

Government Spending between Active and Passive Monetary Policy*

Sebastian Laumer[†] Collin Philipps[‡]

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Abstract

Conventional wisdom suggests that the government spending multiplier is larger when the central bank raises nominal interest rates less than one for one to inflation. However, models supporting this consensus estimate multipliers while holding the monetary policy rule constant after a government spending shock. We show that the multiplier does not depend on monetary policy. We find instead that the monetary policy rule itself changes after a government spending shock. The monetary policy rule evolves quickly to reach a similar regime regardless of its initial condition. This rapid change of monetary policy in response to the economic condition after the government spending shock leaves the multiplier almost completely unaffected by the initial monetary policy regime. An exception to this characterization of monetary policy occurs when nominal interest rates are stuck at zero. We analyze the multiplier at the zero lower bound, and find that the multiplier exceeds one.

Keywords: Fiscal Multiplier, Monetary Policy, Zero-Lower Bound, Nonlinear SVARs

JEL Codes: C32, E32, E62

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[†]University of Illinois at Urbana-Champaign, David Kinley Hall, 1407 W Gregory Dr, Urbana, IL 61801, laumer2@illinois.edu

[‡]University of Illinois at Urbana-Champaign, David Kinley Hall, 1407 W Gregory Dr, Urbana, IL 61801, cphili6@illinois.edu

1 Introduction

Does the government spending multiplier depend on monetary policy? Conventional wisdom suggests that the multiplier is larger when monetary policy is passive.^{1,2} We show that this consensus misleads. Models supporting these predictions estimate multipliers while keeping the monetary policy rule constant after a government spending increase. The shortcoming of that approach is it fails to consider how the central bank adjusts its policy regime in response to the economic conditions after the government raises its spending. We demonstrate that the monetary policy rule itself changes quickly after a government spending intervention, and that it reaches a similar regime regardless of its initial condition. This rapid change of monetary policy challenges the widely accepted link between the multiplier and monetary policy.

This paper proposes a new methodology to analyze this relationship empirically without constraining monetary policy after a government spending intervention. We estimate a Taylor rule with time-varying coefficients, and use its sequence of inflation parameters to inform the monetary policy regimes in a flexible, nonlinear, structural vector autoregression (SVAR) model. We show that the monetary policy regime varies substantially over time, and that the central bank changes its policy regime in response to economic conditions during our sample period. The central bank becomes more active (“hawkish”) when inflation is high, and more passive (“dovish”) during recessions. This leads us to allow the central bank to update its policy regime in response to future economic conditions when we estimate the dynamic effects of a government spending shock. We find that the central bank responds quickly after the shock, and that it transitions rapidly to an active regime even if the initial regime had not previously been active.

The response of the monetary policy regime has a fundamental impact on the multiplier. Once we account for the regime’s reaction to the government spending shock, we find little evidence that monetary policy affects the multiplier. By contrast, when we counterfactually keep the monetary policy regime constant after the shock, our results match those of the theorists. This leads us to conclude that the conventional wisdom is primarily driven by the constant-regime assumption, which itself is difficult to reconcile with the data. In fact, the multiplier-monetary policy relationship vanishes once this assumption is relaxed. Our

¹When monetary policy is passive, the central bank raises nominal interest rates less than one for one to inflation, and the real interest rate decreases. The lower real interest rate leads households to increase consumption. Consequently, output increases more than government spending and the multiplier is predicted to be larger than one. In contrast, when monetary policy is active, the central bank responds more than one for one to inflation, and the multiplier is predicted to be smaller than one.

²See e.g., Kim [2003], Canova and Pappa [2011], Woodford [2011], Christiano et al. [2011], Davig and Leeper [2011], Zubairy [2014], Dupor and Li [2015], Leeper et al. [2017] and Cloyne et al. [2020].

analysis suggests that to boost the impact of government spending on the economy, the central bank would need to accommodate inflation for a long period of time. However, this tactic also creates a dilemma, both because it would conflict with the central bank’s main mandate of maintaining price stability, and because it would violate the *Taylor principle*.

The effect of government spending on the economy is a longlasting research topic in macroeconomics. The core question is whether the government spending multiplier, which measures the change in output in response to a \$1 increase in government spending, exceeds one. Despite extensive research, empirical and theoretical studies reach different conclusions.³ A major point of agreement is that monetary policy is a key determinant of the multiplier. Even before the 2008 financial crisis, interactions between government spending and monetary policy were a major policy consideration. Today, governments and central banks are again working closely to limit the economic consequences of the COVID-19 pandemic.

