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Keywords

Monetary policy shocks, Proxy-SVAR, Weak identification, Output Dynamics

JEL Classification

E31, E32, E43, E44, E52

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Revisiting the macroeconomic effects of monetary policy shocks*

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Abstract

We shed new light on the effects of monetary policy shocks in the US. Gertler and Karadi (2015) suggest that movements in credit costs may result in substantial impact of monetary policy shocks on economic activity. Using the proxy SVAR framework, we show that once the Volcker disinflation period is left out and one focuses on the post-1984 period, monetary policy shocks have no significant effects on output, despite large movements in credit costs. Our finding is robust to weak identification and alternative measure of economic activity.

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1 Introduction

Since the seminal contribution of Sims (1980), vector autoregressions (VARs) have emerged as a modelling tool to study the effects of structural shocks. There is now a growing literature analyzing the effects of monetary policy shock on economic activity using VARs; see e.g., Christiano et al. (1999) for an early survey of this literature. However, recent studies suggest that the VAR-based characterization of monetary policy is sensitive to the choice of the framework, identification strategy and sample period; see e.g., Boivin et al. (2010). In this paper, we revisit the macroeconomic effects of monetary policy shock using a proxy SVAR approach, first introduced by Stock (2008). This approach identifies the effects of monetary policy shock using external instruments, and has been applied in recent studies; see e.g., Stock and Watson (2012), Mertens and Ravn (2013, 2014), Gertler and Karadi (2015), and Mertens and Montiel Olea (2018).

Our focus on proxy SVAR setting is mainly motivated by Gertler and Karadi (2015), who suggest the need to include financial variables (such as credit spread) in order to capture the transmission mechanism of monetary policy to the aggregate economy through the credit channel, and the need to identify policy surprises that are exogenous to both the economic and financial variables in the VAR. Gertler and Karadi (2015) argue that standard identification strategy based on timing restrictions is inappropriate if financial variables are included in the VAR due to the problem of simultaneity.[1] The proxy SVAR framework addresses this simultaneity issue by using high frequency identification (HFI) measures of policy surprises as external instruments. Using these high frequency surprises on Federal Open Market Committee (FOMC) policy announcement dates, Gertler and Karadi (2015) find that monetary policy shocks produce large movements in credit costs accompanied by substantial impact on economic activity.

[1] Within a period, policy shifts may not only influence financial variables, but they may be responding to them as well.

In this paper, we first show that Gertler and Karadi's finding is sensitive to the choice of the sample period. Monetary policy shocks have no statistically significant effect on output during the post-mid-1980s, particularly when the "Volcker disinflation" period is left out from the sample. In contrast, a statistically significant decline in output is observed when the Volcker disinflation period is included. This illustrates a notable difference compared to Gertler and Karadi (2015). An important implication is that widely used hump-shaped response of economic activity provides poor guidance in evaluating the efficacy of structural models and for implementing monetary policy. Second, we show that the proxy SVAR approach using federal funds rate futures surprises around policy announcements as external instruments do not lead to a weak instruments problem, as the identification-robust proxy SVAR method (see Olea et al., 2020) yields confidence sets for the impulse response functions that are similar to their standard strong instrument-based counterparts. Nonetheless, futures surprises are not strong instruments in general. For instance, when using the futures surprises in three month Eurodollar deposits, the identification-robust confidence sets are much wider than the standard ones, thereby suggesting that such instruments are not very strong (see Dufour, 1997).

The ambiguity of the VAR-based approach to identify the transmission of monetary policy shocks to the aggregate economy is well known. Early studies that employed recursive VAR identification approach argue that the response of economic activity to a monetary policy shock is hump-shaped; see e.g., Christiano et al. (1999), who show that output, consumption, and investment all have hump-shaped responses to a monetary policy shock, peaking after about one and half years and returning to their pre-shock levels after about three years. These stylized facts from Christiano et al. (1999) are regularly used in the literature to evaluate the predictions of structural dynamic stochastic general equilibrium (DSGE) models. However, the post-2000 literature on VAR-based characterization of the effects of monetary policy shocks on economic activity provides mixed evidence about the changes in the transmission of monetary policy to the aggregate economy. Boivin and Giannoni (2002, 2006) estimate a VAR over the pre- and post-Volcker periods and identify monetary policy shock using a recursive identification scheme. They find that during the post-Volcker period, the response of output was only about one-quarter of that during the pre-Volcker period. Primiceri (2005), Canova and Gambetti (2009) and Galí and Gambetti (2009) estimate time-varying parameter VARs to characterize the changing transmission of monetary policy. Galí and Gambetti (2009) find that the response of inflation and real activity to demand-type shocks have fallen over time, although they do not separately identify a policy shock. Unlike Galí and Gambetti (2009), Primiceri (2005) find little change in the dynamics of monetary policy shocks in the post-war period. Canova and Gambetti (2009) use sign restrictions to show that the transmission of monetary policy shocks has been relatively stable over time, with stronger effects being observed on inflation and real activity in the post-1990 period. Boivin et al. (2010) suggest that such discrepancy could arise from the way the monetary policy shock is identified. While Canova and Gambetti (2009) leave the impact response of real activity unrestricted, it is constrained to be zero under the recursive identification scheme in Boivin and Giannoni (2002, 2006). Moreover, Boivin et al. (2010) argue that the sign restrictions used by Canova and Gambetti (2009) only produce set identification – their reported impulse response functions are not due to a pure monetary policy shock but rather to a combination of structural shocks including monetary policy. Using the factor-augmented VAR (FAVAR) setting (thus avoiding the omission of potentially important variables), Boivin et al. (2010) find that monetary policy innovations have a muted effect on real activity and inflation in the post-1980 period.

Given this ambiguity, we revisit the macroeconomic effects of monetary policy shocks based on the framework of Gertler and Karadi (2015), who find that monetary policy shocks identified using high frequency surprises around policy announcements as external instruments induce large movements in credit costs accompanied by substantial impact on output. However, we argue that the effects of monetary policy shocks on economic activity depend on the sample period. In particular, we find the effect is statistically insignificant once the Volcker disinflation period is excluded, despite large movements in credit costs. Our finding is in line with Coibion (2012) who, using the Romer and Romer (2004) approach[2], show 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. 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.[3] Mojon (2008) argues that hump-shaped VAR-estimated response of inflation to a monetary policy shock is, therefore, not appropriate to fit stylized models of the response of inflation around a stable trend inflation. Likewise, our result suggests that hump-shaped VAR-estimated response of real economic activity is not appropriate either in evaluating the efficacy of structural models.

A growing body of studies find modest macroeconomic effects of monetary policy shocks during the Great Moderation (Gertler and Lown, 1999; Barth and Ramey, 2001; Hanson, 2004; Boivin and Giannoni, 2002, 2006; Mojon, 2008; Boivin et al., 2010; Castelnuovo, 2016). The literature provides several interpretation of such modest effects including the role played by milder volatility of output (Hanson, 2004), technological progress or financial innovations (Boivin and Giannoni, 2006; Boivin et al., 2010; Castelnuovo, 2016), and a more hawkish monetary policy stance (Boivin and Giannoni, 2006; Boivin et al., 2010). However, most empirical VAR studies documenting modest macroeconomic effects of monetary policy shocks rely on the Cholesky identification scheme, which is known to work poorly when both economic and financial variables are included in the VAR due to the simultaneity issue; see e.g., Gertler and Karadi (2015). Moreover, Castelnuovo (2016) illustrates that such modest effects may as well be an artefact of the timing restrictions in the Cholesky identification scheme. Nonetheless, our results suggest that modest and statistically insignificant response

[2] Romer and Romer (2004) identify monetary policy innovations by first constructing a historical series of interest rate changes decided upon at the Federal Open Market Committee 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.

[3] We also find statistically insignificant response of inflation and so do Gertler and Karadi (2015).

of output and inflation to a monetary policy shock are salient empirical facts of the post-1984 U.S. economy.

The rest of the paper is organized as follows. Section 2 presents the SVAR-IV framework and discusses how external instruments can be used to identify monetary policy shocks. Section 3 presents the results with both the standard proxy SVAR approach and recently developed weak identification-robust method (Olea et al., 2020). Additional robustness checks are performed in Section 4. Section 5 concludes. Further empirical results are available in the Appendix.

2 Econometric framework

2.1 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[4] and follows a stationary p-order structural vector autogression with underlying reduced-form representation:

$$Y_t = \sum_{j=1}^{p} B_j Y_{t-j} + u_t, \tag{2.1}$$

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:

$$u_t = A_0 \varepsilon_t, \tag{2.2}$$

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

$$E(\varepsilon_t) = 0 \text{ and } E(\varepsilon_t \varepsilon_t') = D = \text{diag}(\sigma_1^2, \ldots, \sigma_n^2). \tag{2.3}$$

[4] See the description of the variables in Section 2.2.

From (2.3), the variance-covariance matrix of the reduced form innovations is

$$E\left(u_t u_t'\right) := \Sigma = A_0 D A_0'. \tag{2.4}$$

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

$$Y_t = \sum_{k=0}^{\infty} C_k\left(B\right) A_0 \varepsilon_{t-k}, \tag{2.5}$$

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

$$C_k\left(B\right) = \sum_{m=1}^{k} C_{k-m}\left(B\right) B_m, \quad k = 1, 2, \ldots \tag{2.6}$$

with $C_0\left(B\right) = I_n$ 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

$$\partial Y_{i,t+k} / \partial \varepsilon_t^j = e_i' C_k\left(B\right) A_0 e_j, \tag{2.7}$$

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 (2.7).

Following Gertler and Karadi (2015), we use the one-year government bond rate as the policy indicator, which is associated with exogenous variations due to the structural monetary policy shock ε_t^{mp}. Gertler and Karadi (2015) argue that in order 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^*, i.e., the following assumption (similar to Gertler and Karadi, 2015) must be satisfied:

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

Assumption 1-(a) states the relevance of z_t as an instrument for ε_t^{mp} while Assumption 1-(b) implies that it is a valid (or an exogenous) instrument. Stock and Watson (2018) and Olea et al. (2020) show that if Assumption 1 holds, 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 (2.1)-(2.7), and conduct statistical inference on the identified impulse response coefficients as discussed below.

Identification of the impulse response coefficients. From (2.7), 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 Assumption 1, i.e.

$$\Theta = E\left[z_t A_0 \varepsilon_t\right] = \alpha A_{0,1}. \tag{2.8}$$

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

$$\lambda_{k,i} = \frac{e_i' C_k(B)\Theta}{e_1'\Theta}. \tag{2.9}$$

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:

$$\text{Proj}(z_t|u_t) \;=\; \Theta'\Sigma^{-1}u_t = (\alpha A_0 e_1)'(A_0 D A_0')^{-1}A_0\varepsilon_t = (\alpha/\sigma_1^2)\varepsilon_t^{mp}. \tag{2.10}$$

The projection (2.10) determines ε_t^{mp} up to the scale factor (α/σ_1^2); dividing by $(\Theta'\Sigma^{-1}\Theta)^{1/2}$ 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 (2.1). 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 $X_t = (1, Y_{t-1}', Y_{t-2}', \ldots, Y_{t-p}')'$. 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 (2.9).

