Monetary Policy Counterfactuals: Time Series Evidence on the Fiscal Multiplier*

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Abstract

How does systematic monetary policy shape the effects of macroeconomic shocks? We propose a general time series approach to construct monetary policy counterfactuals, which are robust to the Lucas critique. To operationalize this approach, we leverage historical variation in systematic US monetary policy. Identification is achieved through a novel FOMC rotation instrument. We apply our approach to the question of fiscal-monetary interactions and show that systematic monetary policy affects strongly and significantly the response of the economy to fiscal policy. In a counterfactual, in which monetary policy does not respond to fiscal spending shocks, we find a large fiscal spending multiplier of size 2 at a three-year horizon.

Keywords: policy counterfactual, monetary policy, fiscal multiplier
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1 Introduction

In this paper, we study monetary policy counterfactuals. These counterfactuals allow us to address questions as to how the economy evolves if monetary policy responds more or less aggressively to macroeconomic developments, such as rising inflation or a fiscal expansion. This is important because "what-if" questions of this type are key inputs into monetary policy deliberations. Such counterfactuals are also important to understand how monetary policy shapes the effects of macroeconomic shocks.

We make three contributions to the literature. First, we introduce a general empirical framework and show how monetary policy counterfactuals, that are robust to the Lucas critique, can be identified using historical variation in systematic monetary policy. Second, we operationalize our approach using Istrefi's (2019) narrative account, which classifies policy preferences of Federal Open Market Committee (FOMC) members on a hawk-dove scale, as a measure of perceived systematic monetary policy. To achieve identification, we construct an instrumental variable based on the rotation of voting rights of Federal Reserve Bank (FRB) presidents within the FOMC. Third, we apply our approach to the important and perennial question of fiscal-monetary interactions and show that systematic monetary policy strongly and significantly affects the response of the economy to fiscal shocks. We document that the cumulative fiscal spending multiplier after three years increases from about zero to 2 in a counterfactual in which the FOMC does not respond to the fiscal spending shock.

Policy counterfactuals from time series models are commonly subject to the Lucas (1976) critique because reduced form coefficients generally change when perceived systematic policy changes. We circumvent this challenge by explicitly accounting for historical variation in perceived systematic monetary policy. We capture the dependence of reduced-form relationships on systematic policy using non-linear local projections (LP). The LPs deliver counterfactual impulse response functions (IRF) for different policy configurations. Implementation of our approach requires a measurement of time variation in systematic monetary policy as well as exogenous variation for identification.

We measure perceived systematic monetary policy using Istrefi's (2019) narrative account of FOMC members' policy preferences. This measure is based on the US newspaper coverage of the Federal Reserve and reflects the FOMC composition in terms of hawks and doves, as perceived in real time by Fed watchers for the period 1960 to 2014. FOMC hawks are those members perceived to be more concerned about inflation than unemployment. Conversely, doves are more worried about economic growth and employment than inflation. We propose a novel FOMC rotation instrument to address the endogeneity concern that the variation in the FOMC's hawk and dove composition may relate to the state of the economy. This instrument exploits the annual mechanical rotation scheme that allocates voting rights to four out of eleven FRB presidents and that has been in place since 1943.[1]

We first use our approach to estimate how systematic monetary policy, taking into account the variation in the policy leanings of FOMC members, shapes the federal funds rate (FFR) response to the Federal Reserve Greenbook inflation projections. We show that a hawkish FOMC increases the FFR more aggressively and persistently in response to high inflation forecasts. This supports the validity of our measure of systematic monetary policy.

Second, we apply our approach to an important and perennial question in the literature: How does monetary policy affect the response of the economy to fiscal spending shocks in the US? To address this question, we use identified fiscal spending shocks from Ramey and Zubairy (2018) and estimate non-linear local projections, which account for time variation in systematic monetary policy. Our estimates characterize the importance of systematic monetary policy for the effectiveness of fiscal policies and allow us to construct monetary policy counterfactuals. We consider a counterfactual scenario in which monetary policy is non-responsive to fiscal spending shocks in the sense that the FFR remains unchanged. Our preferred specification suggests a cumulative fiscal spending multiplier of 2 at a three-year horizon. This counterfactual requires only one hawk less or one dove more, compared to the

[1] The FOMC consists of twelve voting members - the seven members of the Board of Governors of the Federal Reserve System, the president of the Federal Reserve Bank of New York and four of the remaining eleven Reserve Bank presidents, who serve one-year terms on a rotating basis.

average composition. While this is a small change compared to the time variation of the FOMC preferences, its effects are sizeable when compared to the average fiscal multiplier, which is close to zero.

Overall, our results highlight a source of state-dependence which is qualitatively and quantitatively important. In particular, the estimates suggest that systematic monetary policy plays a crucial role for successful fiscal stimulus, consistent with theory (e.g. Woodford, 2011) and recent empirical evidence (e.g. Canova and Pappa, 2011; Cloyne et al., 2021).

This paper relates to two strands of literature. First, it contributes to a broad literature that estimates policy counterfactuals using either reduced-form, structural, or semi-structural approaches. The reduced-form approach uses linear VAR (or LP) estimates to implement a counterfactual policy rule through counterfactual monetary policy shocks (Sims, 1980; Bernanke et al., 1997; Leeper and Zha, 2003; Eberly et al., 2019; Antolin-Diaz et al., 2021; Benati, 2021). This approach is valid for modest changes in systematic policy that are perceived as shocks by private agents (Leeper and Zha, 2003). For non-modest interventions, the Lucas critique applies which is a common problem in many applications (e.g. Benati, 2010; Kilian and Lewis, 2011). In comparison, our approach remains valid as long as *systematic* policy does not deviate much from private agents' expectations. Another reduced-form approach is to estimate non-linear VAR models in which systematic monetary policy exogenously varies over time (e.g. Sims et al., 1982; Primiceri, 2005; Sims and Zha, 2006). In comparison, our approach uses a direct measurement of systematic monetary policy and explicitly addresses the endogeneity issues. Finally, the approach in Cloyne et al. (2021) is similar to ours but it relies on cross-country differences in systematic monetary policy, whereas we focus on US time series variation.

The structural approach of constructing policy counterfactuals requires a fully-specified structural model and admits counterfactuals by varying model parameters directly (e.g. Lucas, 1980; Christiano et al., 1999; Woodford, 2003). More recent semi-structural approaches show that valid counterfactuals can be obtained from weaker assumptions. McKay and Wolf

(2021) show that news shocks can be used to construct counterfactuals which are valid in a broad class of linearized DSGE models. In a similar vein, Beraja (2020) finds that it can be sufficient to impose only some cross-equation restrictions to obtain valid counterfactuals. In contrast, our approach does not rely on such structural assumptions and has different informational requirements which makes it complementary to the preceding literature.

Second, we relate to the vast literature on fiscal multipliers. The estimates of the average fiscal spending multiplier in the majority of papers tend to be small (Blanchard and Perotti, 2002; Mountford and Uhlig, 2009; Ramey, 2011). More recent papers ask whether fiscal multipliers are large only under particular circumstances, e.g. during recessions (Auerbach and Gorodnichenko, 2012, 2013; Bachmann and Sims, 2012), or at the zero lower bound (Ramey and Zubairy, 2018), or whether the configuration of systematic fiscal policy (Caldara and Kamps, 2017) or the sign of the shock (Barnichon et al., forthcoming) matters. The configuration of systematic monetary policy is another source of state-dependence found to be important using sign restriction (Canova and Pappa, 2011), Bayesian estimation of structural models (Leeper et al., 2017) or cross-country differences in monetary policy (Cloyne et al., 2021). In comparison, to the best of our knowledge, we are the first to leverage historical variation in measured systematic US monetary policy to study fiscal multipliers.

The paper is organized as follows: Section 2 describes how monetary policy counterfactuals can be identified when variation in systematic monetary policy is observed. Section 3 proposes how to measure and instrument systematic US monetary policy. Section 4 applies our approach to study the response of monetary policy to inflation. Section 5 studies monetary policy counterfactuals for the effects of fiscal policy shocks. Section 6 concludes.

2 Identification of Policy Counterfactuals

Constructing reduced-form monetary policy counterfactuals demands to identify the effects of systematic monetary policy and its interaction with other macroeconomic shocks (e.g. a fiscal shock). This is a challenging endeavor because systematic monetary policy is likely to co-vary with other macroeconomic variables. Further, systematic monetary policy may respond to changes in other components of systematic policy (e.g. fiscal policies) and vice versa. Both arguments render fluctuations in systematic monetary policy endogenous.

This section addresses these threats to identification. We introduce a general empirical framework which naturally lends itself to study dynamic responses with interacted local projection (LP) models. We derive two results which give sufficient conditions for identification of systematic monetary policy. First, we show that ordinary least squares (OLS) achieves identification only under a particularly strong exogeneity assumption. Second, we show that an instrumental variable (IV) estimator achieves identification under standard relevance and exclusion restrictions. Importantly, we show that an instrument that captures exogenous variation in systematic monetary policy satisfies these conditions under relatively mild assumptions on reduced-form parameters. In particular, the IV approach allows for endogenous variation in systematic monetary policy as well as correlation with other policies (e.g. fiscal policy). We further explain how interacted LP models can be used to construct policy counterfactuals and discuss under which conditions the Lucas (1976) critique applies.

2.1 General Framework

2.1.1 Setting

Consider some time series $y_t \in \mathbb{R}$ with a general data-generating process (DGP)

$$y_t = f\left(\{\varepsilon_{t-j}\}_{j=0}^{\infty}, \{x_{t-j}\}_{j=0}^{\infty}\right), \tag{2.1}$$

in which we assume that ε_{t-j}, x_{t-j} and y_t follow a joint stationary process, which is ergodic up to fourth (finite) moments. ε_t is a $N_\varepsilon \times 1$ vector of exogenous *impulses* (i.e., structural shocks) and x_t is a $N_x \times 1$ vector of time-varying coefficients, associated with the *propagation* of impulses, shortly propagation variables. In general, x_t may include the state of the business cycle, financial conditions, or the stance of systematic fiscal and monetary policy. Our proposal is to lever variation in x_t associated with systematic monetary policy to construct monetary policy counterfactuals. The DGP in (2.1) is a fairly general setting which nests a broad class of DSGE models. We next impose some structure on x_t.

Assumption 1. x_t follows the vector auto-regressive process of order 1

$$x_t = Ax_{t-1} + B\varepsilon_t + C\eta_t, \tag{2.2}$$

where A is a $N_x \times N_x$ matrix with all eigenvalues inside the unit circle, B and C are matrices of dimension $N_x \times N_\varepsilon$ and $N_x \times N_\eta$, respectively, and η_t is a $N_\eta \times 1$ vector of innovations.

The innovations η_t affect the outcome y_t only indirectly through their effect on x_t whereas the structural impulses ε_t affect the outcome y_t directly through the first argument of f and indirectly through x_t.[2] Importantly, this setup allows the propagation variables x_t to co-move with each other and with the impulses ε_t. For example, systematic monetary policy, say x_{1t}, may co-move with systematic fiscal policy, say x_{2t}, and may respond to fiscal policy shocks, say ε_{1t}. Random variables ε_t and η_t further satisfy the following assumption.

Assumption 2. ε_{it} and η_{js} are mutually independent, identically distributed over time with finite second moments and $\mathbb{E}[\varepsilon_{it}] = \mathbb{E}[\varepsilon_{it}^3] = \mathbb{E}[\eta_{js}] = 0$ for all i, j, t, s.