Our work analyzing the relationship between the multiplier and monetary policy makes several contributions. First, we augment the smooth-transition VAR (ST-VAR) model, popularized by Auerbach and Gorodnichenko [2012], with the Taylor rule. Our approach models the path of monetary policy as an evolving mix of two extreme monetary policy regimes: one when monetary policy is extremely active, and another when monetary policy is extremely passive. Following, Cogley and Sargent [2001, 2005] and Primiceri [2005], we estimate a Taylor rule with time-varying parameters, and use the corresponding series of inflation parameters to inform the evolution of monetary policy in the ST-VAR model. We find that the monetary policy regime evolves continuously over time. That is, monetary policy is not exclusively active or passive as theory assumes. Instead, it varies in more nuanced ways. For instance, during the 1960s and 1970s, the monetary policy regime was neither strongly active nor strongly passive at various times. Crucially, we also uncover that the central bank changes its policy regime in response to economic conditions. For example, in August 1979, the U.S. Federal Reserve raised nominal interest rates aggressively in response to high inflation after Paul Volcker became its chairman, yet it lowered nominal interest rates in response to the recessions of 1957-58, 1990-91 and 2000-01.

Second, we incorporate these features about monetary policy when we estimate the dynamic effects of a government spending shock. When monetary policy itself changes, and one wants to understand the impact of a government spending shock, one must consider the initial condition of the shock *and* the consequences for the transitions. These transitions

³For theoretical studies, see for example, Aiyagari et al. [1992], Baxter and King [1993], Linnemann and Schabert [2006], Galí et al. [2007], Zubairy [2014] or Leeper et al. [2017] and the citations therein. For empirical papers, see for instance, Ramey and Shapiro [1998], Blanchard and Perotti [2002], Fatás and Mihov [2001], Mountford and Uhlig [2009], Ramey [2011], Ben Zeev and Pappa [2017] or Ellahie and Ricco [2017].

reflect both the direct effect of a shock on the variables, and the indirect effect via the future evolution of the monetary policy regime in response to economic conditions. We start by dividing the initial policy regimes into quintiles, and then allowing the central bank to adjust its policy regime after the government spending shock. Finally, we compare the multiplier estimates across the initial quintiles.

This exercise reveals that the central bank responds quickly after a government spending shock, and that it transitions rapidly to adopt an active policy even if the initial policy had not previously been “very active.” As a result, shortly after the shock, and regardless of its initial condition, the central bank reacts actively to inflation. More importantly, the response of the monetary policy regime has vital consequences for the multiplier. Once we account for the regime’s reaction to the government spending shock, we find little evidence for a relationship between the multiplier and monetary policy. We argue that the dependence of the multiplier on monetary policy vanishes because of a modeling flaw: the literature ignores the response of the monetary policy regime, and, instead, keeps it constant after the shock. This theoretical restriction on the policy regime is not a feature of the data, giving rise to unwarranted conclusions about how the multiplier depends on monetary policy.

Third, we identify a government spending shock using sign restrictions on impulse response functions. We define a government spending shock as a shock that drives up output, inflation, government spending, government tax revenue, and government debt. These restrictions represent joint predictions of theoretical models, and hence, are mostly uncontroversial. The rest of the empirical literature relies on zero restrictions related to the standard Cholesky approach, with scant theoretical foundations [Uhlig, 2005]. Regardless of the monetary policy regime at the time of a shock, our posterior median estimate for the multiplier is almost 5 in the short run, notably above most estimates in the literature. Our multiplier estimates then decline to about 1 after five years.

Next, we conduct counterfactual exercises. We first analyze what happens to the multiplier if the central bank were to keep its monetary policy regime temporarily constant after the shock. We repeat our main exercise but keep the monetary policy regime constant for one, two, and five years. In another exercise, we replicate the framework that underlies the conventional wisdom with our empirical model. To do this, we distinguish between the two most extreme monetary policy regimes, and we fix the regimes for the entire horizon after the shock. These exercises show that the multiplier may depend on monetary policy, but *only* if the central bank keeps its initial policy regime constant for a surprisingly long period of time after the shock (over two years according to our estimates). Because the central bank usually responds quickly after a government spending shock, we conclude that the key driver of this conventional wisdom is the constant-regime assumption — not the data.