Inference about $\lambda_{k,i}$. The plug-in estimator and δ-method type confidence set for $\lambda_{k,i}$ are derived from :

$$\hat{\lambda}_{k,i}\left(\hat{B}_T, \hat{\Theta}_T\right) = e_i' C_k(\hat{B}_T)\hat{\Theta}_T / e_1'\hat{\Theta}_T, \tag{2.11}$$

and

$$\mathscr{C}_\beta^{Plug-in}\left(\lambda_{k,i}\right) = \left(\lambda_{k,i} \middle| \frac{T\left(\hat{\lambda}_{k,i}\left(\hat{B}_T, \hat{\Theta}_T\right) - \lambda_{k,i}\right)^2}{\hat{\sigma}_{T,k,i}^2} \le \chi_{1,1-\beta}^2\right), \tag{2.12}$$

respectively, where $\chi_{1,1-\beta}^2$ is the $1-\beta$ percentile of the χ_1^2 distribution, $\hat{\sigma}_{T,k,i}^2$ is a consistent estimator of $\sigma_{k,i}^2$, which in turn depends on the limiting variance for the estimators $\left(\hat{B}_T, \hat{\Theta}_T\right)$ and the gradient of $\lambda_{k,i}\left(B, \Theta\right)$ with respect to (B, Θ), and can be obtained using the δ-method or a suitable bootstrap procedure.[5] Provided that the instrument z_t is strong, $\mathscr{C}_\beta^{Plug-in}\left(\lambda_{k,i}\right)$ has level $1-\beta$ asymptotically; see Olea et al. (2020).

2.2 Data and estimation

We consider the setting of Gertler and Karadi (2015) where the underlying reduced-form VAR in (2.1) contains six variables: the logarithm of industrial production index (IP), the logarithm of consumer price index (CPI), the one-year government rate, the excess bond premium, the mortgage spread, and the commercial paper spread. Real economic activity is measured by the industrial production index[6] and inflation is measured by the CPI. Mortgage spread is calculated as the difference between the 30-year mortgage rate and the 10-year

[5] Olea et al. (2020) show that $\sqrt{T}\left[\hat{\lambda}_{k,i}\left(\hat{B}_T, \hat{\Theta}_T\right) - \lambda_{k,i}\left(B, \Theta\right)\right]$ converges in distribution to $N\left(0, \sigma_{k,i}^2\right)$.

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

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. We use the one-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 and Karadi, 2015), which 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 three 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 prior to the FOMC meeting, the surprise in the futures rate can be expressed as[7]:

$$surprise = f_{t+j} - f_{t+j,-1}. \tag{2.13}$$

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

[7] 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.

decision itself, the surprises in futures rates are measured within a tight window of thirty minutes as in Gürkaynak et al. (2005). Gertler and Karadi (2015) show that the strength of the "surprise" instrument is not low (see Gertler and Karadi, 2015, Table 3). They also argue that the use of fed funds futures surprises for contracts that expire in the future, e.g. three month ahead in this case, captures shocks to forward guidance in the surprise instrument measure.

We estimate the VAR with 12 lags using monthly data.[8] 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). Notice that the "surprise" instrument FF4 is only available from 1990:1. So, as in Gertler and Karadi (2015), we use the full sample to estimate the VAR lag coefficients and obtain the reduced-form residuals in 2.2. We then use these reduced-form residuals and the instrument for the overlapping period to identify the contemporaneous impact of policy shocks.[9]

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 has become 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. There is also strong evidence of structural breaks in the mid-1980s with a marked decline in macroeconomic volatility (see e.g. Blanchard and Simon, 2001; McConnell and Perez-Quiros, 2000; Stock and Watson, 2002). In particular, McConnell and Perez-Quiros (2000) show the existence of a significant break in output growth volatility in 1984. Therefore, looking at the effects of monetary policy in the post-1984 period (i.e., after the end of the Volcker disinflation) is of paramount importance, as also suggested by Mojon (2008).

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

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

3 Results

3.1 Main results

Figure 1 shows the responses of the variables in the baseline model to a one standard deviation monetary tightening shock using the standard SVAR-IV approach. The confidence bounds are constructed using the wild bootstrap as in Gertler and Karadi (2015).[10] The left column shows the variables' responses for Gertler and Karadi's sample (1979:7-2012:6), while the right column illustrates the responses for the post-1984 sample (1984:1-2012:6). Clearly, the left column replicates Figure 2 in Gertler and Karadi (2015).

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 striking difference: *the response of output (here, IP) is statistically insignificant*. This result points to a notable difference for the effect of monetary policy shock on output than the one documented in Gertler and Karadi (2015), who argue that movements in credit costs may explain why monetary policy shocks substantially impact economic activity. However, our results indicate that once the "Volcker disinflation" period is left out from the estimated sample, the response of output becomes statistically insignificant despite rise in credit costs.

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% in 1980 to 4% in 1983. Monetary policy during this period is better characterized by nonborrowed-reserve targeting (rather than interest rate rules) and aggressive increases in interest rates. The economy experienced two recessions generally attributed to disinflationary monetary policy. While far less than predicted, the output losses were substantial. Therefore, both the

[10] 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.

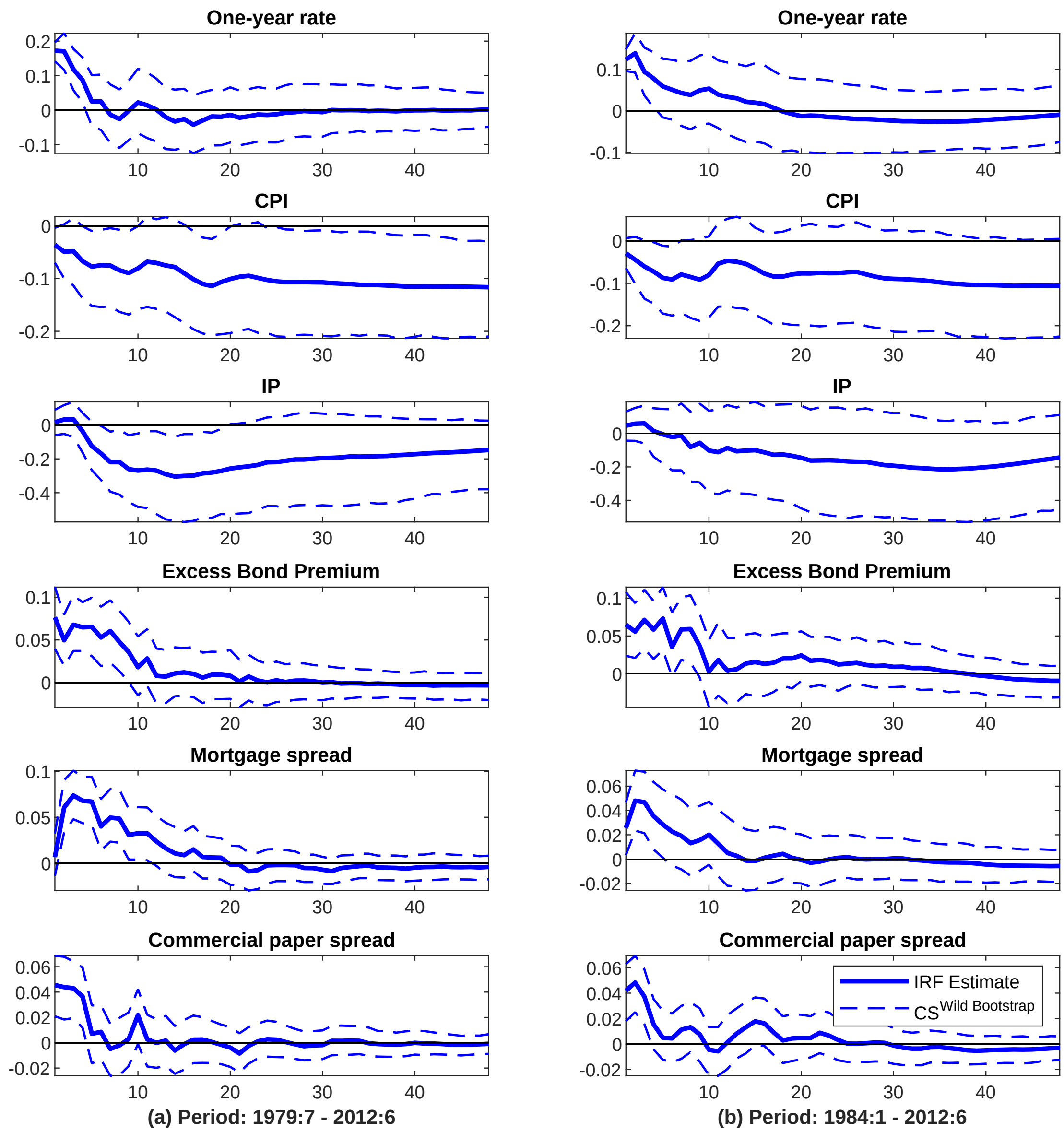

Figure 1: Impulse responses to a one-year rate shock

operating procedures of the Federal Reserve and the dynamics of the economy were different during the Volcker disinflation. Our finding suggest that the estimated VAR in Gertler and Karadi's sample may have mixed up the response of output to monetary policy shock in periods of large swings in economic activity and inflation (such as the Volcker disinflation) and periods when inflation and output are relatively more stable (such as the post-1984 period). Therefore, an important implication of our result is that estimating the VAR using data from the Volcker disinflation period does not approximate well the real effects of monetary policy shocks.

Our finding is in line with the literature, in particular the empirical studies on VARs that often find modest macroeconomic effects of monetary policy shocks during the Great Moderation.[11] Various interpretations have been put forward regarding such modest macroeconomic effects of monetary policy shocks. For example, Hanson (2004) discusses the role played by milder volatility of output during the Great Moderation. Castelnuovo and Surico (2010) and Castelnuovo (2016), among others, point toward technological progress or financial innovations easing households' consumption smoothing. Boivin and Giannoni (2006) and Boivin et al. (2010) point toward aggressive monetary policy response to inflation as the reason for the moderate macroeconomic reactions in the post-1984 period. Mojon (2008) shows that persistent hump-shaped response of inflation disappears when one examines periods without large shifts in the level of inflation. Along the same lines, we find that the effect of monetary policy shocks on real economic activity disappears when the Volcker disinflation period is left out from our sample.

Notwithstanding the evidence, most empirical VAR studies rely on timing restrictions to identify the monetary policy shock. A maintained assumption within this setting is that the policy rate responds to all macroeconomic variables in the VAR within the period but not the other way around. As such, the recursive or Cholesky identification scheme is often adopted. However, Castelnuovo (2016) documents that modest macroeconomic effects of monetary policy shocks may as well be an artefact of the Cholesky identification scheme. Specifically, Castelnuovo (2016) documents through a Monte Carlo experiment that the zero restrictions imposed by the Cholesky identification on the impact response of inflation and output distort the estimated policy shocks. In fact, the monetary policy shock turns out to be a linear combination of genuine monetary surprises and non-policy shocks. Castelnuovo (2016) argues that imposing a recursive identification scheme leads to picking up a combination of demand

[11] See e.g. Gertler and Lown (1999); Barth and Ramey (2001); Hanson (2004); Boivin and Giannoni (2002, 2006); Mojon (2008); Castelnuovo and Surico (2010); Boivin et al. (2010); Castelnuovo (2016).

and supply shocks which induce a modest response on the net effect on inflation and output. However, our findings suggest that the modest effects of monetary policy shocks on inflation and output in the U.S. are likely to be salient stylized facts of the post-1984 period, as the proxy SVAR approach we use does not impose zero restrictions on the contemporaneous response of inflation and output. As mentioned previously, in our proxy SVAR setting identification comes entirely from the external "surprise" instrument.