Next, we consider Taylor approximations of f and introduce LP estimands of interest.

2.1.2 Standard LP

Consider a first-order Taylor approximation of (2.1) with respect to all arguments of f around $\varepsilon_{t-j} = \mathbb{E}\left[\varepsilon_{t-j}\right] = \bar{\varepsilon} = 0$ and $x_{t-j} = \mathbb{E}\left[x_{t-j}\right] = \bar{x} = 0$ for $j \geq 0$.

$$y_t = \sum_{j=0}^{\infty} \left(\bar{f}_{\varepsilon_j} \varepsilon_{t-j} + \bar{f}_{x_j} x_{t-j} \right) \tag{2.3}$$

[2] We assume a first-order Markov process in (2.2) to keep the exposition parsimonious. In principle, we could allow for higher order auto-regressive processes without fundamentally altering our main results below.

We denote the transposed gradients evaluated at the point of approximation by $\bar{f}_{\varepsilon_j} = \left(\nabla_{\varepsilon_{t-j}} f(\cdot,\cdot)|_{\bar{\varepsilon},\bar{x}}\right)'$ and $\bar{f}_{x_j} = \left(\nabla_{x_{t-j}} f(\cdot,\cdot)|_{\bar{\varepsilon},\bar{x}}\right)'$ respectively and impose without loss of generality that $f(\cdot,\cdot)|_{\bar{\varepsilon},\bar{x}} = 0$. Note that $\bar{f}_{\varepsilon_j}$ and $\bar{f}_{x_j}$ depend only on the time lag j and not on t because the function f is time-invariant. Consider the partition $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t}')'$ where ε_{1t} denotes a (scalar) shock of interest ε_{2t} is a $N_\varepsilon - 1 \times 1$ vector containing all other exogenous impulses. Now consider a standard LP to estimate the dynamic effects of ε_{1t} on y_{t+h}:

$$y_{t+h} = \beta^h \varepsilon_{1t} + v_{t+h}^h, \qquad h \geq 0. \tag{2.4}$$

The subsequent proposition characterizes the estimand β^h in terms of the coefficients of the DGP in (2.3) and establishes consistency for the ordinary least-square (OLS) estimator.

Proposition 1. *Assume the data is generated by (2.3) and Assumptions 1 and 2 hold. Then, the OLS estimator* $\hat{\beta}^h$ *of* β^h *in* (2.4) *satisfies:*

$$\hat{\beta}^h \xrightarrow{p} \beta^h \qquad \textit{with} \qquad \beta^h = \left(\bar{f}_{\varepsilon_h}\right)_1 + (\sum_{j=0}^{h} \bar{f}_{x_j} A^{h-j} B)_1$$

Proof. See Appendix A.

The notation $(\cdot)_i$ means that the i-th element of the vector is selected. Note that β^h contains the direct and indirect effects of ε_{1t}. The first term captures the direct effect on the outcome y_{t+h}, while the second term captures the indirect effects on the outcome via changing the propagation variables $\{x_{t+i}\}_{i=0}^h$. This indirect effect enters through (2.2). An econometrician may isolate the direct effects by controlling for x_t. Yet, as long as the total effect is the object of interest, e.g., the GDP response to a fiscal spending shock, one explicitly wants to include the indirect effects. This changes when one is interested in how x_t *shapes* the overall response of y_{t+h}. We study this next by means of an interacted LP.

2.1.3 Interacted LP

We now add second-order cross-effects between ε_{t-j} and x_{t-i} for $j \geq 0$ and $i \geq 0$ to the first-order Taylor approximation in (2.3).

$$y_t = \sum_{j=0}^{\infty} \left(\bar{f}_{\varepsilon_j} \varepsilon_{t-j} + \bar{f}_{x_j} x_{t-j} \right) + \sum_{j=0}^{\infty} \sum_{i=0}^{\infty} \left(\varepsilon'_{t-j} \bar{f}_{\varepsilon_j, x_i} x_{t-i} \right) \tag{2.5}$$

$\bar{f}_{\varepsilon_j,x_i}$ denotes a sub-matrix of the Hessian which capture the cross-effect between ε_{t-j} and x_{t-i}, evaluated at the point of approximation analogous to the gradients. Consider the partition $x_t = (x_{1t,} x'_{2t})'$ where x_{1t} denotes a (scalar) propagation variable of interest and x_{2t} is a $N_x - 1 \times 1$ vector that contains the remaining propagation variables.

Now consider an interacted LP, which adds to the standard LP in (2.4), x_{1t} and the interaction between ε_{1t} and x_{1t}:

$$y_{t+h} = \beta^h \varepsilon_{1t} + \gamma^h (\varepsilon_{1t} \times x_{1t}) + \delta^h x_{1t} + v_{t+h}^h, \qquad h \geq 0. \tag{2.6}$$

This interacted LP allows us to estimate the effect of ε_{1t} on y_{t+h} and its dependence on x_{1t}. Identifying the parameters of equation (2.6) is a challenging problem to which we turn next.

2.2 Identification

2.2.1 Identification Problem

Suppose an econometrician estimates equation (2.6) by OLS. The OLS estimators is only consistent when $\mathbb{E}\left[X_t v_{t+h}^h\right] = 0$, where $X_t = (\varepsilon_{1t}, x_{1t}, \varepsilon_{1t} \times x_{1t})'$. However, this is not generally satisfied because x_{1t} is endogenous. For example, x_{1t} may correlate with ε_{2t} and x_{2t}, which are contained in the residual v_{t+h}^h. Eliminating the endogeneity of x_{1t} requires additional assumptions.

Assumption 3. $B_{1,\cdot} = 0, \quad A_{1,2:N_x} = 0, \quad A_{\cdot,1} = 0, \quad C_{1,2:N_\eta} = 0, \quad C_{2:N_\eta,1} = 0$

This assumption is sufficient to ensure that x_{1t} is *not* serially correlated and independent of ε_t and η_{2t}, where η_{2t} is a $N_\eta - 1 \times 1$ vector based on the partition $\eta_t = (\eta_{1t}, \eta_{2t}')'$. Next, we state our first identification result in Proposition 2. The proposition characterizes the estimands of the interacted LP in (2.6) and establishes consistency for the OLS estimator under Assumptions 1 to 3. The notation $(\cdot)_{2:N.}$ means that one takes elements from the vector starting with the second element and ending with the $N.$-th element.

Proposition 2. *Assume the data is generated by (2.5) and Assumptions 1–3 hold. Then, the OLS estimator $\hat{\theta}^h$ of $\theta^h = \left(\beta^h, \gamma^h, \delta^h\right)'$ in (2.6) satisfies:*

$$\hat{\theta}^h \xrightarrow{p} \theta^h \qquad \textit{with} \qquad \beta^h = \left(\bar{f}_{\varepsilon_h}\right)_1 + \left(\sum_{j=0}^{h} \left(\bar{f}_{x_j} A^{h-j}\right)_{2:N_x} B_{2:N_x,\cdot}\right)_1,$$

$$\gamma^h = \left(\sum_{i=0}^{h} \bar{f}_{\varepsilon_h, x_i} A^{h-i}\right)_{1,1} \qquad \textit{and} \qquad \delta^h = \left(\sum_{j=0}^{h} \bar{f}_{x_j} A^{h-j}\right)_1$$

Proof. See Appendix A.

The key message of Proposition 2 is that the OLS estimate of (2.6) can be consistent, albeit only under strong assumptions. Conversely, the identification problem is that the OLS estimate is inconsistent under less restrictive assumptions than Assumption 3, for example, if we allow x_{1t} to co-move with x_{2t} and to respond to shocks ε_{2t}.

The estimands in Proposition 2 deserve some discussion. Both δ^h and γ^h capture the fact that the propagation variable $\{x_{t+j}\}_{j=0}^{h}$ shapes the response of y_{t+h} according to (2.5). Intuitively, the A terms give expectations over future propagation variables, given x_{1t} and given the law of motion for x_t in equation (2.2). The $\bar{f}_{\cdot}$ terms map these conditional expectations into (expected) outcomes. As for β^h, the first and second terms reflect the direct and indirect effects of ε_{1t}, respectively. The term $B_{2:N_x,\cdot}$ maps the the vector of (contemporaneous) shocks ε_t into x_t. The first row of matrix B is excluded because the effect of x_{1t} is absorbed by δ^h. Only the first entry of the entire vector is contained in β^h because ε_{2t} is not included as regressor. It is further instructive to compare the estimand β^h with the one from the standard LP without interactions in Proposition 1. The estimand in Proposition 1 also includes the indirect effects of x_{1t}. This can be easily seen as the B matrix includes the first row.

Next, we propose an IV approach. The IV approach no longer requires Assumption 3. In particular, we can allow for x_{1t} to respond to ε_t and η_{2t} .

2.2.2 Instrumental Variable Approach

We propose an IV approach to identify the parameters in equation (2.6). Consider a variable $z_t \in \mathbb{R}$.

Assumption 4. For all t and all $h \geq 0$, it holds that

$$\mathbb{E}\left[z_t x_{1t}\right] \neq 0 \quad \text{and} \quad \mathbb{E}\left[(\varepsilon_{1t} \times z_t)(\varepsilon_{1t} \times x_{1t})\right] \neq 0, \tag{Relevance}$$

$$\mathbb{E}\left[z_t v_{t+h}^h\right] = 0 \quad \text{and} \quad \mathbb{E}\left[(\varepsilon_{1t} \times z_t) v_{t+h}^h\right] = 0, \tag{Exclusion}$$

and $z_t,\ x_t,\ \varepsilon_t$ and y_{t+h} follow a joint stationary process, which is ergodic up the fourth (finite) moments.

Assumption 4 implies that variables z_t and $\varepsilon_{1t} \times z_t$ can be used as instrumental variables for endogenous regressors x_{1t} and $\varepsilon_{1t} \times x_{1t}$. The first condition states that the instruments are correlated with the endogenous regressors (Relevance). The second condition imposes that the instruments are uncorrelated with the residual (Exclusion). For example, Assumption 4 is satisfied if z_t=η_{1t}, $C_{1,1} \neq 0$, and $C_{2:N_x,1} = 0$. The former condition on C implies relevance and the latter condition on C implies exclusion, and can be seen as a timing restriction. It is apparent that these are much weaker conditions than those imposed by Assumption 3. Restrictions on C become weaker, and A and B are not restricted.

The next proposition gives our main identification result which states that the IV estimator yields consistent estimates of the estimands from the interacted LP (2.6).

Proposition 3. *Assume the data is generated by (2.5) and Assumptions 1,2 and 4 hold. Then, the IV estimator* $\hat{\theta}^h_{IV}$ *of* $\theta^h = \left(\beta^h, \gamma^h, \delta^h\right)'$ *in (2.6) satisfies:*

$$\hat{\theta}^h_{IV} \xrightarrow{p} \theta^h \qquad \textit{with} \qquad \beta^h = \left(\bar{f}_{\varepsilon_h}\right)_1 + \left(\sum_{j=0}^{h} \left(\bar{f}_{x_j} A^{h-j}\right)_{2:N_x} B_{2:N_x,\cdot}\right)_1,$$

$$\gamma^h = \left(\sum_{i=0}^{h} \bar{f}_{\varepsilon_h, x_i} A^{h-i}\right)_{1,1} \qquad \textit{and} \qquad \delta^h = \left(\sum_{j=0}^{h} \bar{f}_{x_j} A^{h-j}\right)_1$$

Proof. See Appendix A.

