grotteria, M. (2022), 'Risk-Free Interest Rates', *Journal of Financial Economics*, **143**, 1–29.