3.2 Identification-robust inference

The SVAR-IV approach of Section 2.1 yields consistent estimates of the VAR parameters and the impulse responses only if the instrument z_t is strong and valid for ε_t^{mp}, i.e., if Assumption 1 is satisfied. Recent literature (see e.g. Olea et al., 2020) emphasizes the possibility that the external instrument may be weakly correlated with the target structural shock, thus biasing the standard SVAR-IV estimates of the impulse responses. Gertler and Karadi (2015) also point to this issue of instrument strength and show that the baseline instrument FF4 is not weak when used in conjunction with the one-year rate as the policy indicator to identify the monetary policy shock. However, their first-stage F statistics indicate that most of their other instruments, which all come from the event study analysis of Gürkaynak et al. (2005), are weak, particularly when used together with the two-year government bond rate as the policy indicator.[12] Notwithstanding this evidence, Gertler and Karadi (2015) do not investigate the consequences of using weak instruments for the standard confidence sets for the impulse responses to monetary policy shocks. Given recent developments on identification-robust methods within the proxy SVAR framework (see e.g. Olea et al., 2020), we first investigate the robustness of our results to weak identification and then illustrate the extent to which some of Gürkaynak et al.'s (2005) surprises variables are uninformative as external instruments in identifying monetary policy shocks.

[12] Gertler and Karadi (2015) suggest that the two-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 two year horizon.

Olea et al. (2020) show how small values of $e_1'\Theta$ in the denominator of (2.9), which arise when $cov(z_t, \varepsilon_t^{mp}) = \alpha$ is small, may lead to poor coverage of the confidence set $\mathscr{C}_\beta^{Plug-in}(\lambda_{k,i})$ in (2.12). They propose a weak IV-robust method based on Anderson and Rubin (1949, AR-statistic) to build confidence sets for the impulse response coefficients. Specifically, using similar notations as in Section 2.1, define

$$H_T = \begin{pmatrix} e_i' C_k(B)\,\Theta_T \\ e_1'\Theta_T \end{pmatrix} \equiv \begin{pmatrix} H_{T,1} \\ H_{T,2} \end{pmatrix}, \tag{3.1}$$

where $\Theta_T = \alpha_T A_{0,1}$, $E[z_t \varepsilon_t^{mp}] = \alpha_T \to \alpha$ as $T \to \infty$, and $\alpha = 0$ is allowed. This framework allows for 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 (2.9), the impulse response coefficient can be written as $\lambda_{k,i} = H_{T,1}/H_{T,2}$. Olea et al. (2020) show that $\hat{H}_T \overset{a}{\sim} N\left(H_T, T^{-1}\Omega\right)$, 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 $\lim\limits_{T\to\infty} H_T$ with respect to (B,Θ), and W is the asymptotic variance of $\sqrt{T}\left[vec\left(\hat{B}_T - B\right)', \hat{\Theta}_T - \Theta_T\right]'$. 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 $H_{T,2}$, i.e., $H_0 : H_{T,1} - \lambda_0 H_{T,2} = 0$. To assess H_0, Olea et al. (2020) propose to use the Wald statistic

$$q_T(\lambda_0) = \frac{T\left(\hat{H}_{T,1} - \lambda_0 \hat{H}_{T,2}\right)^2}{\hat{\omega}_{T,11} - 2\lambda_0\hat{\omega}_{T,12} + \lambda_0^2\hat{\omega}_{T,22}}, \tag{3.2}$$

where $\hat{\omega}_{T,ij}$ are consistent estimates of the elements of Ω. Under H_0, $q_T(\lambda_0) \xrightarrow{d} \chi_1^2$, 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 and Taamouti, 2005; Doko Tchatoka and Dufour, 2014; Olea et al., 2020):

$$\mathscr{C}_\beta^{AR}(\lambda_{k,i}) = \left\{\lambda_{k,i} : q_T(\lambda_{k,i}) \leq \chi_{1,1-\beta}^2\right\}. \tag{3.3}$$

Figure 2 shows the estimated impulse responses and the corresponding 95% confidence interval with the plug-in method (i.e., $\mathscr{C}_{\beta}^{Plug-in}(\lambda_{k,i})$ given in (2.12)) and the weak-IV robust method (i.e., $\mathscr{C}_{\beta}^{AR}(\lambda_{k,i})$ given in (3.3)). To enable comparison with Figure 1, the size of the monetary policy shock is normalized so that the response of the one-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 2 are based on the δ-method as in Olea et al. (2020).[13] From the plots in Figure 2, we can make two important observations.

First, we note that the weak-instrument robust confidence sets (gray areas) mostly coincide with their standard plug-in counterparts (areas delimited by the two blue dashed lines), thus corroborating Gertler and Karadi's (2015) finding that the three month ahead federal funds futures surprise instrument ($FF4$) is not weak. Second, the impulse responses in Figure 2 are essentially the same as those in Figure 1. Both the $\mathscr{C}_{\beta}^{Plug-in}(\lambda_{k,i})$ and $\mathscr{C}_{\beta}^{AR}(\lambda_{k,i})$ confidence sets in Figure 2 show modest and statistically insignificant effects of monetary policy shocks on inflation and output despite substantial increase in credit spreads, thereby confirming our previous results.

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 six month, nine month and one year ahead futures surprises in three month Eurodollar deposits (henceforth ED2, ED3 and ED4, respectively).[14] The impulse responses are mostly similar to those reported in Figure 2 when FF1 and ED2 are used as external instruments, and also the $\mathscr{C}_{\beta}^{Plug-in}(\lambda_{k,i})$ and $\mathscr{C}_{\beta}^{AR}(\lambda_{k,i})$ 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 3 shows the impulse responses when ED4 is used as the external instrument. As seen in the figure, there are substantial discrepancies between the $\mathscr{C}_{\beta}^{Plug-in}(\lambda_{k,i})$ and $\mathscr{C}_{\beta}^{AR}(\lambda_{k,i})$ confidence sets,

[13] A bootstrap method can also be used; see (see Olea et al., 2020, Appendix A.4) .
[14] The results are reported in Figures 8-10 in the Appendix.

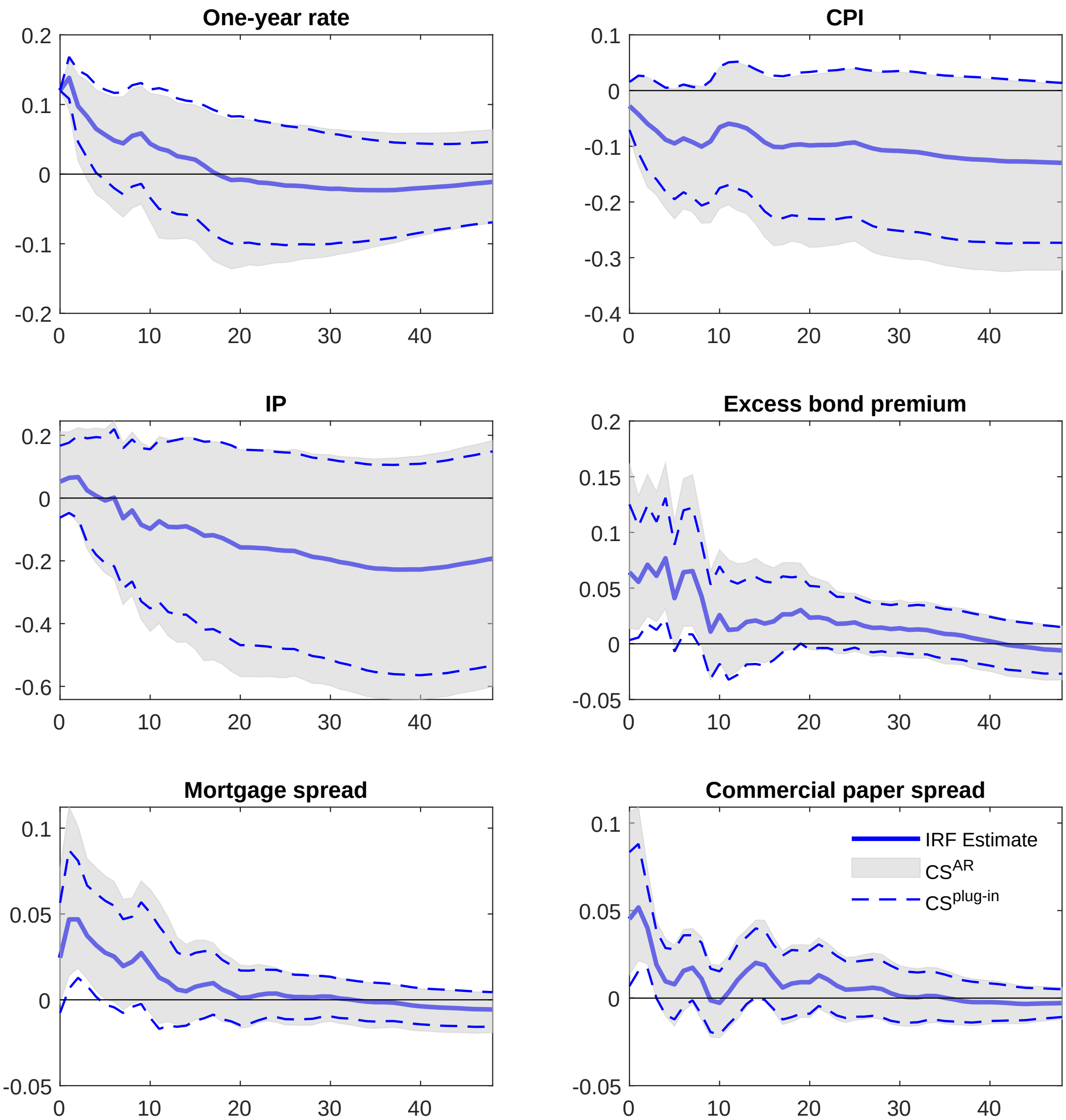

Figure 2: Weak-instrument robust impulse responses to a one-year rate shock; Period: 1984:1-2012:6

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 3 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% AR-type confidence sets show statistically insignificant effects of monetary policy shocks on inflation and output.