The estimands are identical to those in Proposition 2. Identification can be achieved through Assumption 3 (OLS) or Assumption 4 (IV). We next explain how to construct counterfactuals.

2.3 Policy Counterfactuals

2.3.1 Constructing Policy Counterfactuals

The interacted LP discussed above can be used to construct policy counterfactuals. As we argue below, these policy counterfactuals are robust to the Lucas critique. We first define a conditional LP impulse response function (LP-IRF) in the spirit of Jordà (2005) by the difference of two forecasts.

Definition. A conditional LP-IRF is given by

$$IRF^h_{y\varepsilon}(\bar{\varepsilon}_1, \bar{x}_1) = \mathbb{E}\left[y_{t+h} | \varepsilon_{1t} = \bar{\varepsilon}_1, x_{1t} = \bar{x}_1\right] - \mathbb{E}\left[y_{t+h} | \varepsilon_{1t} = 0, x_{1t} = \bar{x}_1\right]$$

Applying this definition to the interacted LP (2.6) yields

$$IRF^h_{y\varepsilon}(\bar{\varepsilon}_1, \bar{x}_1) = \beta^h_y \bar{\varepsilon}_1 + \gamma^h_y \bar{\varepsilon}_1 \times \bar{x}_1 \tag{2.7}$$

A standard LP-IRF depends only on the shock $\bar{\varepsilon}_1$. In our setting, the IRF is also a function of the propagation variable $\bar{x}_1$. This gives us scope to construct an IRF where the initial value of $\bar{x}_1$ can be varied counterfactually. For instance, a researcher may study how the shock $\bar{\varepsilon}_1$ affects the economy if the monetary authority did not respond to the shock in terms of some policy instrument i_t. In this case, one may think of $\bar{x}_1$ as a parameter (or combination of parameters) in the Taylor rule that governs the monetary response. The construction of the counterfactual for a pre-specified horizon h follows three steps.

1. Obtain the conditional IRF for the policy instrument i_{t+h}:

$$IRF^h_{i\varepsilon}(\bar{\varepsilon}_1, \bar{x}_1) = \beta^h_i \bar{\varepsilon}_1 + \gamma^h_i \bar{\varepsilon}_1 \times \bar{x}_1$$

2. Solve for $\bar{x}^h_1$ which satisfies: $IRF^h_{i\varepsilon}(\bar{\varepsilon}_1, \bar{x}^h_1) = 0$, $\forall \bar{\varepsilon}_1$:

$$\bar{x}^h_1 = -\beta^h_i / \gamma^h_i$$

3. The IRF for outcome y_{t+h} under $\bar{x}^h_1$ gives the counterfactual:

$$IRF^h_{y\varepsilon}(\bar{\varepsilon}_1, \bar{x}^h_1) = \beta^h_y \bar{\varepsilon}_1 + \gamma^h_y \bar{\varepsilon}_1 \times \bar{x}^h_1$$

One can see that this approach generates exact counterfactuals, i.e. the policy instrument does not react at all at the pre-specified horizon h. However, the policy instrument may well react at other horizons. If one is interested in the entire path of the policy instrument, one can adjust the second step as follows. Given some norm $|| \cdot ||$, one may minimizes deviations of the policy instrument for the entire path $h = 0, ..., H$.

$$\bar{x}^h_1 = \text{argmin}_{\tilde{x}} \; ||IRF_{i\varepsilon}(\tilde{x})|| \tag{2.8}$$

where $IRF_{i\varepsilon}(\tilde{x}) = \left(IRF^0_{i\varepsilon}(\varepsilon_{1,t}, \tilde{x}), ..., IRF^H_{i\varepsilon}(\varepsilon_{1,t}, \tilde{x}) \right)'$.
The value $\bar{x}^h_1$ is chosen to match all $H + 1$ targets as close as possible. The counterfactual

does not exactly set the policy instrument to zero, but admits to be constructed from a parsimonious empirical model.

2.3.2 Lucas Critique

The Lucas (1976) critique asserts that reduced-form estimates become invalid if agents perceive (understand) that systematic policy has changed. This is a consequence of optimizing agents who forecast future values of policy instruments. The underlying forecasting model generally affects the standard LP coefficient β^h (as in equation (2.4)). In our setting however, we explicitly model that policy x_{1t} (and x_{2t}) varies over time. Therefore, the Lucas critique only applies when agents' perceive that the *process* governing x_{1t} has changed. Intuitively, changes in x_{1t} do not cause expectations to change as long as they are sufficiently plausible to occur. Thus, we have scope to counterfactually vary x_{1t}. Note that this plausibility criterion cannot be formally tested because the change of this process is a zero probability event in our framework.

To investigate whether a given policy counterfactual is susceptible to the Lucas critique, we propose a heuristic approach, motivated by Leeper and Zha (2003). One may consider the following uni-variate auto-regressive model.

$$x_{1t} = ax_{1t-i} + b\varepsilon_{1t} + u_t \tag{2.9}$$

$\Delta^h = \bar{x}_1^h - ax_{1t-1} - b\varepsilon_{1t}$ gives the size of the innovations needed for $\bar{x}_1^h$ to occur, conditional on the shock ε_{1t} and the previous value x_{1t-1}. Thus, Δ^h can be compared with the distribution of error terms u_t in the spirit of a t-test. For instance, one may reject the null that $\bar{x}_1^h$ is sufficiently reasonable whenever $G(\Delta^h) \notin [0.025, 0.975]$ where $G(\cdot)$ is the CDF of the error terms. In practice, the (empirical) distribution of residuals can be used as $G(\cdot)$ is typically unknown and potentially time-dependent whenever errors are auto-correlated. However, notice that this is no formal test whether agents update their expectations. This is infeasible as the Lucas critique is concerned with a zero probability. Finally, it is important to emphasize that this testing procedure depends on x_{1t-1} which makes the robustness of the counterfactual state-dependent. In particular, the robustness depends on the (historical) episode for which the counterfactual is supposed to be used.

3 Systematic Monetary Policy in the US

Monetary policy is thought of having a systematic and a shock component. The systematic portion of policy is predictable and describes how strongly the monetary authority is perceived to respond to inflation and real activity (the policy reaction function), while shocks are random deviations from systematic monetary policy (Primiceri, 2005; Ramey, 2016). In this section, we propose a measurement of systematic US monetary policy based on Istrefi's (2019) narrative classification of FOMC members' policy preferences as hawks and doves. We further propose a FOMC rotation instrument to account for potential endogeneity in the FOMC's preference composition.

3.1 Narrative Account of Hawks and Doves at the FOMC

The FOMC is the body of the Federal Reserve System that sets US monetary policy. It consists of twelve members – the seven members of the Board of Governors of the Federal Reserve System, the president of the Federal Reserve Bank of New York and four of the remaining eleven FRB presidents, who serve one-year terms on a rotating basis. All 12 of FRB presidents attend FOMC meetings and participate in FOMC discussions, but only five of them who are Committee members at the time may vote on policy decisions. The seven members of the Board of Governors, including the Fed Chair, are nominated by the US President and confirmed by the US Senate. Governors are appointed for a 14-year term, however the Chair and Vice Chair of the Board serve only four-year terms, with the possibility for reappointment. In contrast, FRB presidents are appointed by their respective board of directors.[3] In our sample, between 1966 and 2007, FRB presidents have had an average term length of more than 12 years, compared to 9 years for governors.

The FOMC's monetary policy decision making involves the aggregation of diverse individual member preferences and views into a collective decision. Policy preferences are important in this process and, thus, a constant interest for academics (e.g. Rogoff, 1985) and for financial market participants and Fed watchers.[4] In the public debate central banker's preferences are often characterized as "hawks" – more concerned about inflation than unemployment – or "doves" – more concerned about unemployment than inflation. In the context of a Taylor rule, we can think about these preferences as a relatively larger coefficient on inflation (hawks) versus a relatively larger coefficient on the output gap (doves).

The challenge is that the true policy preferences of FOMC members are unobserved in practice. To our purpose, we use Istrefi (2019)'s measure of FOMC preferences that takes into account public perceptions on the type (hawk or dove) as formed in real time. This measure is based on narrative records in US newspapers regarding the policy leanings of each FOMC member with respect to the Fed's dual mandate: maximum employment and stable prices. Istrefi (2019) categorizes as hawk and dove about 93% of the 130 FOMC members who have served in the FOMC since the early 1960s, furthermore distinguishing those members that are perceived consistently as either hawks or doves (69% of the sample) and those perceived as switching camps over their tenure (i.e swingers, 24%).[5]

Istrefi (2019)'s Hawk-Dove classification is a panel across FOMC meetings and FOMC members. We aggregate the preferences of individual members in one measure, reflecting the balance between hawks and doves in the FOMC. We define $Hawk_{i\tau}$ as the policy preference

[3]Two thirds of the appointing directors are elected by commercial banks, which arguably makes their appointment less political than appointments of governors. In addition, the 5-year legal terms of presidents are commonly extended.

[4]Rogoff (1985) shows that society can make itself better off by appointing as head of an independent central bank an agent whose dislike for inflation relative to unemployment is known to be stronger than average. This agent is known as a 'conservative' central banker or a hawk in our terminology.

[5]The FOMC's Hawk-Dove index is based on the reading of about 20,000 articles or reports, from more than 30 newspapers and business reports of Fed watchers, referencing to 130 FOMC members that served between the period 1960 to 2015. For more details please see Istrefi (2019).

of FOMC member i in FOMC meeting τ. The number of voting members can vary across meetings but maximum is 12. We set $Hawk_{i\tau} = 1$ for a hawk and $Hawk_{i\tau} = -1$ for a dove. For every FOMC meeting, we compute the aggregate Hawk-Dove balance of policy preferences as follows:

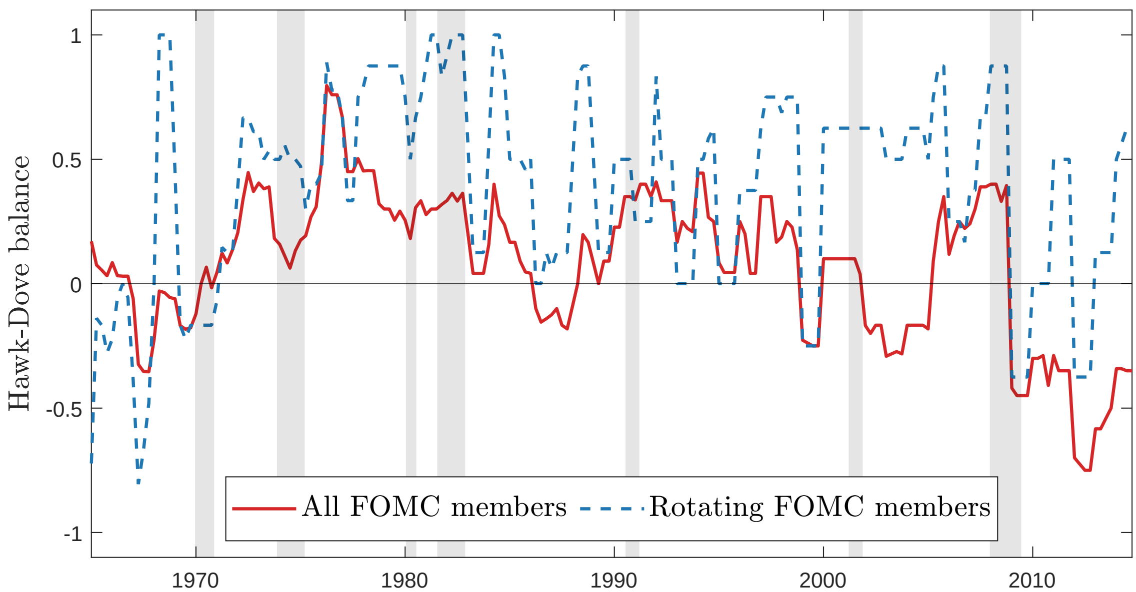
$$Hawk_{\tau} = \sum_{i} \alpha_i Hawk_{i\tau}. \tag{3.1}$$

where α_i represents the weight we give to our four types, hawk, dove, swinger hawk, swinger dove. We assign a full weight ($\alpha_i = 1$) to persistent hawks and doves, where persistent means that the member has been perceived to not change type during the whole tenure in the FOMC. We assign a smaller weight ($\alpha_i = 0.5$) to swingers, as their preference is perceived to change over time and likely dependent on the state of the economy (Bordo and Istrefi, 2021).[6] Finally, we aggregate $Hawk_{\tau}$ into monthly and quarterly frequency. If multiple FOMC meetings occurred in the same month or quarter, denoted by t, we compute $Hawk_t$ as the arithmetic average across $Hawk_{\tau}$ for all τ in t. Otherwise, we compute $Hawk_t$ as the last observed $Hawk_{\tau}$ until the end of month or quarter t.