The only exception to the active/passive characterization of monetary policy occurs when nominal interest rates are stuck at zero in response to a severe economic downturn, such as the 2008 financial crisis. This leads us to analyze the multiplier at the zero lower bound separately. We provide evidence that the zero lower bound represents a third extreme monetary policy regime that we cannot capture with our two-regime model. Additionally, between 2008Q4 and 2015Q4, the Fed held nominal interest rates at zero, and it did not change its policy regime despite unprecedented fiscal policy interventions such as the American Recovery and Reinvestment Act (ARRA) of 2009. This behavior differs substantially from previous periods.

Accordingly, we analyze the multiplier at the zero lower bound by holding the monetary policy regime constant after the government spending intervention. Then, we compare the multiplier estimates across different strategies for identifying government spending shocks. When we employ the standard Cholesky approach, our multiplier estimates are smaller than one, a finding which is in line with that of Ramey and Zubairy [2018]. On the contrary, when we apply our sign-restriction method, the multiplier estimates are considerably larger. The posterior median estimate is 4.5 on impact. Five years later, it falls to 3. We argue that the timing restrictions related to the Cholesky approach, as used in Ramey and Zubairy [2018], may be violated, especially during this era. The Cholesky approach assumes that governments change their spending in response to shocks to the business cycle only with a delay. Contrary to this premise, however, recent events such as the Coronavirus Aid, Relief and Economic Security (CARES) Act of 2020 reveal that governments can and do react quite quickly during crises, for example, when central banks cut nominal interest rates to zero in response to a deep economic recession. Our results suggest that the multiplier at the modern zero-lower bound in the United States is larger than has previously been found, and that recent fiscal stimulus packages, such as the ARRA and the CARES Act, might also have larger economic effects than the literature has formerly concluded.

This paper proceeds as follows: Section 2 introduces the methodology. Section 3 describes the evolution of monetary policy in the United States. Section 4 presents our main results. Section 5 shows our counterfactual analysis. Section 6 includes our zero-lower bound analysis. Section 7 concludes.

2 Model

2.1 Model

This paper studies how the government spending multiplier depends on the responsiveness of monetary policy to inflation. To allow for policy-dependent multipliers, we use the following smooth transition VAR (ST-VAR) model:

$$X_t = (1 - G(z_{t-1}))\Pi_{AM}X_{t-1} + G(z_{t-1})\Pi_{PM}X_{t-1} + u_t \quad (1)$$

$$u_t \sim N(0, \Omega_t) \quad (2)$$

$$\Omega_t = (1 - G(z_{t-1}))\Omega_{AM} + G(z_{t-1})\Omega_{PM} \quad (3)$$

$$G(z_t) = \frac{1}{1 + \exp(\gamma(z_t - c))} \quad (4)$$

where X_t is a vector of endogenous variables that represents the economy. X_t consists of real government spending, real tax receipts, real GDP, Ramey’s news shocks, GDP inflation, the federal funds rate, and real government debt.⁴ We include Ramey’s news shock to account for the issue of fiscal foresight.^{5,6}

Equation (1) says that the economy X_t evolves as a convex combination of two ideal monetary policy regimes: “purely active” monetary policy (AM) and “purely passive” monetary policy (PM) regimes. The transition function $G(\cdot)$ governs the transition between the AM and PM regimes. The state-determining variable z_t characterizes the underlying state of the economy. To describe the evolution of monetary policy over time, we estimate a Taylor rule with time-varying coefficients, and use the estimated sequence of the inflation parameter as z_t . This choice allows us to distinguish between different monetary policy regimes for each quarter of our sample period. We provide more details about our choice of z_t in Section 2.2. In Equation (1), u_t is a vector of reduced-form residuals. We assume that the residuals are normally distributed with zero mean and a time-varying covariance matrix Ω_t . The covariance Ω_t also evolves as a convex combination of the pure regime covariance matrices, Ω_{AM} and Ω_{PM} .

⁴Real government spending, real tax receipts, real GDP, and real government debt are measured in growth rates. The remaining variables are in levels.

⁵Fiscal foresight arises because changes in fiscal policy are often announced before they are implemented. Hence, agents can react before the policy change occurs. This creates a misalignment between the information sets of economic agents and the econometrician. As a result, the VAR cannot consistently estimate impulse response functions [Leeper et al., 2013, Forni and Gambetti, 2014]. The standard approach to deal with this issue is to include Ramey’s news shocks to control for agents’ expectations [Ramey, 2011]. We follow this method.