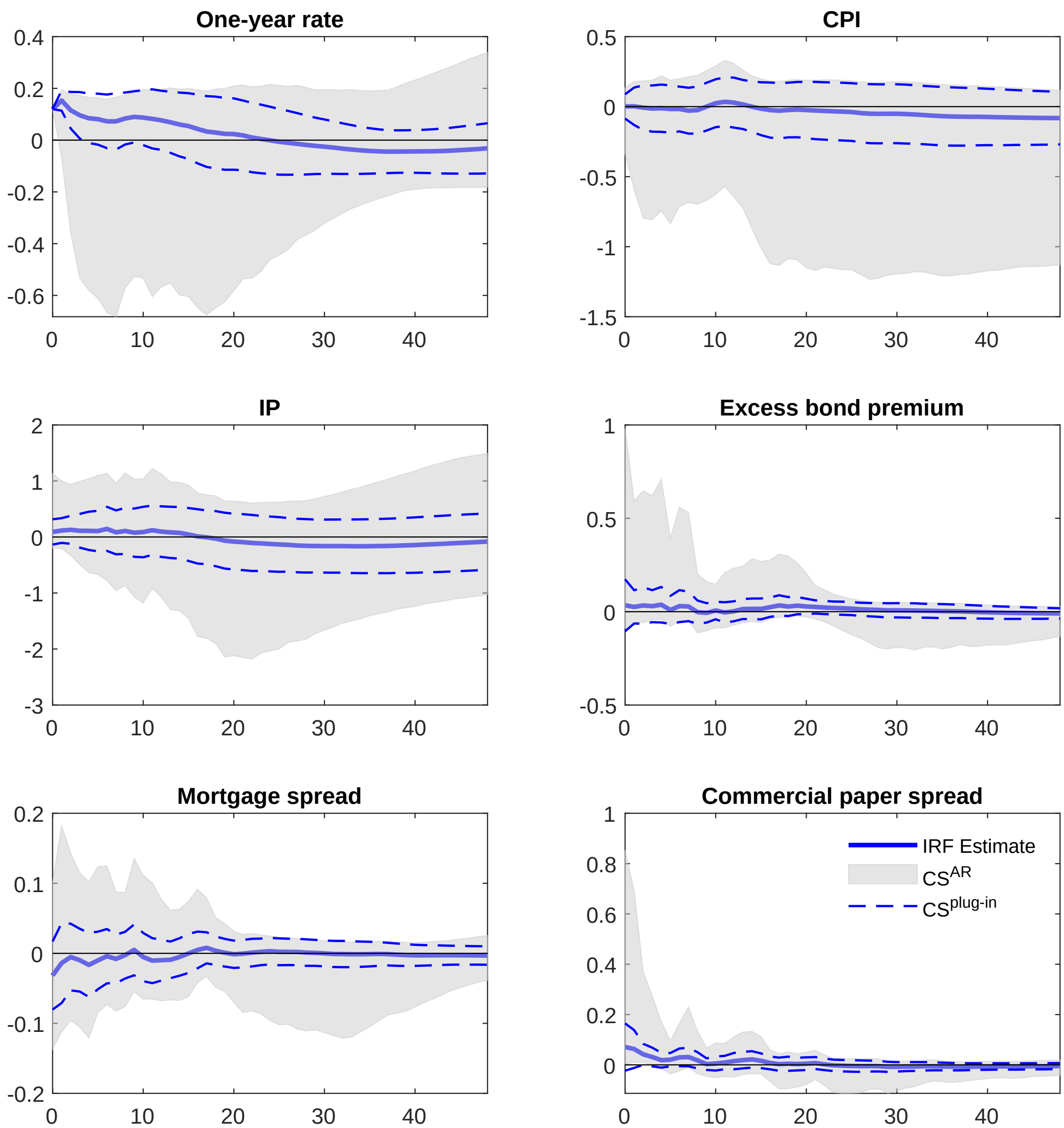

Figure 3: Impulse responses to a one-year rate shock with ED4 as instrument; Period: 1984:1-2012:6

4 Robustness checks

In this section, we conduct various robustness checks with respect to the measure of economic activity, the policy indicator, and the sample period.

4.1 Alternative measure of economic activity

The unemployment rate is a commonly used indicator of real economic activity in the empirical VAR literature. Therefore, we check the robustness of our results to this alternative measure of real economic activity. For this, we replace the industrial production index in the baseline VAR (2.1) with the unemployment rate. As in Section 3.1, the change in three month ahead federal funds futures (FF4) is used as external instrument to identify the effects of monetary policy shocks. Figure 4 shows the impulse responses to a one standard deviation monetary tightening shock, where the size of the shock is normalized such that the impact response of the one-year rate is the same as in the baseline estimation. As seen in this figure, the responses of CPI, the one-year bond rate, and the spread variables are similar to those observed in Figure (2). In particular, there is a decline in the CPI but the response is relatively modest, thus aligning with theory. With regards to the unemployment rate, although it goes up as expected, the response is not significantly different from zero, despite significant increases in credit costs.

4.2 Alternative policy indicator

Next we investigate the robustness of our results to using an alternative policy indicator. Swanson and Williams (2014), Hanson and Stein (2015), and Gertler and Karadi (2015) all suggest that the Federal Reserve's forward guidance strategy operates with roughly two year horizon. As such, we now use the two-year government bond rate as the policy indicator and estimate a seven-variate VAR. As in the baseline model, the change in the three month ahead federal funds futures (FF4) is used as external instrument to identify the effects of monetary policy shocks. The size of the two-year rate shock is normalized such that the impact response on the one-year rate is similar to the one in the baseline estimation. Figure 5 contains the results. As seen in the figure, the two-year rate shock produces responses similar to the baseline model. Once again we observe modest macroeconomic effects of monetary policy shocks on inflation and output. We also estimate the VAR by replacing FF4, one at a

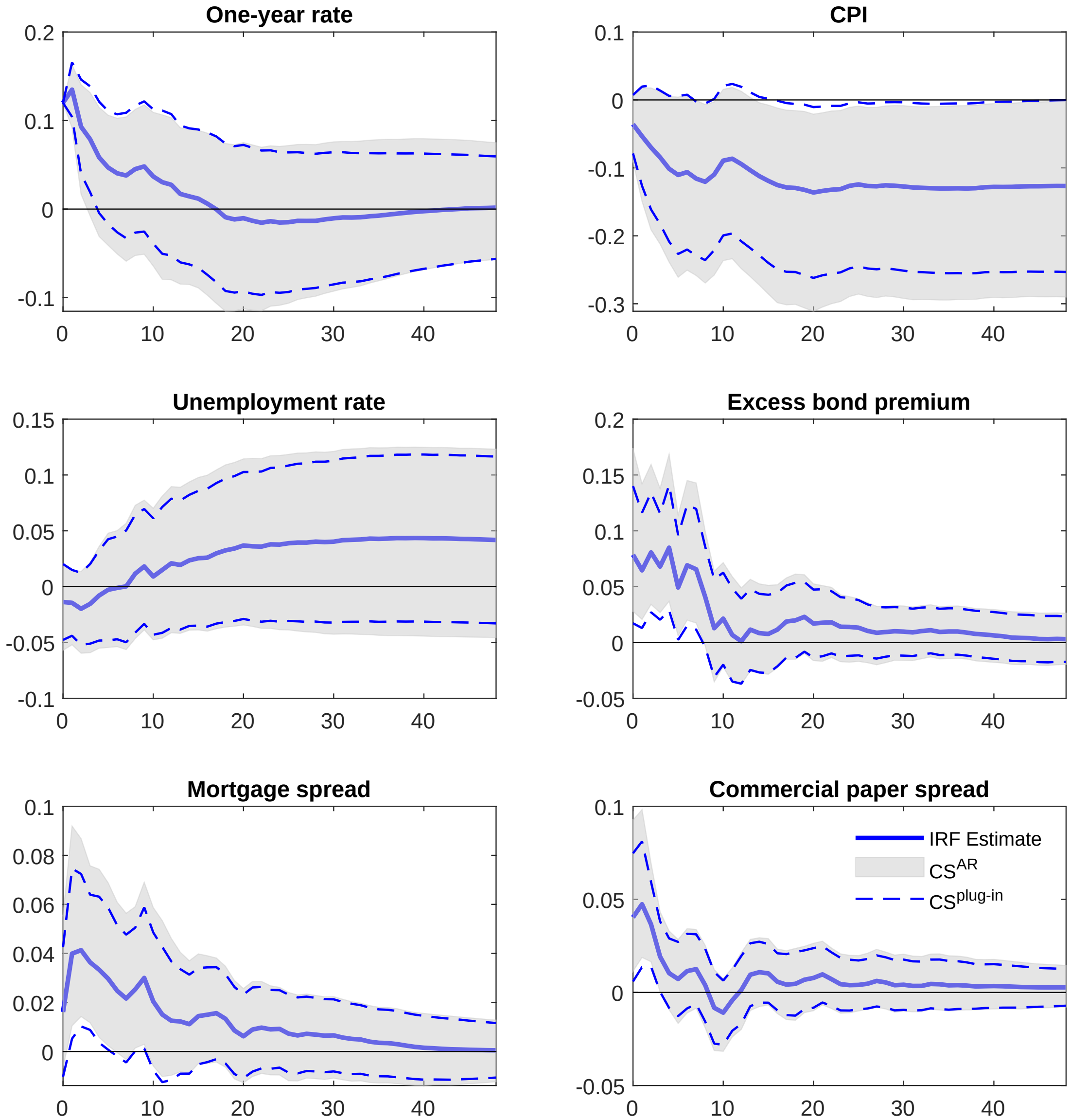

Figure 4: Impulse responses to a one-year rate shock with alternative economic activity indicator; Period: 1984:1-2012:6

time, with FF1, ED2, ED3 and ED4 as the external instrument. These results are reported in Figures 11-14 in the Appendix. 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 when FF4 is used. However, ED3 and ED4 remain weak even with the two-year bond rate as the policy indicator, which is reflected in the wide[15] weak-IV

[15] See Dufour (1997).

robust confidence sets for the impulse responses (see Figures 13 and 14 in the Appendix for ED3 and ED4, respectively). In fact, the identification-robust confidence sets for ED4 turn out to be completely unbounded suggesting that ED4 is completely uninformative as external instrument.