The aggregate Hawk-Dove balance $Hawk_t$ represents our proxy measure of systematic Fed's monetary policy. This measure not only varies over time but is also perceived by the public. In the context of Section 2, we consider $Hawk_t$ as measure of x_{1t}.

[6] Other aggregation approaches are possible, e.g., a higher weight on the preference of the Fed chair or equal weight for swingers and non-swingers. We show that our results are robust under such alternative weights.

Figure 1: FOMC's Hawk-Dove Composition



Notes. Red solid line shows the cross section of FOMC members' policy preferences aggregated to quarterly frequency. The blue dashed line shows the same measure for the sub-group of FRB presidents that had voting rights within the FOMC. Grey bars indicate NBER dated recessions.

We present the evolution of the FOMC's Hawk-Dove balance from 1966 until 2014 in Figure 1 (the solid red line). It is noteworthy that there is considerable variation in this balance, featuring both hawkish and dovish majorities. There are several reasons behind this variation. A big part of it is due to the turnover of FOMC members (Governors particularly) and to the annual rotation of the FRB presidents. In addition, as this is a real time measure, there are new members for which the preference is unknown for some meetings and they are counted only when a perception on their type is formed.[7]

Istrefi (2019) shows that the FOMC's Hawk-Dove balance matches well with narratives of monetary policy in the U.S but also with 'true' policy tendencies, not known in real time to the public, as expressed by preferred interest rates (from FOMC transcripts in Chappell et al. (2005)), by forecasting patterns of individual FOMC members and by dissents.[8] Moreover,

[7]The series exhibits persistent fluctuations between a minimum of -0.75 (2012q3) an a maximum of 0.80 (1976q2) with an auto-correlation of 0.90.

[8]For a detailed discussion of this narrative please refer to Istrefi (2019).

Bordo and Istrefi (2021) show that the composition of the FOMC in terms of hawks and doves helps to explain deviations of the FFR from the path described by a conventional forward-looking Taylor rule.

3.2 FOMC Rotation Instrument

While we discussed some mechanical reasons that drive the time series variation in the Hawk-Dove balance (turnover of members, rotation of voting rights, yet unknown preferences), we can not rule out the possibility that $Hawk_t$ is partly driven by the state of the economy. The latter could be affected by shocks, be it policy, demand or supply shocks (cf. ε_t in Section 2), or by other systematic changes in macroeconomic propagation, e.g., through changes in systematic fiscal policy or financial conditions (cf. x_{2t} in Section 2).

Changes in the macroeconomic environment in t (ε_t or x_{2t}), may change the Hawk-Dove balance $x_{1t} = Hawk_t$ through the extensive and intensive margin. At the extensive margin, the macroeconomic environment may influence the choice for filling the vacant seats (Governor or FRB president) with members with a certain policy preference. For example, rising inflation might lead to the appointment of new governors, which are more hawkish than the governors they replace. At the intensive margin, $Hawk_t$ may change because existing FOMC members' preferences change. For example, in times of rising inflation, some previously dovish FOMC members might become more hawkish.

This endogeneity concern leads to biased OLS estimates of an interacted LP as in (2.6) that regresses some outcome, e.g., GDP, on the interaction of a shock of interest with $Hawk_t$. For example, if $Hawk_t$ is high, partly because fiscal policy has become more expansionary, then the correlation between GDP and the interaction may partly reflect the change in fiscal policy.

To address this concern, we propose a novel FOMC rotation instrument based on the rotation of voting rights of FRB presidents. As discussed earlier, four of the eleven FRB presidents (excluding the president of the New York Fed as a permanent voter) mechanically rotate voting rights on an annual basis. More precisely, some FRB presidents have voting rights only every second year (Cleveland and Chicago) or every third year (Philadelphia, Richmond and Boston; Dallas, Atlanta and St. Louis; Minneapolis, San Francisco and Kansas City). As the rotation of voting rights is mechanical and not dependent on the state of the economy, it creates exogenous variation in $Hawk_t$ at the extensive margin.

This rotation is important in practice, as the interest of Fed Watchers shows. Each year, before the first FOMC meeting with the new line up of the FOMC, media starts commenting on FRB presidents who leave and take voting rights, debating whether the new voters 'will rock the boat' or not. A typical quote reads as follows:

> *An annual rotation will strip regional Fed presidents of their voting rights [...] will bring a set of four different presidents to the table, [...] the addition of three decidedly hawkish individuals and one moderate dove has the potential to create one of the most vigilant monetary policy committees in recent memory."*
>
> FXCM, 25 January 2007

We construct the FOMC rotation instrument through

$$Hawk_{\tau}^{\mathrm{IV}} = \sum_{i \in Z_{\tau}} \alpha_i Hawk_{i\tau}, \tag{3.2}$$

where set Z_τ is the set of indices i of FRB presidents attending the FOMC as voting members at meeting τ and IV stands for instrument variable. We then aggregate $Hawk_{\tau}^{\mathrm{IV}}$ to monthly or quarterly frequency, similarly as for $Hawk_\tau$.

In Figure 1, the dashed line shows the time series of the FOMC rotation instrument $Hawk_t^{\mathrm{IV}}$, representing the Hawk-Dove balance of rotating FRB presidents. The instrument is typically above the overall FOMC Hawk-Dove balance, reflecting the fact that FRB presidents tend to be more hawkish than governors (Istrefi, 2019; Bordo and Istrefi, 2021). At the same time, we do observe large variation in the instrument, mainly due to the mechanic rotation of voting rights and the turnover (retirements and subsequent new appointments) of FRB presidents.

The instrument has a correlation of 0.51 with $Hawk_t$, which is important for the relevance of the instrument. To assess its relevance more formally, we regress $Hawk_t$ on the instrument $Hawk_t^{\mathrm{IV}}$ and perform the weak instrument test suggested by Olea and Pflueger (2013). The test rejects the null of weak instruments at 1 percent.

We further argue that the exclusion restriction is satisfied. First, a large source of variation in the instrument is due to the mechanic rotation. Second, FRB presidents are appointed by their respective board of directors (with the approval of the Board of Governors) and tend to have a substantially longer tenure than governors, shielding them from potential political cycle influence. In support of this, Bordo and Istrefi (2021) have shown that differently from Governors, there is no correlation between the preferences of the FRB presidents and the US president's party at the time of their appointment. FRB presidents are perceived mainly as hawks, irrespective of the US president's party (Republican or Democratic).

4 Monetary Policy Response to Inflation

In this section, we show that our measure of historical variation in systematic monetary policy in the US is meaningful, as the Fed responds significantly more aggressively to high expected inflation when the FOMC is more hawkish.

We show this effect by estimating a dynamic Taylor rule, where the dynamic response of the FFR, i_{t+h} depends on the Greenbook forecast of inflation (GDP deflator), $\hat{\pi}_t$, the Greenbook forecast of the unemployment rate $\hat{u}_t$, and on the perceived Hawk-Dove balance, $Hawk_t$.[9] Formally, we estimate interacted local projections, similar to (2.6),

$$i_{t+h} = \alpha^h + \beta_\pi^h \hat{\pi}_t + \beta_u^h \hat{u}_t + \gamma_\pi^h \left(\hat{\pi}_t \times Hawk_t\right) + \gamma_u^h \left(\hat{u}_t \times Hawk_t\right) + \delta^h Hawk_t + v_{t+h}^h, \quad h \geq 0. \quad (4.1)$$

[9] Greenbook forecasts are prepared by the Federal Reserve Board staff ahead of each FOMC meeting, and as such they reflect the information set available to the FOMC members before deliberating and deciding on policy. We use the arithmetic average of forecasts over the current and subsequent month. The inflation forecast is based on the GDP deflator.

The mean effects β_π^h and β_u^h measure the average response of systematic monetary policy to increased inflation and unemployment forecasts respectively. Our proxy $Hawk_t$ is demeaned and therefore captures deviation in systematic monetary policy from the sample average.[10] In particular, interaction terms, $\hat{\pi}_t \times Hawk_t$ and $\hat{u}_t \times Hawk_t$ show how much stronger the Fed responds when the FOMC is more hawkish. Additionally, $Hawk_t$ is normalized to have unit standard deviation and forecasts $\hat{\pi}_t$ and $\hat{u}_t$ are demeaned. Both choices ease the interpretation of the regression coefficients. We estimate (4.1) using monthly data from 1966m1 until 1996m12.

Figure 2 presents the response of the FFR to a one percentage point increase in the inflation forecast, taking into account the estimates for inflation coefficients β_π^h and γ_π^h. Panels (a) and (b) show OLS estimates and panels (c) and (d) show IV estimates using the FOMC rotation instrument, $Hawk_t^{\mathrm{IV}}$.[11] In both figures, the left panels shows the impulse responses of the FFR for different values of $Hawk_t$. In particular, we vary systematic monetary policy from (one standard deviation) more hawkish ($\beta_\pi^h + \gamma_\pi^h$) to one standard-deviation more dovish ($\beta_\pi^h - \gamma_\pi^h$), relative to the sample average (β_π^h). The right panels shows the corresponding interaction effect, γ_π^h, with 68 and 95 percent confidence bands.