⁶In Appendix C, we conduct a robustness check replacing Ramey’s news shock by another measure for agents’ expectations.

Our ST-VAR model approximates the variation in monetary policy over time as a smoothly evolving convex combination of the two pure *AM* and *PM* regimes. In each period, $G(z_{t-1})$ represents the relative weight on the *PM* regime and is given by Equation (4). We choose $G(z_{t-1})$ to be a logistic sigmoid function, as it is the standard choice in the related literature (see, e.g., Auerbach and Gorodnichenko [2012], Caggiano et al. [2015] or Ramey and Zubairy [2018]). Equation (4) says that $G(z_{t-1})$ is continuous, monotonically decreasing, and bounded between zero and one. The transition function $G(z_{t-1})$ depends on the transition parameter γ , the state variable z_t , and on the threshold parameter c . As the variable z_t moves from below c to above c , the value of $G(z_{t-1})$ decreases, and the model puts relatively more weight on the *AM* regime. The rate of this transition between regimes is determined by γ . If $\gamma \approx 0$, then the ST-VAR model collapses to a linear VAR model, and no transition occurs. Conversely, if $\gamma \approx \infty$, then the model jumps directly from the *PM* to *AM* regime as soon as z_t surpasses c . Thus, the ST-VAR model nests a threshold-VAR model, which also nests Markov-type models. For any value of γ between those two extremes, the transition between the two extreme regimes is smooth.

The literature suggests that monetary policy activism has varied substantially over time, but that the nature of this variation is unclear. To allow the data to speak freely, we use fully Bayesian methods to estimate the joint posterior distribution of the model. This approach lets the data inform the model structure – particularly for the state-transitioning parameters. This would not be possible when parameters are calibrated. First, we define prior distributions for all model parameters, i.e. $\Pi_{AM}, \Pi_{PM}, \Omega_{AM}, \Omega_{PM}, \gamma$ and c . Second, because the joint posterior distribution is analytically untractable, we employ the multi-move Gibbs sampler proposed by Gefang and Strachan [2010] and Galvão and Owyang [2018] to let the data update our prior beliefs. The Gibbs sampler partitions the vector of model parameters into different groups. Then, the Gibbs sampler generates draws for each group separately from the corresponding marginal posterior distributions, conditional on the remaining parameters. This procedure simplifies the drawing process because the marginal posterior distributions are often known and easy to sample from. Finally, the draws from the marginal posteriors approximate the joint posterior of the model.⁷

In our context, it is crucial that Bayesian methods allow the data to be informative about the transition parameters γ and the threshold parameter c , which in turn inform our beliefs about the evolution of monetary policy. Conditional on the marginal posterior distributions of γ and c , we can compute the posterior distribution of $G(z_{t-1})$ as a function of z_t . Our estimate of $G(z_{t-1})$ captures how monetary policy activism has evolved over time. This exercise reveals that the monetary policy regime behaves differently from how theory models

⁷See Appendix A for a detailed description of our Bayesian Sampler.

it. Figure 1 illustrates that the monetary policy regime evolves continuously over time and changes in response to inflation and recessions.⁸

2.2 State-Determining variable z_t

The key ingredient of the smooth-transition VAR model, equations (1) - (4), is the state-determining variable z_t . The choice of z_t characterizes the underlying nonlinearity of the economy. Therefore, z_t should embody economically meaningful and empirically informative attributes of monetary policymakers' behavior. In models similar to ours, the standard approach is to use an observable variable as z_t . For example, to replicate the business Cycle of the U.S. economy, Auerbach and Gorodnichenko [2012] use the growth rate of real GDP, and Ramey and Zubairy [2018] use the unemployment rate.

Monetary policy is not determined by any single variable, and there is no unique measure of monetary policy activism. However, when theorists make predictions about how the effect of government spending depends on monetary policy, they typically refer to the central bank's responsiveness to inflation as the determinant that distinguishes between the monetary policy regimes [Leeper et al., 2017]. When monetary policy is active, the central bank raises nominal interest rates more than one for one to inflation. As a result, the real interest rate increases and induces households to decrease consumption. Consequently, output increases, but it increases by less than government spending, so the government spending multiplier is less than one. When, instead, monetary policy is passive, the central bank reacts only weakly to inflation, and the real interest rate decreases. As a result, households increase consumption, so the government spending multiplier exceeds one. We believe that this measure of policy *responsiveness* likely influences the dynamics of the real economy.