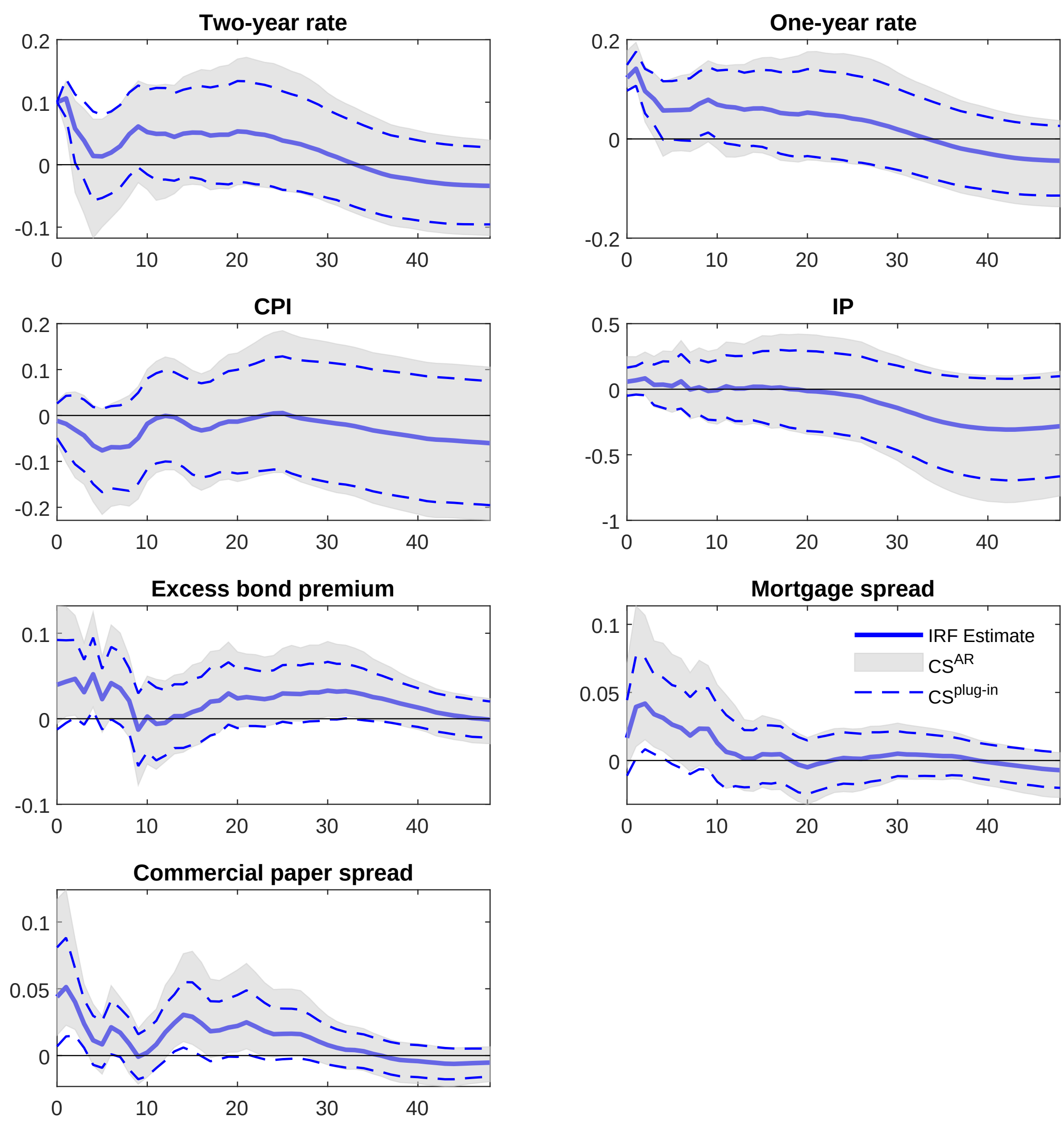

Figure 5: Impulse responses to a two-year rate shock with FF4 as instrument; Period: 1984:1-2012:6

4.3 Alternative sample period

We also extend the analysis by estimating the VAR over the period 1984:1-2020:2. Monthly data on fed funds futures surprises for this extended sample were constructed as in Gürkaynak et al. (2005) and Gertler and Karadi (2015). Swanson and Williams (2014) and Gertler and Karadi (2015) suggest that the zero lower bound was not a constraint on the Federal Reserve's ability to manipulate the two-year rate. So, we use the two-year government bond rate as the policy indicator in order to address concerns related to the zero lower bound while the policy shock is instrumented using the change in three 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 6 shows the responses of the variables to a one standard deviation monetary tightening shock. As seen in this figure, the effects of the shock on both inflation (CPI) and output (IP) are modest and statistically insignificant, thus confirming our previous analyses. The results remain essentially the same when the one-year government bond rate is used as the policy indicator or when the surprises in the current month's fed funds futures (FF1) are used as instrument. The results are also similar to those reported here when the Great Recession and the subsequent zero lower bound periods are excluded, i.e. when the subsample 1984:1-2008:6 is used in the estimation.

5 Conclusion

This paper revisits the macroeconomic effects of monetary policy shocks using an external instrument identification approach (Stock and Watson, 2012) and econometric methods that are robust to the weak instrument problem (Olea et al., 2020). Using the framework of Gertler and Karadi (2015), we analyze the joint response of financial and real economic variables to monetary policy shock, upon incorporating innovations in forward guidance into the measure of the shock. Our results suggest that the real effects of monetary policy shock on economic activity is sensitive to the choice of the sample period. In particular, leaving out Volcker

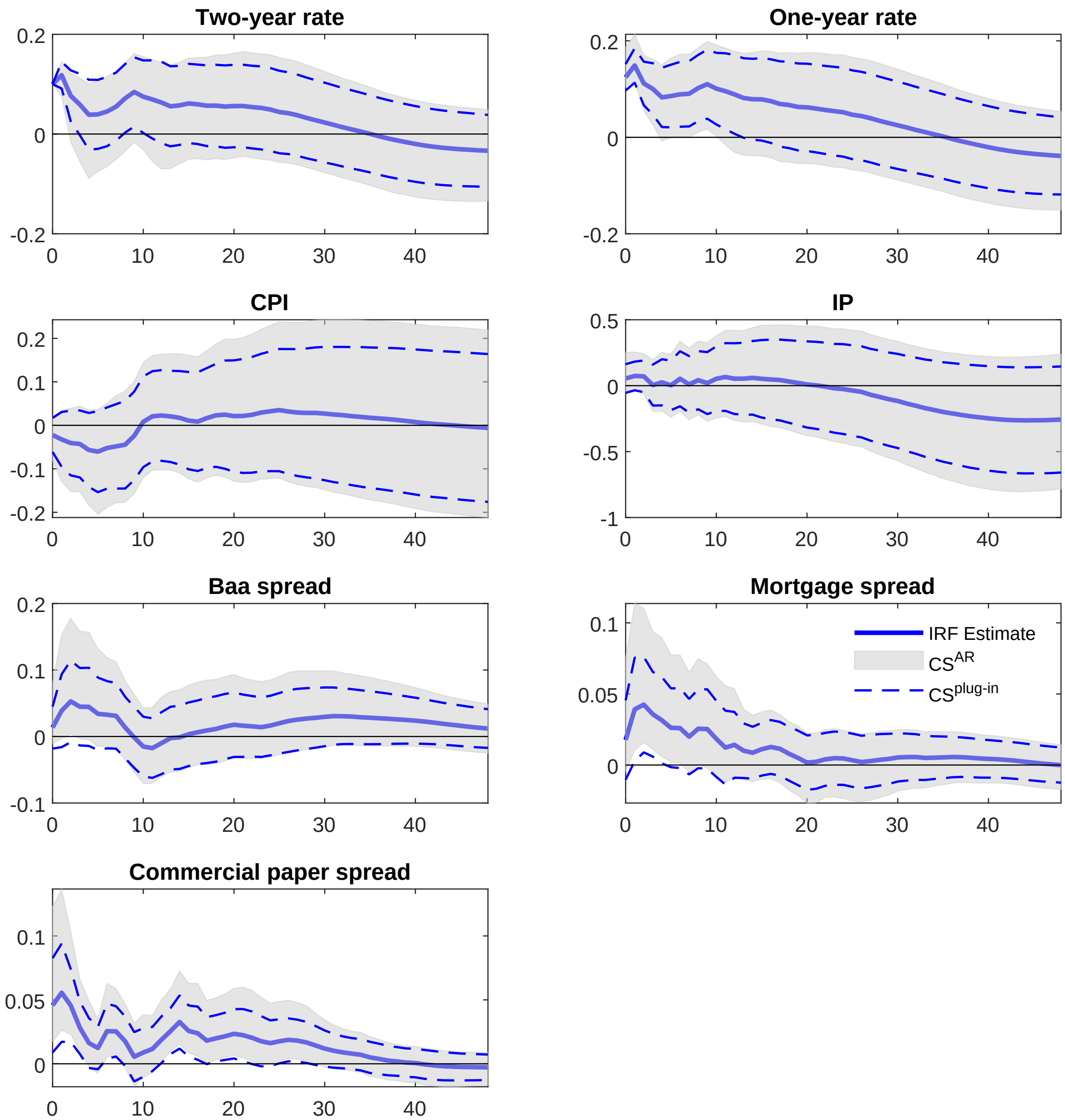

Figure 6: Impulse responses to a two-year rate shock with FF4 as instrument, Period: 1984:1-2020:2

disinflation from the estimated sample and focusing on the post-1984 period, we find that estimated monetary policy shocks have no significant effects on real economic activity and inflation. This holds despite large movements in credit costs arising due to the policy shock, a finding that deviates from Gertler and Karadi (2015) who suggest that large movements in credit costs may explain their finding of a substantial impact of monetary policy shocks on economic activity. One important implication of our finding is that widely considered stylized fact of a hump-shaped response of output to monetary policy shock provides poor guidance for implementing monetary policy and evaluating the efficacy of structural models.

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Appendix

A Futures rates surprises

We extend Gertler and Karadi's (2015) data set on futures rates surprises.[16] First, we reconstruct a list of dates and times of monetary policy announcements between 2012 and February 2020, corresponding to the press release times after every FOMC meeting. Table 1 shows the dates and times of each FOMC announcement based on an analysis of time-stamps of Dow Jones newswires (2012:1-2015:12) and FOMC press releases from their meeting calendars (2016:1-2020:2).[17] Second, high frequency data on federal funds futures contracts (at 15-minute intervals) are obtained from the *Refinitiv Tick* History database. For each announcement, we measure the surprise component of the current month's fed funds futures (FF1) and the three month ahead monthly fed funds futures (FF4) as in Gertler and Karadi (2015).[18] In particular, letting f_{t+j} denote 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 prior to the FOMC meeting, we can express the surprise in the futures rate as:[19]

$$surprise = f_{t+j} - f_{t+j,-1} \tag{A.1}$$

Following Gürkaynak et al. (2005) and Gertler and Karadi (2015), we measure the surprises within a "tight" window of the announcement in order to ensure that these surprises reflect only news about the FOMC decision. In particular, we consider the last available price

[16] The original sample in Gertler and Karadi (2015) covers the period 1990:1-2012:6.

[17] See Gürkaynak et al. (2005) and Lucca and Moench (2015) for dates and times for FOMC announcements between 1990 and 2011.

[18] Other measures on the surprise component suggested by Gürkaynak et al. (2005) are based on the six month, nine month and one year ahead futures surprises in three month Eurodollar deposits (henceforth ED2, ED3 and ED4, respectively). We focus here on FF1 and FF4 given the earlier evidence that these two external instruments are strong.

[19] 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. For surprises in the last 7 days of the month we consider the first-difference in the closest to maturity contract, for example for a meeting in the last 7 days of a month the surprise in FF1 is the first-difference of FF2.

between 5-20 minutes before the announcement and the first available price between 15-25 minutes after the announcement.[20] Finally, to convert futures surprises on FOMC days into monthly average surprises we proceed as in Gertler and Karadi (2015). Specifically, for each day we first cumulate the surprises on FOMC days during the last 31 days. Then, the monthly surprise for a given month is computed as the average of the cumulative surprises over the month.