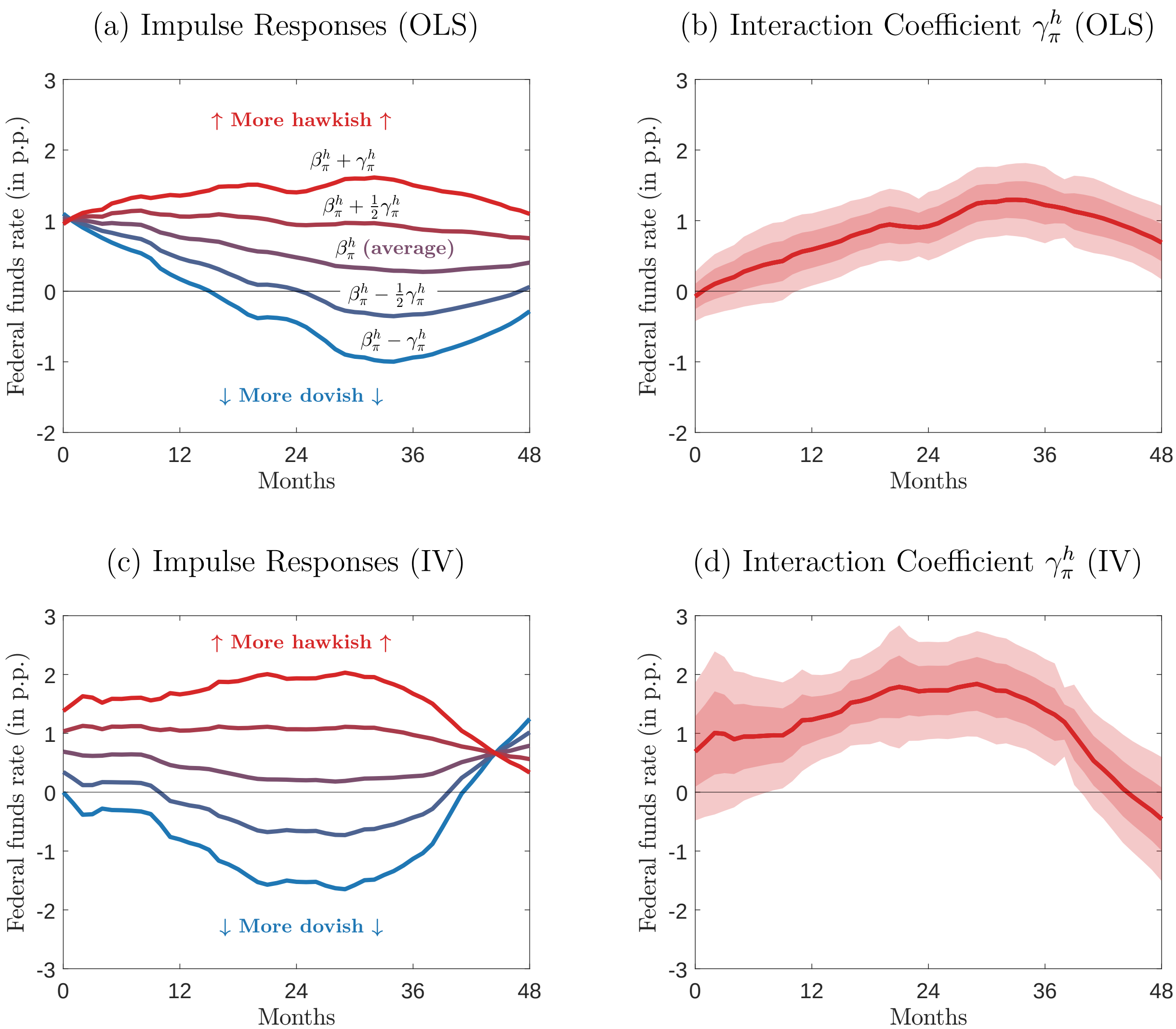
The central line of panel (a) shows that high inflation forecasts are on average followed by a higher FFR. In addition, a more hawkish FOMC responds by a prolonged rate hike to increased inflation forecasts. In contrast, a more dovish FOMC lets the FFR decay much faster. For example, a (one standard deviation) more hawkish FOMC holds the FFR by 0.90 percentage points higher two years after a one percentage point increase in the inflation forecast. The differences across the more hawkish (dovish) lines are significant at horizons longer than $h = 12$ months and strikingly different (see panel (b)).

Note that the estimates of β_π^h and γ_π^h do not condition on the source of high inflation, but are conditional on the forecasted unemployment rate. Thus, our estimates should be understood

[10] For simplicity, we will continue to refer to $Hawk_t$ as proxy for systematic monetary policy with the understanding that this refers to deviations from average systematic monetary policy in our sample.

[11] We use $Hawk_t^{\mathrm{IV}}$ and its interaction with both forecasts as instruments for regressors involving $Hawk_t$.

Figure 2: FFR Response to Greenbook Inflation Forecast



Notes. Based on estimates of β_π^h and γ_π^h in (4.1). These estimates reflect the effect of a one percentage point increase in the inflation forecast. The left panels present IRFs where systematic policy varies from one standard-deviation more hawkish ($\beta_\pi^h + \gamma_\pi^h$) to one standard-deviation more dovish ($\beta_\pi^h - \gamma_\pi^h$). The right panels show the estimate of the interaction coefficient (γ_π^h) with 68 and 95 percent confidence bands, based on Newey-West standard errors.

as business cycle correlations and not as causal effects of inflation on the FFR.

Panels (c) and (d) of Figure 2 present the IV estimates of β_π^h and γ_π^h. Overall, using the FOMC rotation instrument to account for endogeneity in $Hawk_t$ leads to similar conclusion as the OLS estimates. The estimated differential effect of a more hawkish FOMC, γ_π^h, are again highly significant. Compared to panel (b), the IV estimate in (d) attains large values at a shorter lag and is slightly larger at its peak.

Overall, we consider these estimates as strong evidence that the FFR response to inflation crucially depends on $Hawk_t$. This outcome is in line with theoretical predictions, e.g., from a New Keynesian model with a time-varying Taylor rule as described in Appendix A.1. These results suggest that our measure of historical variation in systematic monetary policy captures important aspects of Fed's monetary policy making.[12]

5 Fiscal Spending Shocks and Monetary Policy

In this section, we apply our approach of constructing policy counterfactual to an important question in the literature: How does monetary policy affect the response of the economy to fiscal spending shocks in the US? More specifically, we investigate the extent to which systematic monetary policy shapes the effects of fiscal spending shocks and the size of fiscal spending multipliers.

5.1 Impulse Responses to Fiscal Shocks

5.1.1 Setting

We estimate the dynamic effects of fiscal spending shocks ε_t^g on some outcome y_{t+h} when accounting for measured systematic monetary policy, $Hawk_t$, and a vector of control vari-

[12]Figure B.1 in Appendix, shows estimates with respect to the unemployment forecast. The results are less clear, as most estimates are either zero or slightly negative but also insignificant. We therefore suppress a more detailed discussion and conclude that we find no meaningful responses on the unemployment forecast.

ables, X_t, formally as below

$$y_{t+h} = \alpha^h + \beta^h \varepsilon_t^g + \gamma^h \left(\varepsilon_t^g \times Hawk_t\right) + \delta^h Hawk_t + \zeta^h X_t + v_{t+h}^h, \quad h \geq 0. \tag{5.1}$$

This is an interacted local projection as in (2.6). As response y_{t+h}, we consider real GDP and real government spending, both relative to potential output, and the FFR.[13] We use the military spending news shock form Ramey and Zubairy (2018) as our baseline fiscal spending shock. Finally, the vector of control variables X_t includes 4 lags of real GDP, real government spending, and the fiscal shock. Our LP specification, construction of variables, and choice of control variables follows Ramey and Zubairy (2018) who study how the state of the business cycle and the zero lower bound shape the effects of fiscal spending shocks.

We estimate (5.1) on data from 1966Q1 through 2007Q4. We start in 1966 because this is the earliest year for which we measure $Hawk_t$.[14] By ending our sample in 2007, we avoid the Great Recession and the period where the Federal Reserve has increasingly used unconventional monetary policy with the FFR stuck at the zero lower bound. However, we show that our results are robust to extending the sample until 2014. In the following we present our results for the OLS and IV estimator. The IV estimator uses the FOMC rotation instrument analogously to Section 4.

5.2 Results

Figures 3 and 4 show the estimated effects of a fiscal spending shock – precisely, military spending news amounting to 1% of potential GDP – conditional on the hawkishness of the FOMC. In both figures, the left panels shows the impulse responses of a particular outcome variable for different values of $Hawk_t$. In particular, we vary systematic monetary policy from

[13]Following Ramey and Zubairy (2018), we estimate potential output as the explained part of real GDP when fitting a polynomial of degree 6 in time to real GDP.

[14]In principle, we could extend the $Hawk_t$ series back in time by expanding the narrative approach in Istrefi (2019). However, prior to 1966 news coverage of individual FOMC members is relatively sparse, which would leave us with a unreliable measure of $Hawk_t$ for periods before 1966.

(one standard deviation) more hawkish ($\beta^h + \gamma^h$) to (one standard deviation) more dovish ($\beta^h - \gamma^h$), relative to the sample average (β^h). The right panels shows the corresponding interaction effect, γ^h, with 68 and 95 percent confidence bands.

Federal Funds Rate. Figure 3, panel (a), shows the FFR increasing on average by around 50 basis points one year after the expansionary fiscal shock. Importantly, we find that a more hawkish FOMC is associated with a larger increase in the FFR (see panel (b) of Figure 3). The FFR increases by up to 2 percentage points after one year when the FOMC is one standard-deviation more hawkish.

Panels (a) and (b) of Figure 4 show the same responses when addressing endogeneity concerns with our IV estimator. The results confirm that a (one standard-deviation) hawkish FOMC raises the FFR by around 2 percentage points within a year. Conversely, a more dovish FOMC even cuts the FFR in response to this shock.

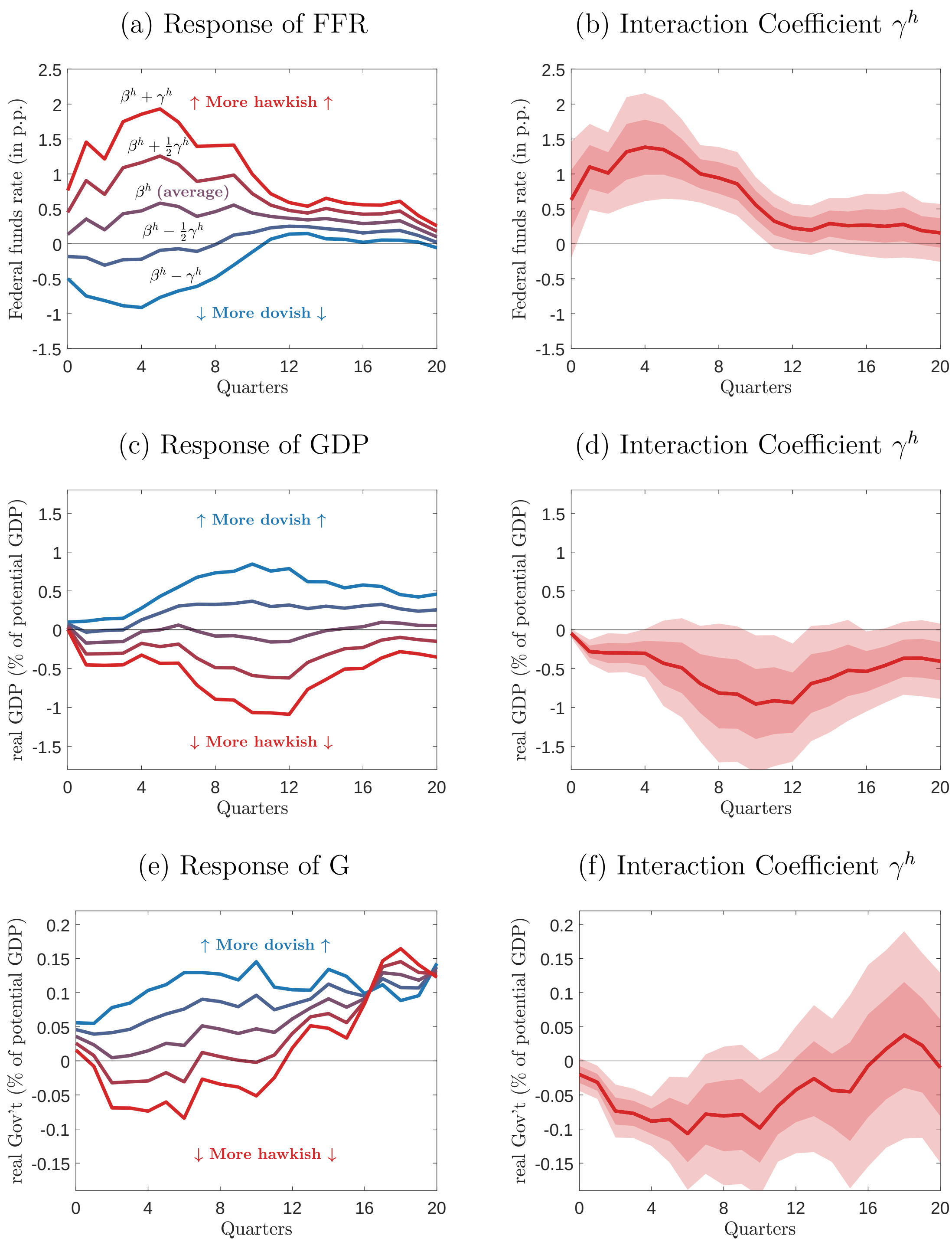
These findings are consistent with the validation exercise in Section 4 and suggest a meaningful role of systematic monetary policy for the effects of fiscal spending shocks.