The central bank's responsiveness to changes in inflation is a major theme in both applied and theoretical literatures. In the wake of Taylor [1993], theorists have typically used the Taylor rule to model monetary policy. According to the Taylor rule, the central bank sets its policy instrument in response to inflation and output. Because ample evidence exists that monetary policy has changed over time,⁹ we estimate the following ex-post Taylor rule with time-varying parameters:

$$i_t = c_t + \phi_{\pi,t}\pi_t + \phi_{y,t}y_t + v_t, v_t \sim N(0, \sigma_t^2) \quad (5)$$

where i_t is the federal funds rate, π_t inflation and y_t is the growth rate of real GDP. c_t is a time-

⁸We describe the evolution of monetary policy in more detail in Section 3.

⁹See e.g., Taylor [1999], Clarida et al. [2000], Cogley and Sargent [2005], Primiceri [2005], Boivin [2006], Wolters [2012] or Carvalho et al. [2019].

varying constant, and $\phi_{\pi,t}$ and $\phi_{y,t}$ capture the central bank's time-varying responsiveness to inflation and output, respectively. σ_t^2 corresponds to the variance of the residuals that can also vary over time. To allow for permanent and temporary changes in the model coefficients, we assume that $[c_t, \phi_{\pi,t}, \phi_{y,t}]$ and $\log(\sigma_t)$ both follow a Random Walk. We then exploit $\phi_{\pi,t}$ to distinguish between different monetary policy regimes in each quarter of our sample period: when $\phi_{\pi,t}$ exceeds one, then the central bank raises nominal interest rates more than one for one to inflation and the monetary policy regime is *active* in that period. By contrast, if $\phi_{\pi,t}$ is smaller than one, then the central bank responds less than one for one to inflation and the monetary policy regime in that period is *passive*.

The model in (5) can be written with a state-space representation

$$i_t = \mathbf{z}'_t \phi_t + e_t, e_t \sim N(0, \sigma_t^2) \quad (6)$$

$$\phi_t = \phi_{t-1} + v_t, v_t \sim N(0, q) \quad (7)$$

$$\log(\sigma_t) = \log(\sigma_{t-1}) + \eta_t, \eta_t \sim N(0, w) \quad (8)$$

where $\mathbf{z}_t = [1, \pi_t, y_t]$ and $\phi_t = [c_t, \phi_{\pi,t}, \phi_{y,t}]$. (6) and (7) form a linear-Gaussian state-space model, which implies that standard Kalman-filtering techniques can be used to estimate ϕ_t . In contrast, (6) and (8) build a Gaussian, but nonlinear state-space model. In this case, standard Kalman filtering is not directly applicable. However, Kim et al. [1998] and Primiceri [2005] show how the model in (6) and (8) can be transformed into a linear-Gaussian state-space model so that standard Kalman-filtering techniques can be used to estimate σ_t . We explain the estimation procedure in greater detail in Appendix A.

We employ Bayesian methods to estimate (6) - (8) and then use the median estimate of $\phi_{\pi,t}$ to determine the monetary policy regime for each quarter of our sample period. Therefore, our model is not the standard smooth-transition VAR model as in Auerbach and Gorodnichenko [2012], but involves a prior step in which we estimate the state variable as a parameter with time-varying coefficients. We refer to our model as smooth-transition VAR model with time-varying parameters (TVP-ST-VAR model). Our model is preferable to the standard ST-VAR model in situations in which the underlying state of the economy is not well-captured by a single observable variable but can be characterized via a parameter that varies over time.

The estimation of the Taylor rule is an interesting topic in the literature. In Appendix B, we review the literature, and conduct a battery of robustness checks. We now describe the details of our baseline specification of (5).

First, we use current inflation and the current growth rate of real GDP as regressors. These choices raise endogeneity concerns [Clarida et al., 2000]. de Vries and Li [2013] argue

that usual instruments (i.e., lagged regressors, real-time data or the instrument set of Clarida et al. [2000]) are not fully exogenous if the residuals of the Taylor rule are serially correlated; at the same time, however, they insist that the resulting bias will be small. In addition, Carvalho et al. [2019] estimate the Taylor rule before and after 1980 using ordinary least squares (OLS) and different instruments and uncover comparable estimates. We find that the residuals of our baseline specification are serially correlated (see Appendix B). Moreover, we find that the estimate of $\phi_{\pi,t}$ of our baseline specification is comparable to those of specifications with lagged regressors. Fortunately, each method captures a similar pattern of variation over time. And because our transition function contains an estimated γ, c , the choice of the state variable is only unique up to an affine transformation.