Table 1: FOMC meeting dates and times 2012 - 2020

year	1st	2nd	3rd	4th	5th	6th	7th	8th	9th
2012	2012-01-25 12:36	2012-03-13 14:16	2012-04-25 12:32	2012-06-20 12:32	2012-08-01 14:13	2012-09-13 12:31	2012-10-24 14:30	2012-12-12 12:35	
2013	2013-01-30 14:15	2013-03-20 14:01	2013-05-01 14:00	2013-06-19 14:00	2013-07-31 14:15	2013-09-18 14:00	2013-10-30 13:59	2013-12-18 14:00	
2014	2014-01-29 14:00	2014-03-19 14:02	2014-04-30 14:00	2014-06-18 14:00	2014-07-30 14:00	2014-09-17 14:00	2014-10-29 14:00	2014-12-17 14:00	
2015	2015-01-28 14:00	2015-03-18 14:00	2015-04-29 14:00	2015-06-17 14:00	2015-07-29 14:00	2015-09-17 14:00	2015-10-28 14:00	2015-12-16 14:00	
2016	2016-01-27 14:00	2016-03-16 14:00	2016-04-27 14:00	2016-06-15 14:00	2016-07-27 14:00	2016-09-21 14:00	2016-11-02 14:00	2016-12-14 14:00	
2017	2017-02-01 14:00	2017-03-15 14:00	2017-05-03 14:00	2017-06-14 14:00	2017-07-26 14:00	2017-09-20 14:00	2017-11-01 14:00	2017-12-13 14:00	
2018	2018-01-31 14:00	2018-03-21 14:00	2018-05-02 14:00	2018-06-13 14:00	2018-08-01 14:00	2018-09-26 14:00	2018-11-08 14:00	2018-12-19 14:00	
2019	2019-01-30 14:00	2019-03-20 14:00	2019-05-01 14:00	2019-06-19 14:00	2019-07-31 14:00	2019-09-18 14:00	2019-10-11 11:00	2019-10-30 14:00	2019-12-11 14:00
2020	2020-01-29 14:00								

[20] In times when no federal funds futures were traded exactly at the beginning of the specified window, we use the most recent price. Similarly, when there were no trades exactly at the end of the specified window, we use the next available trade price.

B Additional results

Figure 7 shows the impulse responses for the baseline VAR with 24 lags (as opposed to 12 lags in our main analysis). We have also used 6 lags and 18 lags and the results remain essentially the same as those reported here and so are omitted to shorten the presentation. As seen in the figure, the impulse responses are similar to the baseline VAR with 12 lags, thereby corroborating our main findings.

Figures 8 to 10 show the impulse responses for the baseline VAR using FF1, ED2 and ED3, one at a time, as the external instrument. While the confidence sets (both with the plug-in and identification-robust methods) for the impulse responses with FF1 and ED2 as instruments are similar to those with FF4, the identification-robust confidence sets with ED3 are very wide, indicating that ED3 is not a strong instrument (see Dufour, 1997).

Figures 11 to 14 show the impulse responses for the seven-variate VAR with the two-year government bond rate as the policy indicator and FF1, ED2, ED3 and ED4 as the instrument (one at a time). As seen in these figures, the impulse responses and their associated confidence sets with FF1, ED2 and ED3 as external instruments are similar to those in Figures 8-10. However, Figure 14 show that the identification-robust confidence sets (CS^{AR}) with ED4 as the instrument are completely unbounded, suggesting that this external instrument (i.e., ED4) is completely uninformative. This is depicted in Figure 14 as the shaded area spanning the entire space in the plot (in fact the confidence sets span the entire real line). The figure also indicates clearly that when instruments are weak the confidence sets based on the plug-in method (seen as dotted blue lines in the figure) are invalid in the sense that their true coverage probability (level) is zero (see Dufour, 1997).

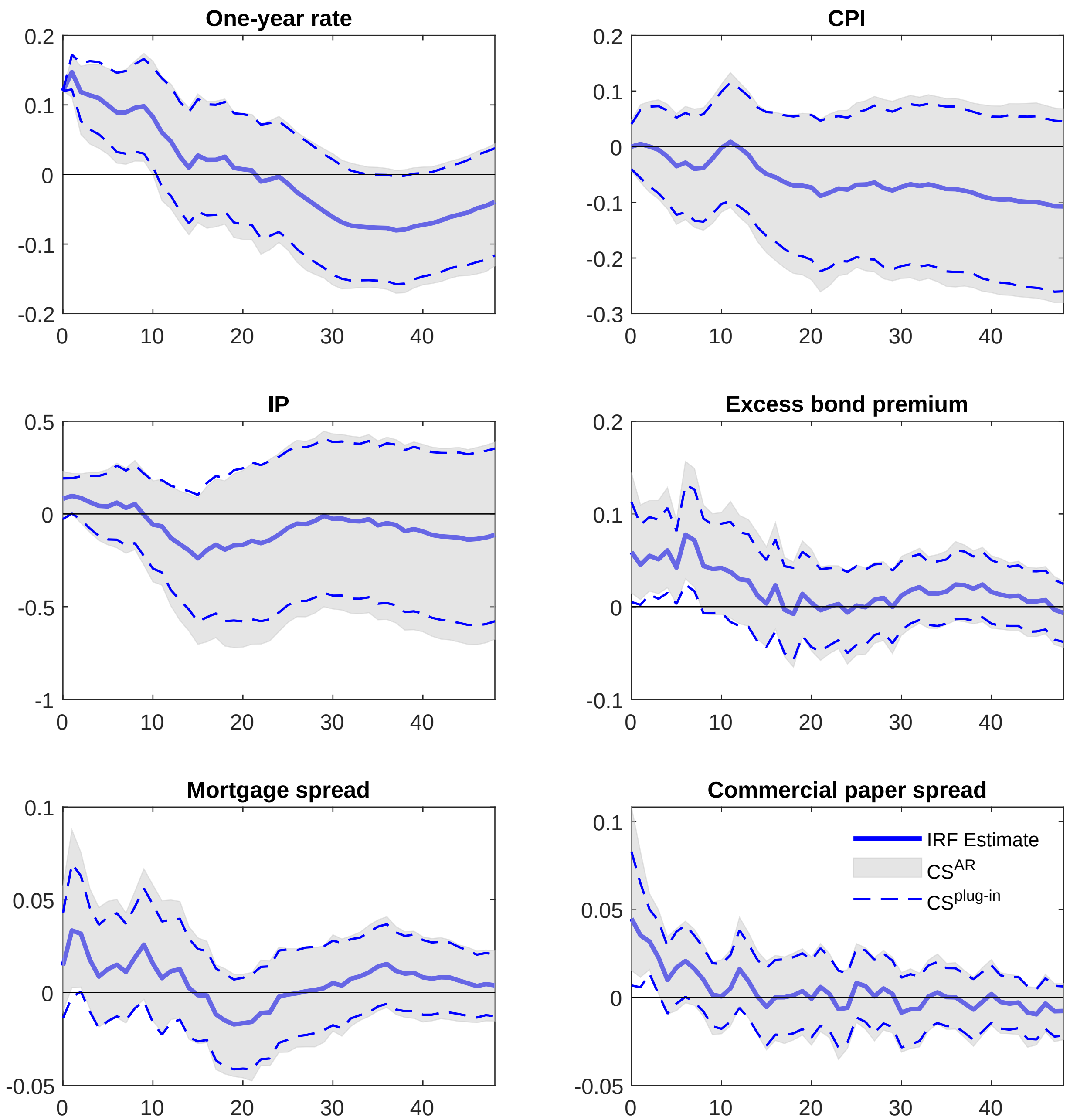

Figure 7: Impulse responses to a one-year rate shock in the baseline VAR with 24 lags; Period: 1984:1-2012:6

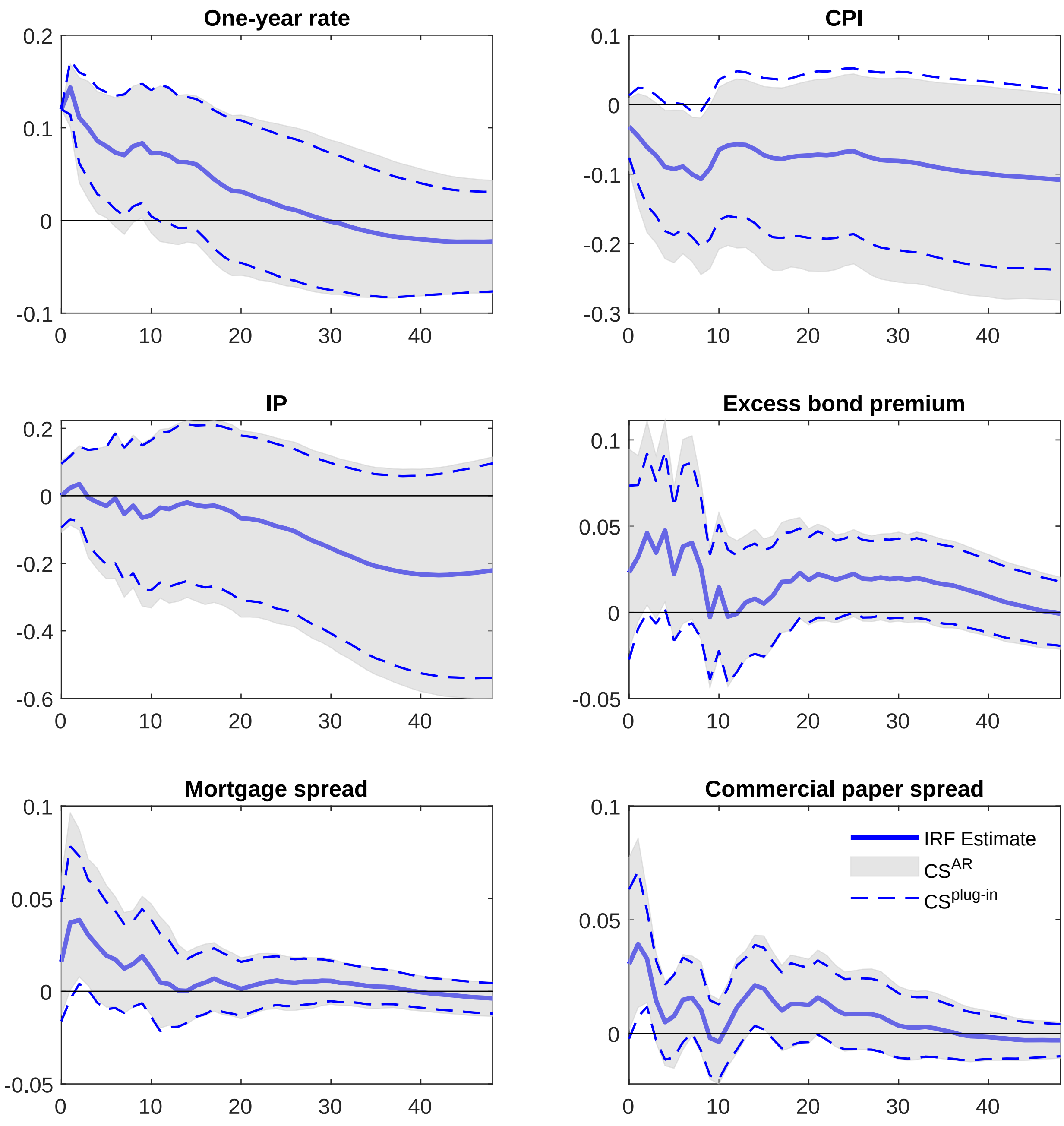

Figure 8: Impulse responses to a one-year rate shock with FF1 as instrument; Period: 1984:1-2012:6