Real GDP and Real Government Spending. The FFR hike dampens aggregate demand and counteracts the expansionary fiscal impulse. Interestingly, our estimates for real GDP in Figure 3 imply that, on average, the monetary authority almost perfectly stabilizes output, as the estimate for β^h is not significantly different from zero. Further, the government spending response is only modest suggesting that the fiscal authority reduces non-war related expenses, perhaps due to higher interest rates.[15]

In stark contrast with the small average responses of GDP and government spending, a more hawkish or dovish systematic monetary policy has large effects on these responses. Figure 3, panel (d), shows that GDP falls by up to 1% more when the FOMC is more hawkish. The

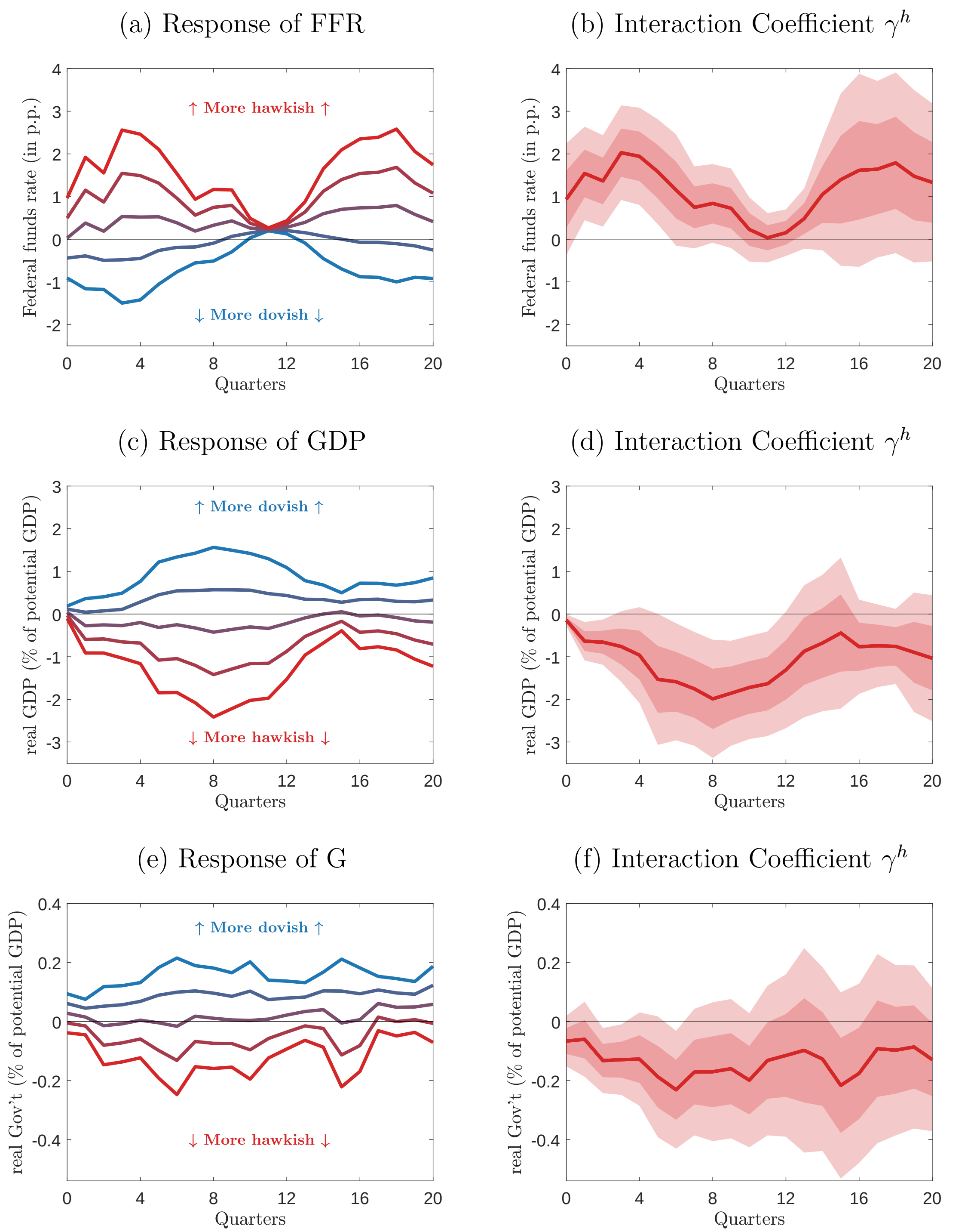
[15] It is noteworthy that we obtain insignificant average effects of GDP and government spending because we omit the Korean war from the sample. The military spending news shock has substantially less explanatory power when the Korean war is excluded, as Ramey (2011) remarks. We cannot accommodate a longer time series however because Istrefi's (2019) narrative account starts in 1960 only.

Figure 3: Conditional Impulse Responses to Fiscal Spending Shock (OLS)



Notes: Based on OLS estimates of β^h and γ^h in (5.1). These estimates reflect the effect of military spending news amounting to 1% of potential GDP. The left panels present IRFs where systematic policy varies from one standard-deviation more hawkish ($\beta^h + \gamma^h$) to one standard-deviation more dovish ($\beta_{h-\gamma^h}$). The right panels show the estimate of the interaction coefficient with 68 and 95 percent confidence bands, based on Newey-West standard errors.

Figure 4: Conditional Impulse Responses to Fiscal Spending Shock (IV)



Notes: Based on IV estimates of β^h and γ^h in (5.1). These estimates reflect the effect of military spending news amounting to 1% of potential GDP. The left panels present IRFs where systematic policy varies from one standard-deviation more hawkish ($\beta^h + \gamma^h$) to one standard-deviation more dovish ($\beta_{h-\gamma^h}$). The right panels show the estimate of the interaction coefficient with 68 and 95 percent confidence bands, based on Newey-West standard errors.

IV estimate in Figure 4, panel (d), even reaches 2%. Similarly, the response of government spending is between 0.1% and 0.2% lower with a more hawkish FOMC. These differential effects are large but may not be entirely surprising given the large differential response of the FFR.

Overall, our estimates suggest that systematic monetary policy plays a crucial role for successful fiscal stimulus, consistent with theory (e.g., Woodford, 2011) and recent empirical evidence (e.g., Canova and Pappa, 2011; Cloyne et al., 2021). Our results thus highlight a source of state-dependence which is qualitatively and quantitatively important.

Our results show that a dovish (hawkish) FOMC amplifies (depresses) aggregate demand, through their FFR reaction to the fiscal shock. Thus, it is natural to ask what the effects of fiscal stimulus are, when the FOMC is nonresponsive to fiscal shocks in terms of the FFR. The next section estimates the fiscal spending multiplier in such a counterfactual scenario.

5.3 Counterfactual Fiscal Multipliers

5.3.1 Setting

A key object for the evaluation of fiscal policies is the fiscal spending multiplier. This multiplier is defined as the dollar amount by which GDP increases, per dollar fiscal spending (both in real terms). Typically, one compares the cumulative change in GDP with the cumulative change in government spending over some horizon h (e.g., Mountford and Uhlig, 2009; Ramey and Zubairy, 2018). Formally, the cumulative multiplier at date t is given by

$$M_t^h = \frac{\sum_{j=0}^{h} GDP_{t+j}}{\sum_{j=0}^{h} G_{t+j}}, \tag{5.2}$$

where GDP_t and G_t denote GDP and government spending respectively. We could construct the fiscal multiplier M_t^h via the estimates in the previous subsection, by summing up the responses for GDP and government spending. However, a more efficient one-step procedure is to directly estimate the cumulative responses, which we obtain through a cumulative version of (5.1). Formally, we consider

$$\sum_{j=0}^{h} y_{t+j} = \tilde{\alpha}^h + \tilde{\beta}^h \varepsilon_t^g + \tilde{\gamma}^h \left(\varepsilon_t^g \times Hawk_t\right) + \tilde{\delta}^h Hawk_t + \tilde{\zeta}^h X_t + \tilde{v}_{t+h}^h, \quad h \geq 0, \tag{5.3}$$

where y_t is either GDP or government spending. We use the same data as in Section 5.1. We can then construct a state-dependent fiscal multiplier via

$$M^h(\overline{Hawk}) = \frac{\tilde{\beta}_{GDP}^h + \tilde{\gamma}_{GDP}^h \overline{Hawk}}{\tilde{\beta}_G^h + \tilde{\gamma}_G^h \overline{Hawk}}, \tag{5.4}$$

where $\overline{Hawk}$ is a given hawkishness of systematic monetary policy. The subscripts GDP and G indicate the coefficients in (5.3), when y_t is GDP or government spending, respectively. Importantly, this multiplier is a function of our measure of systematic monetary policy, $\overline{Hawk}$.

The state-dependent multiplier $M^h(\overline{Hawk})$ allows us to construct monetary policy counterfactuals. In particular, we estimate the size of the fiscal multiplier in the counterfactual scenario in which the monetary authority is non-responsive to the fiscal spending shock, i.e., it keeps the FFR fixed. Following the algorithm in Section 2.3, we compute the hawkishness $\overline{Hawk}^h$ that is required to keep the response of the FFR to the fiscal spending shock in minimum distance, defined by the Euclidean norm, from zero over a horizon up to h quarters after the shock. This counterfactual is of interest because it delivers a *pure* fiscal multiplier which is not (directly) affected by monetary policy.

5.3.2 Results

Table 1 presents our main results. More specifically, it presents the OLS and IV estimates of the numerator and denominator in equation (5.4), as well as the implied cumulative fiscal multiplier and the value $\overline{Hawk}^h$ that implements the policy counterfactual.[16] We conduct

[16]The OLS and IV estimates of numerator and denominator in (5.4) broadly conform with the non-cumulative estimates in Section 5.1.

the exercise for forecast horizons h between one and five years after the shock, as shown in the columns of the table. The chosen forecast horizon corresponds with the horizon over which we set $\overline{Hawk}^h$ to minimize the federal funds rate response.