Second, Orphanides [2001, 2002, 2004] and Boivin [2006] advocate using real-time data rather than revised data because revised data may contain information that was not available when policy makers made their monetary policy decisions. Despite their claim, we use revised data in our baseline specification of (5). We are more interested in the interaction between policy and the economy, and less interested in how policy decisions were made. For example, a central bank’s decision can turn out differently than had been intended, and this can only be seen using revised data. In contrast, real-time data may be more useful in capturing the central bank’s intentions. In Appendix B, we show that the estimates of $\phi_{\pi,t}$ using both revised and real-time data are very similar.

Third, we use the growth rate of real GDP to measure output rather than the output gap as suggested by Taylor [1993]. This is reasonably common in the literature because the natural rate of output is not an observable variable, and estimates of it are believed to be imprecise. In addition, Orphanides and Williams [2002] argues that interest rate rules based on output growth are preferable to those that are based on the output gap of models in which the natural rate of unemployment is uncertain. Walsh [2003] shows that loss functions based on output growth rather than the output gap result in lower losses. Finally, Coibion and Gorodnichenko [2011] illustrate that the Taylor principle breaks down under positive trend inflation, but that it can be restored if the central bank responds to output *growth* rather than to the output gap. As we detail in Section 2.4, we allow the central bank to adjust its policy regime after the government spending shock. This idea requires us to update the inflation parameter in our forecasts of the real economy. Hence, we must also include all variables of the Taylor rule in the VAR part of our model. Because the impulse response of output is a key ingredient in the formula for the multiplier, we measure output using the growth rate of real GDP instead of using the output gap.

Finally, the Taylor rule in (5) can be contrasted with our model of the economy in (1), which also contains an equation for the nominal interest rate. Though the simple Taylor rule

is nested in the larger model, the two equations do not embody the same information about the economy. While even simple versions of the Taylor rule have been shown to successfully approximate the central bank’s policy instrument, the nominal interest rate equation of (1) includes variables that are typically not included in the Taylor rule (e.g., fiscal variables, and several lags of all endogenous variables). In addition, the current response parameters in Equation (1) are included in the impact matrix, which is not statistically identified. Because the Taylor rule is the most widely accepted way to model monetary policy activism, we choose $\phi_{\pi,t}$ from equation (5) as the state variable that distinguishes between different monetary policy regimes in our analysis. This choice ensures that we capture a recognizable component of “true” monetary policy behavior, and that we avoid imposing any additional restrictions on the overall economy in Equation (1).

2.3 Generalized Impulse Response Functions

In nonlinear models such as ours impulse responses depend on the shock’s sign, size and the timing. In addition, the state of the economy can respond to economic conditions – both during the sample period and after shocks. If one wants to understand how the impact of a government spending shock depends on monetary policy, and how monetary policy itself changes, then one must consider the initial state of the economy and the dynamics that arise from both the direct effect of a shock on the variables and its *indirect* effects via the future evolution of the monetary policy rule after the shock.

To incorporate these features, we follow Koop et al. [1996], and we employ generalized impulse response functions. Generalized impulse response functions are defined as the expected difference between two simulated paths of the economy. Formally, they can be written as

$$GIRF(h) = E[(1 - G(z_{t+h-1}))\Pi_A\hat{X}_{t+h-1}^\epsilon + G(z_{t+h-1})\Pi_P\hat{X}_{t+h-1}^\epsilon + \epsilon_{t+h}] \\ - E[(1 - G(z_{t+h-1}))\Pi_A\hat{X}_{t+h-1}^u + G(z_{t+h-1})\Pi_P\hat{X}_{t+h-1}^u + u_{t+h}]. \quad (9)$$

The first part of (9) represents the simulated path of the economy hit by a government spending shock, \hat{X}_{t+h}^ϵ . The second part corresponds to the simulated path when the economy is *not* hit by the government spending shock, \hat{X}_{t+h}^u .