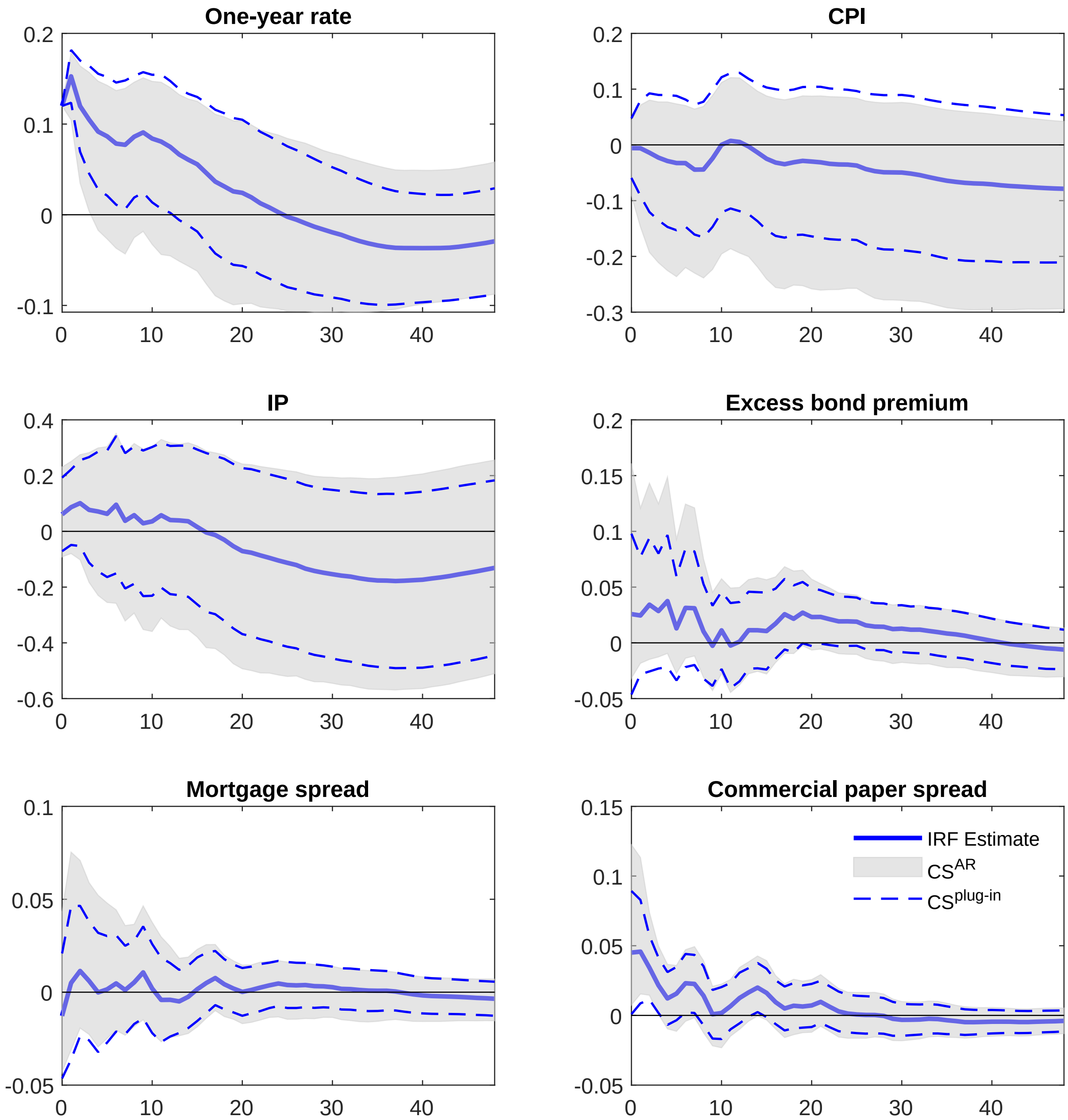

Figure 9: Impulse responses to a one-year rate shock with ED2 as instrument; Period: 1984:1-2012:6

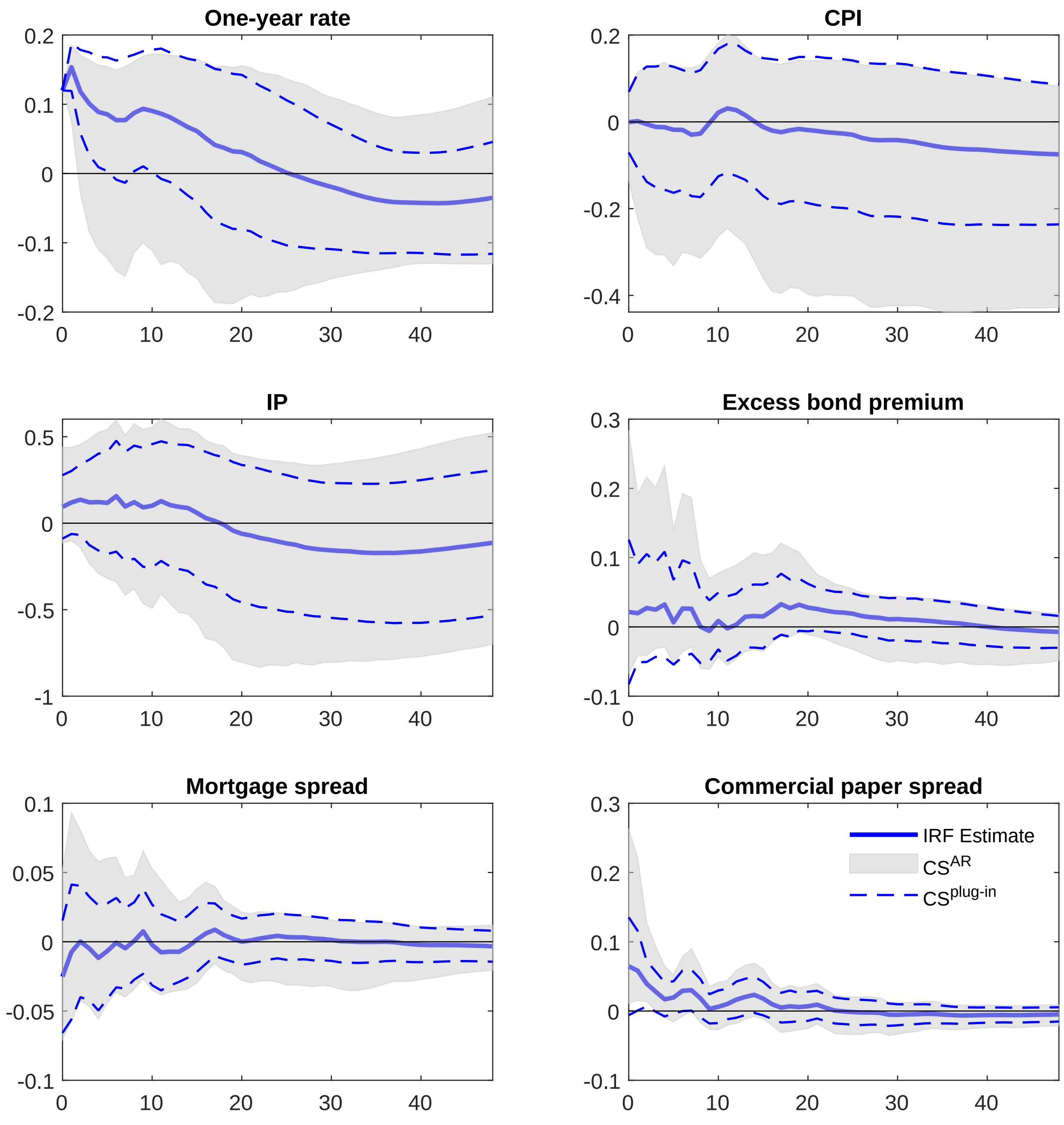

Figure 10: Impulse responses to a one-year rate shock with ED3 as instrument; Period: 1984:1-2012:6

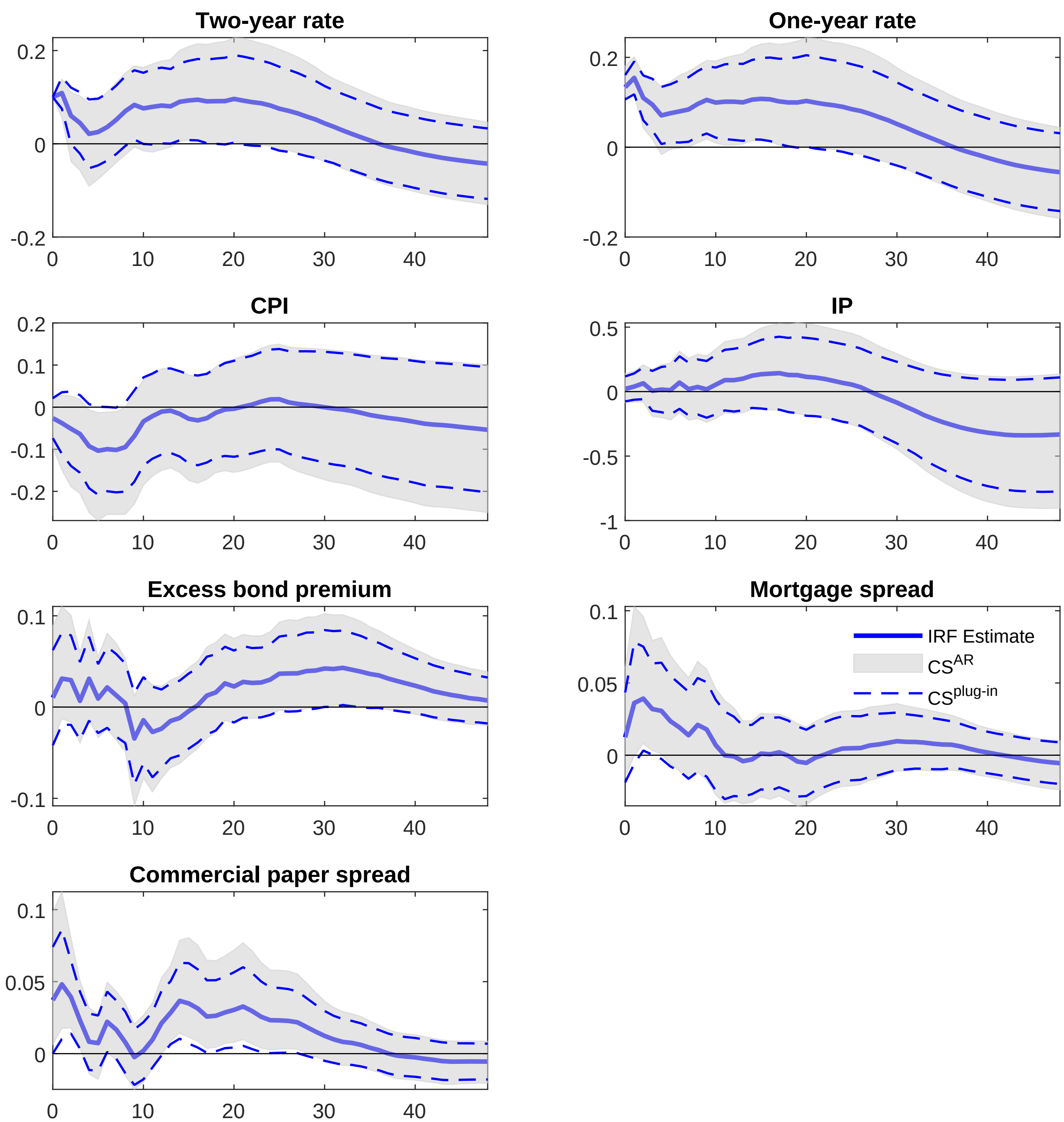

Figure 11: Impulse responses to a two-year rate shock with FF1 as instrument; Period: 1984:1-2012:6