Table 1: Counterfactual Fiscal Spending Multipliers

Horizon	h=4	h=8	h=12	h=16	h=20
			OLS		
Fiscal multiplier	-0.5298	1.7692	2.4843	2.5618	2.2888
Hawk-Dove balance	-0.0648	-0.0801	-0.0903	-0.0946	-0.0978
Response of GDP	-0.0009	0.0083	0.0201	0.0311	0.0410
	(0.0033)	(0.0080)	(0.0122)	(0.0181)	(0.0229)
Response of G	0.0017	0.0047	0.0081	0.0122	0.0179
	(0.0010)	(0.0021)	(0.0032)	(0.0035)	(0.0043)
			IV		
Fiscal multiplier	-2.4152	0.5228	1.8313	2.8556	3.0527
Hawk-Dove balance	-0.0479	-0.0571	-0.0611	-0.0710	-0.0804
Response of GDP	-0.0031	0.0018	0.0106	0.0264	0.0431
	(0.0056)	(0.0156)	(0.0210)	(0.0251)	(0.0310)
Response of G	0.0013	0.0035	0.0058	0.0092	0.0141
	(0.0011)	(0.0024)	(0.0036)	(0.0045)	(0.0045)

Notes: Based on OLS and IV estimates of $\tilde{\beta}^h$ and $\tilde{\gamma}^h$ in (5.3). These estimates reflect the effect of military spending news amounting to 1% of potential GDP. Columns 1 to 5 show different counterfactuals in which we minimize the FFR response until h quarters after the shock, by setting an appropriate value of $\overline{Hawk}^h$. The first and second rows show the cumulative fiscal spending multiplier and the value $\overline{Hawk}^h$ to implement the counterfactual. The third and fourth row show the associated cumulative GDP and government spending response until horizon h, with Newey-West standard errors in parenthesis.

We find that, implementing a policy counterfactual characterized by a non-responsive monetary policy requires a more dovish FOMC compared to the sample average. Across forecast horizons, we require $\overline{Hawk}^h$ between -0.04 and -0.10.

Further, Section 2.3 discusses that the robustness of the counterfactual hinges on a plausible value for $\overline{Hawk}^{h}$, given the empirical distribution of the $Hawk_t$ series. We can expect that the counterfactual is robust to the Lucas critique when $\overline{Hawk}^{h}$ is not unlikely to occur. In fact, the $\overline{Hawk}^{h}$ are relatively small compared to a standard deviation of $Hawk_t$ of 0.23. The fiscal multipliers associated with $\overline{Hawk}^{h}$ build up over time. The estimated multipliers are approximately of size 2 after three to four years, for both the OLS and IV estimates. This is large compared with an average fiscal multiplier close to zero for such horizons in our sample. Our estimates tend to be significant for OLS but less for the IV estimator, especially for short horizons. Overall, our results suggest that fiscal multipliers are sizeable if the FOMC is unresponsive.

We show the average and counterfactual responses to the fiscal shock for $h = 12$ in Figure B.2 in the Appendix. The counterfactual response of the FFR is indeed close to zero between 0 and 12 quarters after the shock. It slightly undershoots in the first few quarters and overshoots thereafter, compared with the target of setting this response to zero. We further find that the responses of both GDP and the government spending are persistently larger in the counterfactual scenario.

Finally, we relate our fiscal multiplier estimates to the literature. First, our counterfactual multipliers are large compared with estimates of average multipliers which tend to be below unity (e.g. Mountford and Uhlig, 2009; Ramey, 2011). Our estimates are also larger than fiscal multipliers during recessions (e.g. Auerbach and Gorodnichenko, 2012, 2013; Bachmann and Sims, 2012) or at the zero lower bound (e.g. Ramey and Zubairy, 2018). However, our findings are broadly consistent with the previous literature that can speak to the role of monetary policy. Canova and Pappa (2011) find fiscal multipliers around 2 in the short run. Nakamura and Steinsson (2014) estimate open economy multipliers which are also free of the monetary response but also free of (nationwide) general equilibrium effects. They find cumulative fiscal multipliers between 1.5 and 1.8 after two years, which is similar to our estimates. Cloyne et al. (2021) find cross-country multipliers of up to 2 depending on systematic monetary policy. We note that in comparison to these existing studies, we deliver direct evidence on the cumulative US fiscal multiplier, explicitly addressing identification.

5.4 Robustness

5.4.1 Alternative Specifications

Below, we explore the robustness of our results and show the resulting multiplier estimates in Table 2. In particular, we show counterfactual multipliers for $h = 12$ which can be compared with our main results in Table 1, (column 3).

Fed Chair Weight. In our main specification, the Fed chair carries the same weight in the Hawk-Dove composition as any other FOMC member with voting rights. However, several papers have stressed the importance of the Fed Chair in setting the FOMC agenda and building consensus (e.g. Sims and Zha, 2006; Bianchi et al., forthcoming). The first column of Table 2 shows fiscal multiplier estimates when we double the weight of the Fed chair in the aggregate Hawk-Dove balance of policy preferences. Both, the OLS and the IV estimate for the fiscal multiplier is very similar to the estimates presented in Table 1.

Swinger Weights. Above, we assume that the weight for swingers was only one-half whereas it was assumed to be unity for FOMC members with permanent (constant) policy preferences. We consider an alternative scenario where all FOMC members receive the same weight to explore whether this choice was consequential for our results. Column 2 of Table 2 gives the results which are very similar to the results we found before.

Include Great Recession. We cut the sample at the end of 2007 due to the onset of the great financial crisis. We made this choice because this period is characterized by exceptional monetary policy behavior with unconventional policy measures. In particular, for several years after the crisis the FFR has been stuck at the zero lower bound, and thus, is not a suitable measure of monetary policy. Yet, in our set up, the FFR response is crucial for the construction of monetary policy counterfactuals. We expand the sample until the end of 2014 to explore the robustness of our results, irrespective of unconventional policy measures.

Table 2: Robustness of Counterfactual Multipliers for $h = 12$

Setting	Fed Chair weight	Swinger weight	Great Recession
		OLS	
Fiscal multiplier	2.6001	2.3524	-0.9829
Hawk-Dove balance	-0.0776	-0.1083	0.0905
Response of GDP	0.0204	0.0192	-0.0052
	(0.0112)	(0.0153)	(0.0180)
Response of G	0.0078	0.0082	0.0053
	(0.0033)	(0.0029)	(0.0043)
		IV	
Fiscal multiplier	1.8555	1.8247	3.3760
Hawk-Dove balance	-0.0481	-0.0861	-0.0041
Response of GDP	0.0106	0.0120	0.0303
	(0.0231)	(0.0239)	(0.0132)
Response of G	0.0057	0.0066	0.0090
	(0.0038)	(0.0030)	(0.0042)

Notes: Based on OLS and IV estimates of $\tilde{\beta}^h$ and $\tilde{\gamma}^h$ in (5.3). These estimates reflect the effect of military spending news amounting to 1% of potential GDP. We show counterfactuals in which we minimize the FFR response until 12 quarters after the shock, by setting an appropriate value of $\overline{Hawk}^{\,h}$. In column 1, we double the weight of the Fed chairman in the aggregation. In column 2, we aggregate with a constant weight and therefore ignore swingers. In column 3, we expand the sample until 2014q4. The first and second rows show the cumulative fiscal spending multiplier and the value $\overline{Hawk}^{\,h}$ to implement the counterfactual. The third and fourth row show the associated cumulative GDP and government spending response until horizon h, with Newey-West standard errors in parenthesis.

Column 3 of Table 2 shows the resulting multipliers. The IV estimate becomes substantially larger compared with our main results. Yet, the required hawkishness declined by approximately 50 percent. One reason for this larger multiplier might be the zero lower bound period which has been shown to amplify multipliers (e.g. Woodford, 2011; Ramey and Zubairy, 2018). The OLS estimates turn even negative, due to a negative GDP response. This is driven by the FFR response which is completely unreliable in the extended sample. In fact, the FFR response implies that a more hawkish FOMC is required to implement the counterfactual. Thus, it is not surprising that GDP contracts. However, the seemingly different multipliers can easily be reconciled with our main estimates in Section 5.3.2. When one is willing to estimate the FFR response on the reduced sample until 2007q4 to obtain $\overline{Hawk}^{\,h}$ but uses the extended sample until 2014q4 for GDP and government spending, one obtains fiscal multipliers close to 2, after three years. This holds true for both, the OLS and the IV estimator.

6 Conclusion

This paper studies how monetary policy shapes the effects of macroeconomic shocks. We provide a framework for identification of monetary policy counterfactuals that leverages observed time variation in systematic monetary policy. The narrative account by Istrefi (2019) provides a proxy for systematic monetary policy and allows us to apply our framework. Identification is achieved using a novel FOMC rotation instrument, based on the rotation of voting rights within the FOMC. We apply our framework to study fiscal-monetary interactions empirically. We document that the responses of GDP, government spending and the federal funds rate depend strongly and significantly on monetary policy. In particular, the cumulative US fiscal spending multiplier after three years increases from close to 0 to 2 in a counterfactual where monetary policy does not respond to fiscal shocks.

Our paper suggests that the consequences of discretionary fiscal policy crucially depend on the monetary policy response. Thus, our results imply that there have been historical episodes with both, large positive but also negative fiscal multipliers. This may tempt policy makers to increase discretionary spending in periods of large multipliers, i.e. when monetary policy is dovish. Yet, we caution that the Lucas (1976) critique is lurking: Our estimates retain validity only if the fiscal authority does not try to exploit dovish monetary regimes to a greater extend than in the historical sample. This limitation echos the original Lucas critique in our setting. Conversely, our estimates retain valid for different configurations of monetary policy when fiscal policy behaves similar as in the historical sample. We therefore consider our estimate as important empirical benchmark to quantify the role of systematic US monetary policy for the consequences of fiscal shocks. Further, our general framework admits to study other non-policy shocks and systematic components of other policies. Expanding along these lines remains an important avenue for further research.

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A Theory: Proofs and Discussion

A.1 New-Keynesian Example

We next show that our results in Sections 2.1-2.3 apply to a New-Keynesian model. We consider a simple non-linear modification of the textbook New-Keynesian (NK) model outlined in Galí (2015, chapter 3).

$$\pi_t = \beta \mathbb{E}_t \pi_{t+1} + \kappa y_t \tag{A.1}$$

$$y_t = -\frac{1}{\sigma}\left(i_t - \mathbb{E}_t \pi_{t+1} - r_t^n\right) + \mathbb{E}_t y_{t+1}, \qquad r_t^n = -\psi_{na} a_t + \psi_{ng} g_t \tag{A.2}$$

$$i_t = \phi_t \pi_t + v_t \tag{A.3}$$

The variables π_t, y_t and i_t denote inflation, the output gap and the real interest rate, all in log-deviations from steady state. Further, a_t, g_t and v_t are exogenous state variables, which follow AR(1) processes, with the respective innovations being structural shocks to technology, government spending and monetary policy.