The generalized impulse response functions require initial conditions from a starting period. Given a particular starting period, we use (1) to roll the model forward in both simulations. Thus, we can estimate the effects of a government spending shock for each monetary policy regime by choosing starting conditions that correspond to a particular pol-

icy. Later, we divide the monetary policy regimes into quintiles to distinguish between different monetary policy regimes (e.g., between “very active,” “weakly active,” “neutral,” “weakly passive,” and “very passive” regimes). Then, we randomly draw the initial conditions from these quintiles, and present the average difference between two simulations as in (9) for each quintile.

2.4 Updating Rules for Monetary Policy

In nonlinear models such as ours, the state of the economy can respond to its economic environment, such as in response to shocks. In Section 3, we see that the central bank changes its monetary policy regime frequently in response to inflation and recessions. The generalized impulse response functions also allow for scenarios in which the underlying state of the economy changes after the shock. Based on the forecasts for the economies with and without a government spending shock, X_{t+h}^e and X_{t+h}^u , we can also forecast the corresponding values of z_{t+h} and then of $G(z_{t+h})$. The new values of $G(z_{t+h})$ and $1 - G(z_{t+h})$ represent the weights assigned toward to purely active and purely passive monetary policy h periods after the shock. The updated weights then enter the forecasts of the economies with and without the government spending shock in the next periods, X_{t+h+1}^e and X_{t+h+1}^u . Hence, the monetary policy regime can change after the government spending shock, and can further affect how the government spending shock disseminates through the economy.

To implement this feature, we run the Kalman filter (as described in Section 2.2) based on the forecasted values of the federal funds rate, inflation rate, and output growth. Each time we predict the economy one step ahead, the Kalman filter obtains a new $\phi_{\pi,t+h}$ and $\phi_{y,t+h}$, which represent the central bank’s responsiveness to inflation and output growth, respectively, for that period. At each step, we use that estimate of $\phi_{\pi,t+h}$ as z_{t+h} , and we plug it into the transition function to update the weights of the two monetary policy regimes. This method requires the Taylor rule to use the same information set as the rest of our model. We believe that our model is the most intuitive approach. (We discuss other possibilities in Appendix C.¹⁰) These updating rules can only reflect average behavior according to the historical data; the Federal Reserve relies on its own discretion when updating its policy.

¹⁰For example, we can also run a regression of the forecasted values of the federal funds rate on the forecasted values on inflation and output growth, and use the point estimate of the inflation parameter, $\hat{\phi}_{\pi,t+h}$, as z_{t+h} . Additionally, we can update $\phi_{\pi,t+h}$ via an AR(1) process in which we predefine the persistence parameter.

2.5 Identification

Identification of structural shocks in our nonlinear model is similar to identification in linear models. The reduced-form errors, u_t , do not have an economic interpretation. Macroeconomic theory assumes that the reduced-form errors are linear combinations of structural shocks ϵ_t . Formally,

$$u_t = A\epsilon_t, \tag{10}$$

where

$$A = Chol(\Omega_t)Q. \tag{11}$$

$Chol(\Omega_t)$ is the Cholesky decomposition of the covariance matrix Ω_t , and Q is an orthogonal matrix, $QQ' = I$. The matrix A is not identified statistically. To identify the structural shocks, we must impose economically meaningful restrictions on A .

We identify government spending shocks using sign restrictions on impulse response functions. This strategy has been used in the literature by Mountford and Uhlig [2009], Canova and Pappa [2011] and Laumer [2020]. We follow Laumer [2020], and define a government spending shock as a shock that drives up output, inflation, government spending, government tax revenue, and government debt. This set of restrictions is consistent with a large set of neoclassical and new Keynesian models (e.g., Woodford [2011]). In this study, we assume that this set of restrictions holds regardless of the monetary policy regime.

The sign-restriction strategy has major advantages over other methodologies. First, the imposed restrictions are derived from theoretical models. Our restrictions represent joint predictions of theoretical models, and are therefore mostly uncontroversial. Furthermore, sign restrictions impose only weak restrictions on the behavior of the economy, as we only restrict the signs of certain impulse responses. Finally, sign restrictions allow all variables to contemporaneously react to all shocks while other strategies (e.g., the recursive approach) impose specific timing restrictions that govern which variables can react to the shocks on impact. Sign restrictions in linear VAR models have been widely applied and are well documented (e.g., see Baumeister and Hamilton [2018] for an overview). We follow the approach of Laumer and Philipps [2020], who implement sign restrictions on generalized impulse response functions.¹¹

¹¹Details are provided in Appendix A.