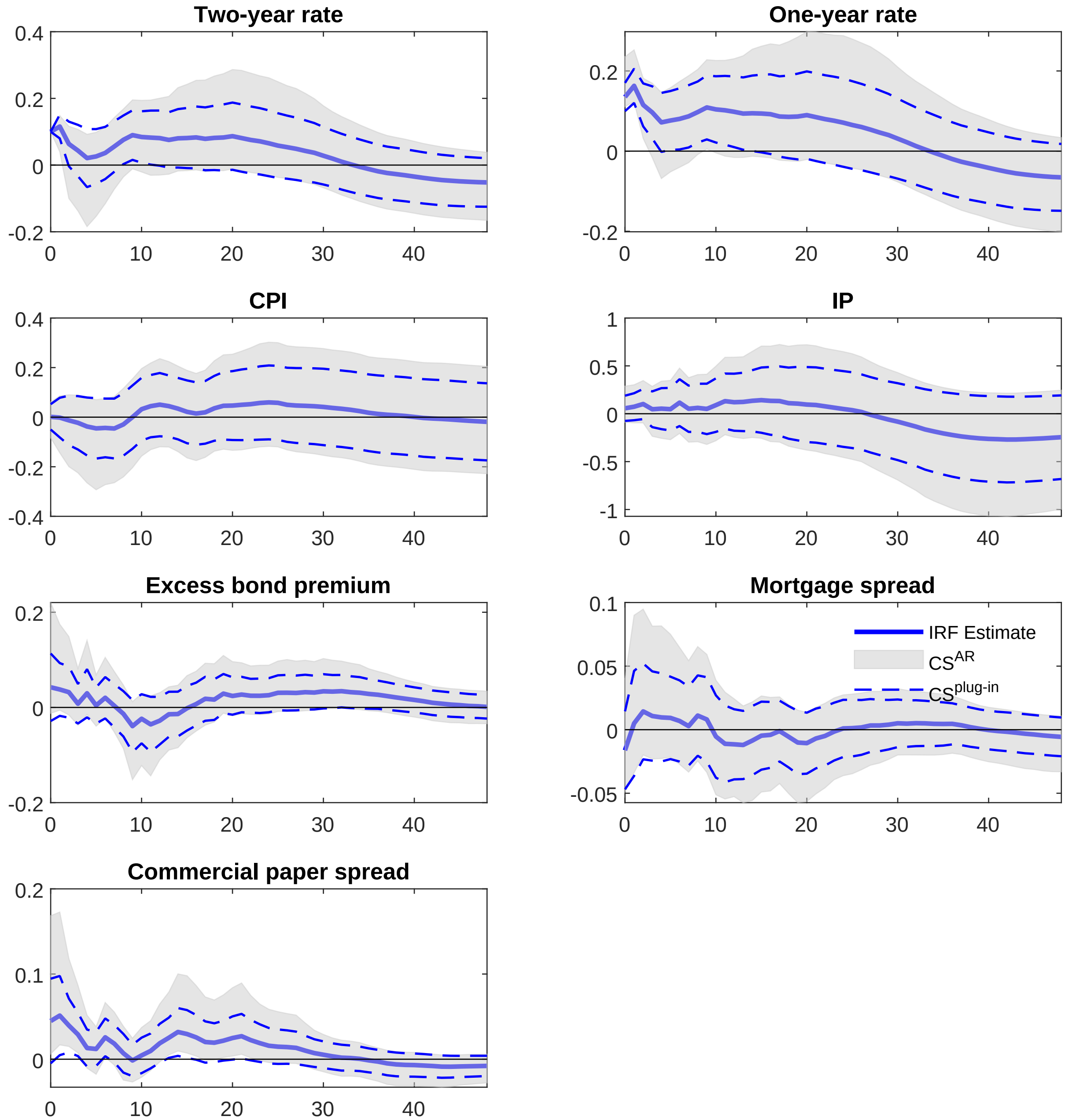

Figure 12: Impulse responses to a two-year rate shock with ED2 as instrument; Period: 1984:1-2012:6

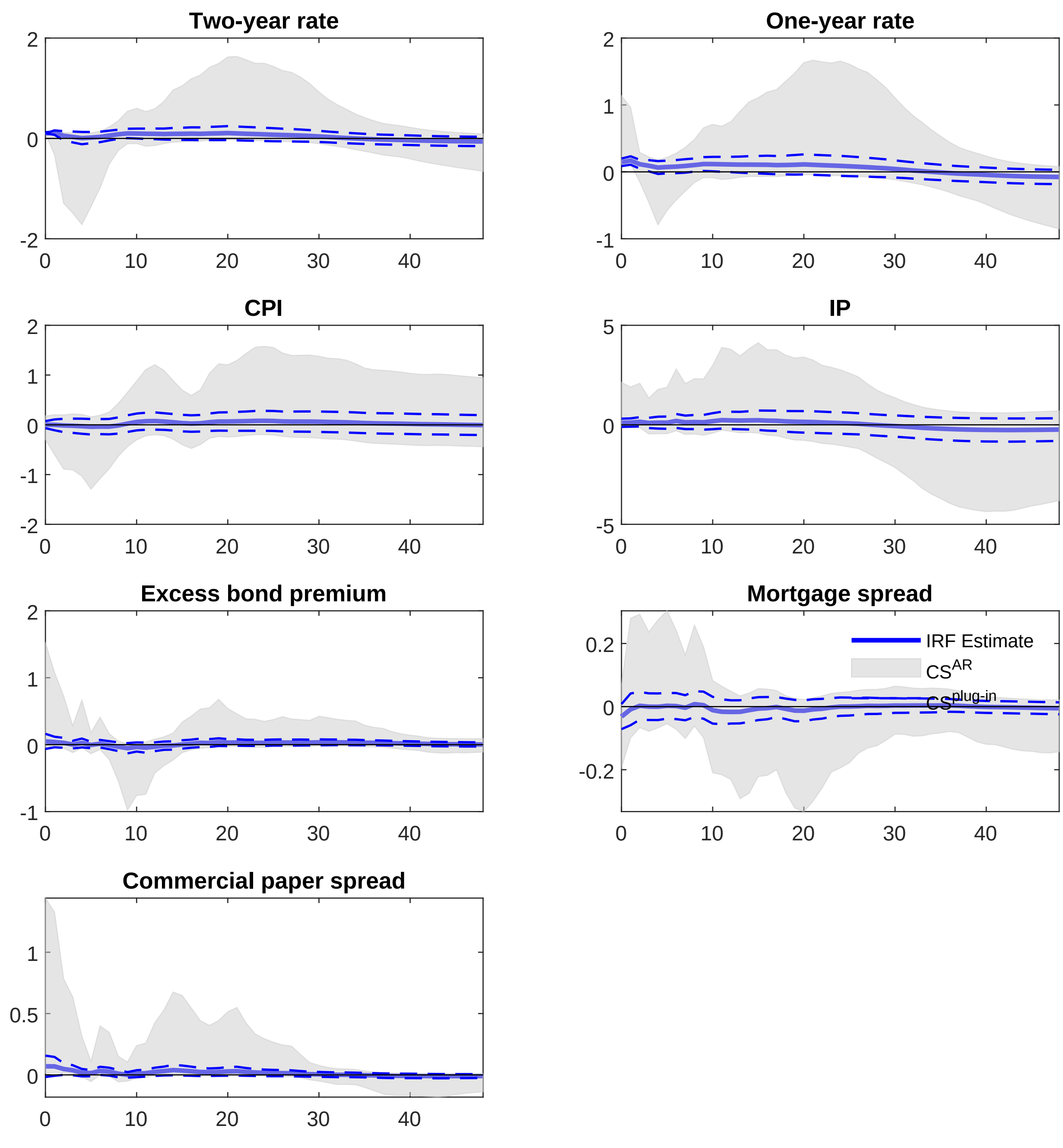

Figure 13: Impulse responses to a two-year rate shock with ED3 as instrument; Period: 1984:1-2012:6

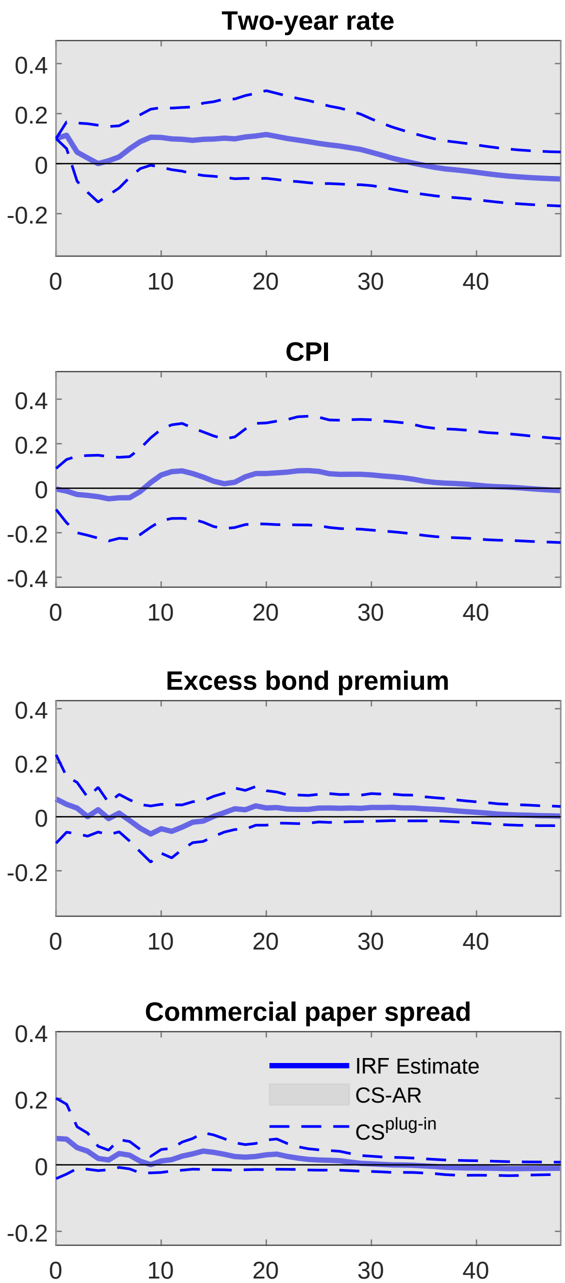

Figure 14: Impulse responses to a two-year rate shock with ED4 as instrument; Period: 1984:1-2012:6