The non-linear modification is that the Taylor rule coefficient is time-variant

$$\phi_t = \bar{\phi} + e^{x_t} \tag{A.4}$$

$$x_t = a x_{t-1} + b \varepsilon_t^u + c \eta_t \tag{A.5}$$

In the context of the general framework presented in Sections 2.1-2.3, in this model x_t is a scalar, which describes systematic monetary policy. The equilibrium conditions can be represented more compactly in matrix notation

$$\begin{aligned}
\begin{pmatrix} y_t \\ \pi_t \end{pmatrix} &= \omega(x_t) \begin{pmatrix} \sigma & 1 - \beta\left(\bar{\phi} + e^{x_t}\right) \\ \sigma\kappa & \kappa + \beta\sigma \end{pmatrix} \begin{pmatrix} \mathbb{E}_t\left[y_{t+1}\right] \\ \mathbb{E}_t\left[\pi_{t+1}\right] \end{pmatrix} + \omega(x_t) \begin{pmatrix} 1 \\ \kappa \end{pmatrix} u_t \\
&= A(x_t) \begin{pmatrix} \mathbb{E}_t\left[y_{t+1}\right] \\ \mathbb{E}_t\left[\pi_{t+1}\right] \end{pmatrix} + B(x_t) u_t,
\end{aligned} \tag{A.6}$$

where $\omega(x_t) = \left(\sigma + \kappa\left(\bar{\phi} + e^{x_t}\right)\right)^{-1}$ and $u_t = \rho_u u_{t-1} + \varepsilon_t^u$ is an auto-correlated innovation, which is a function of the three structural shocks.

In the standard linear model, the solution to (A.6) is a mapping from u_t to y_t and π_t. The unknowns of this mapping are the coefficients of the mapping, which can be determined via guess and verify. In our non-linear model, the coefficients of this mapping are functions that depend on x_t. In particular, we guess the following mapping.

$$\begin{pmatrix} y_t \\ \pi_t \end{pmatrix} = \begin{pmatrix} \psi_y(x_t) \\ \psi_\pi(x_t) \end{pmatrix} u_t + \begin{pmatrix} \Omega_y(x_t) \\ \Omega_\pi(x_t) \end{pmatrix}. \tag{A.7}$$

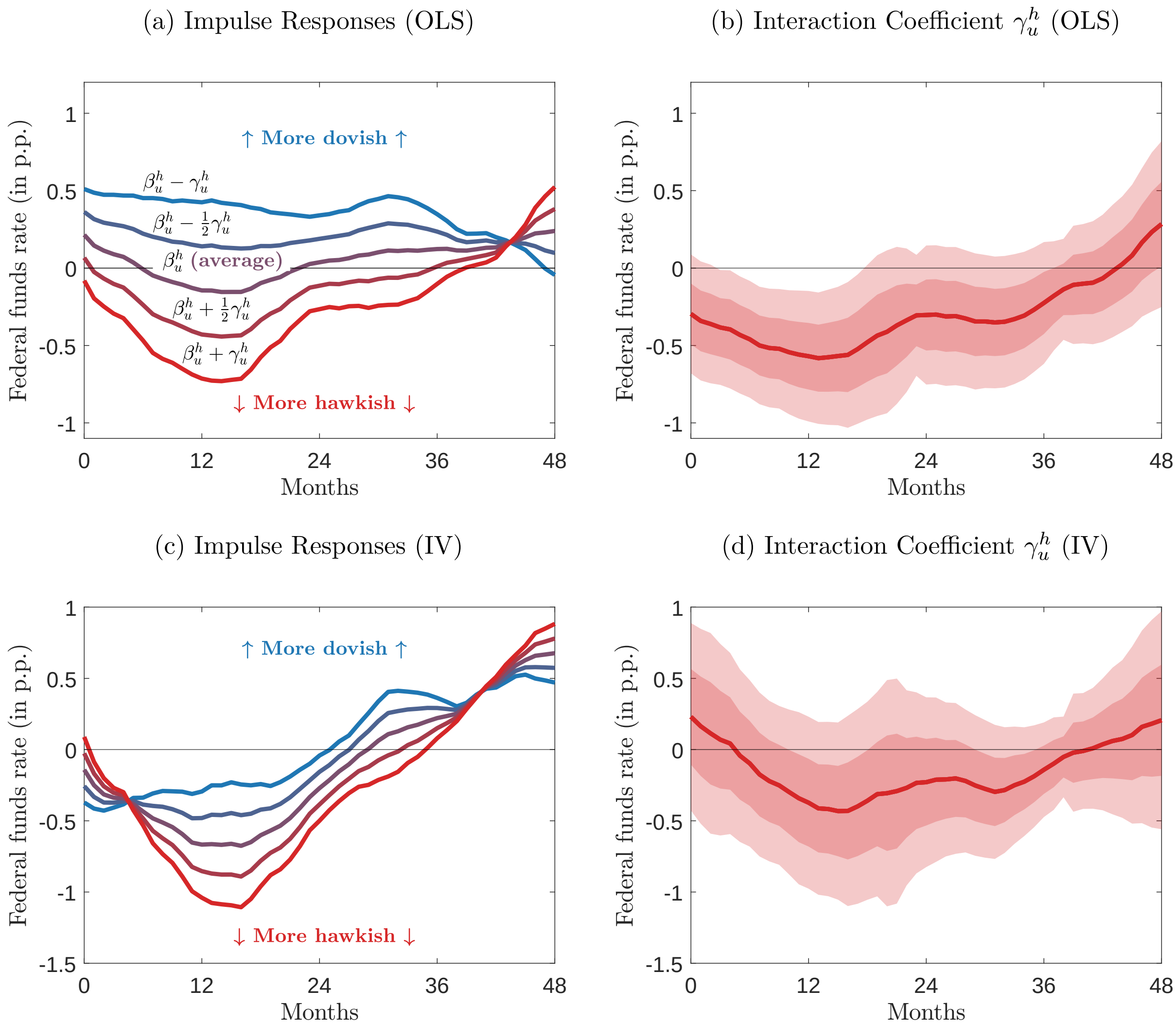
Substituting this guess into (A.6) we obtain

$$\begin{aligned} \begin{pmatrix} y_t \\ \pi_t \end{pmatrix} &= \left\{ A(x_t) \begin{pmatrix} \mathbb{E}_t \left[\psi_y(x_{t+1}) \rho_u \right] \\ \mathbb{E}_t \left[\psi_\pi(x_{t+1}) \rho_u \right] \end{pmatrix} + B(x_t) \right\} u_t \\ &\quad + A(x_t) \begin{pmatrix} \mathbb{E}_t \left[\psi_y(x_{t+1}) \varepsilon^u_{t+1} + \Omega_y(x_{t+1}) \right] \\ \mathbb{E}_t \left[\psi_\pi(x_{t+1}) \varepsilon^u_{t+1} + \Omega_\pi(x_{t+1}) \right] \end{pmatrix}, \end{aligned} \tag{A.8}$$

which verifies the guess. The solution fits our general DGP in equation (2.1). Thus, our results in Sections 2.1-2.3 apply, provided that the maintained assumptions are satisfied. Equation (A.5) directly implies Assumption 1 when $|a| < 1$. Additionally, Assumption 2 must be imposed on the structural shock ε^u_t and on innovation η_t. Finally, Assumption 3 demands that x_t is simply proportional to η_t which makes OLS consistent. Similarly, Assumption 4 can be satisfied with an instrumental variable being proportional to η_t which delivers a consistent IV estimator.

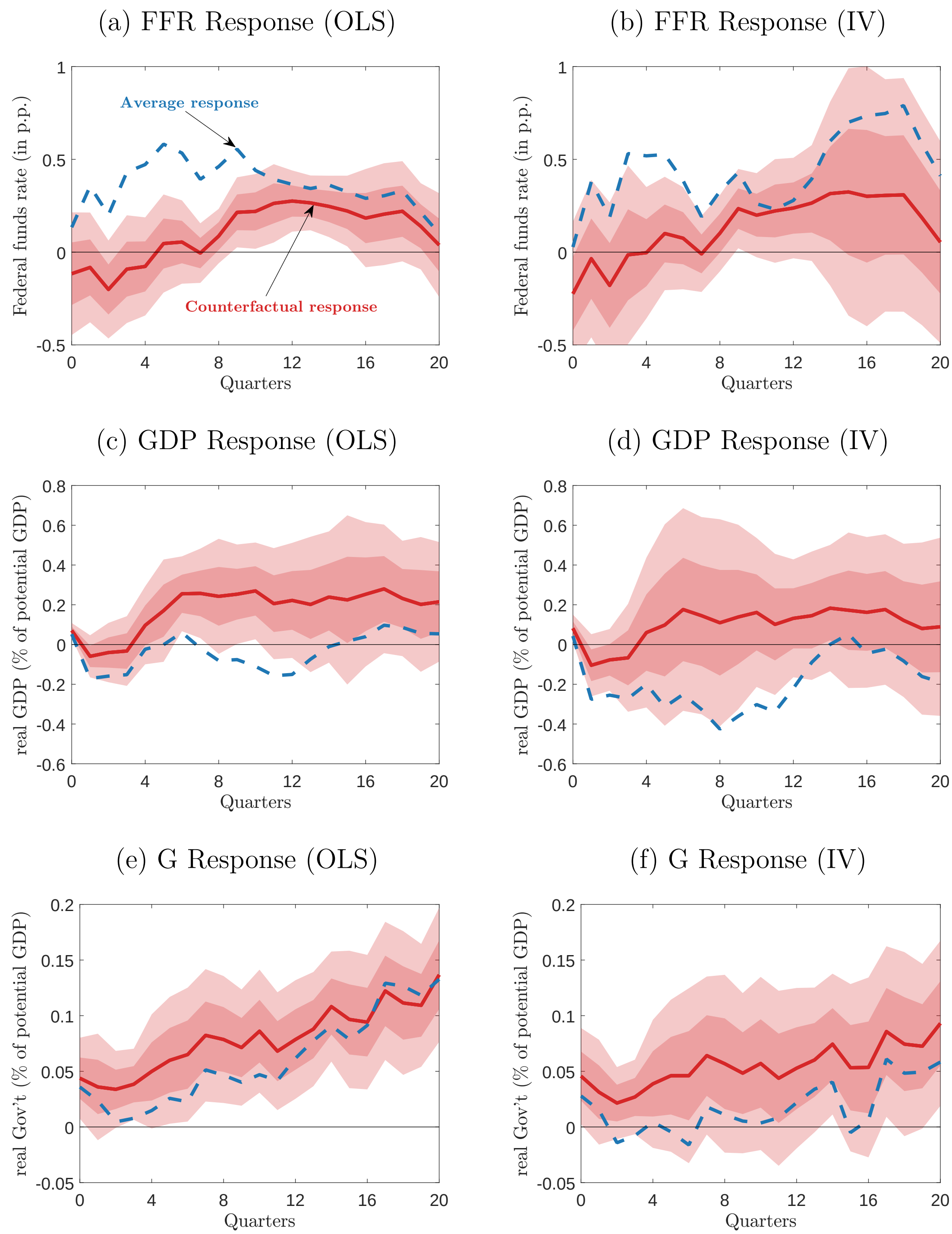
B Results

Figure B.1: FFR Response to Unemployment Forecast



Notes. We show estimates of (4.1). The responses correspond to a one percentage point innovation in the unemployment forecast. The left panels present IRFs where systematic policy varies from one standard-deviation more hawkish ($\beta_u^h + \gamma_u^h$) to one standard-deviation more dovish ($\beta_u^h - \gamma_u^h$), relative to the sample average. The right panels show the estimate of the interaction coefficient with 68 and 95 percent confidence bands, based on Newey-West standard errors.

Figure B.2: Counterfactual of Non-Responsive Monetary Policy



Notes: Based on estimates of β^h and γ^h in (5.1). These estimates reflect the effect of military spending news amounting to 1% of potential GDP. We show the average response (β^h) and a counterfactual ($\beta^h + \gamma^h \overline{Hawk}^{\,h}$) where we minimize the FFR response over the first three years ($h = 0, .., 12$), by setting an appropriate value of $\overline{Hawk}^{\,h}$. The confidence bands are at the 68 and 95 percent level based on Newey-West standard errors.