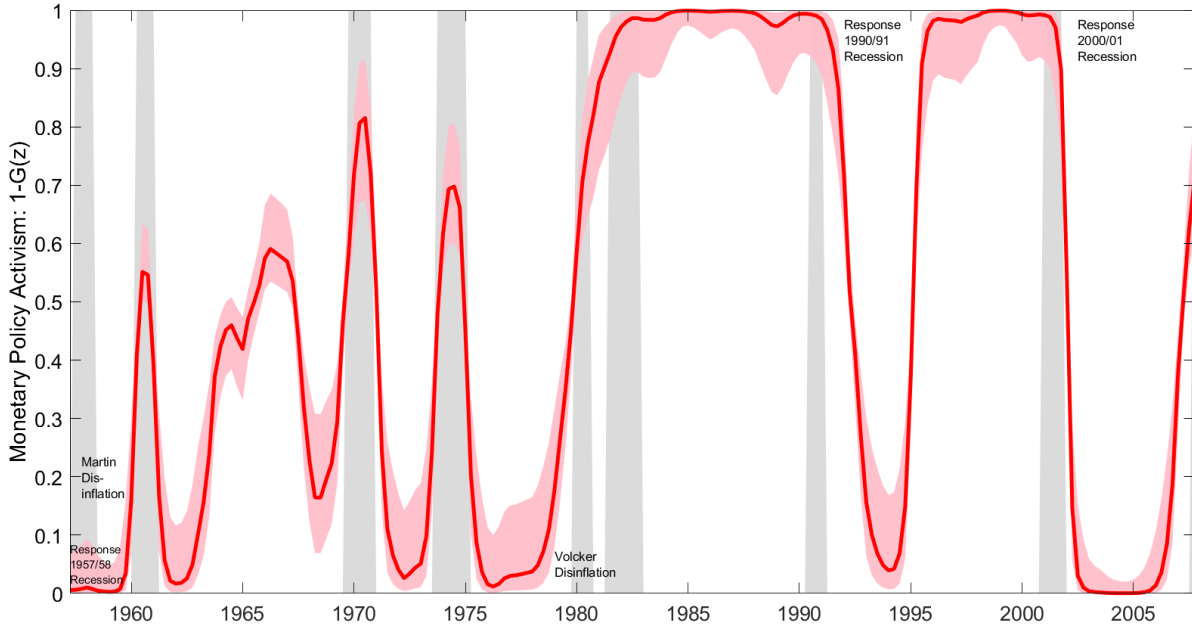
3 History of Monetary Policy

In this section, we describe how monetary policy has evolved over our sample period. We are using quarterly data for the U.S. economy from 1954Q3 to 2007Q4. In Section 6, we extend the sample period to 2015Q4 to include the zero lower bound epoch. Figure 1 plots the median estimate of $1 - G(z_{t-1})$, the weight assigned to the purely active monetary policy regime in Equation (1), along with the 68 percent credible bands. We interpret $1 - G(z_{t-1})$ as a measure of the central bank’s activism because larger values of $1 - G(z_{t-1})$ occur when the central bank responds more strongly to inflation. Low values of $1 - G(z_{t-1})$ indicate a more “passive” monetary policy regime.

Figure 1 reveals that the monetary policy regime varies substantially over time. Our estimate of $1 - G(z_{t-1})$ suggests that monetary policy was very passive during the second half of the 1950s, while the model puts some weight on the purely active monetary policy regime during the 1960s. In the 1970s, the monetary policy regime was very passive, only interrupted by the summer of 1970 and the winter of 1974/75. Consistent with the existing narrative in the literature (e.g., Romer and Romer [2004]), our model indicates that monetary policy changed dramatically after Paul Volcker had become the chairman of the Federal Reserve. Between 1979 and 1982, monetary policy transitioned from a very passive regime to a very active regime, and remained very active throughout the 1980s. During the first half of the 1990s, our model assigns a low weight to the active regime, while it assigns a relatively high weight to the active regime during the second half. Since 2000, monetary policy has been mostly passive, save for the very end of our sample period.

Our results regarding $1 - G(z_{t-1})$ lend support to existing narrative evidence given by Romer and Romer [2004]. Romer and Romer [2004] describe how monetary policy was very passive during the late 1950s before the Federal Reserve increased nominal interest rates in response to high inflation in 1959 (Martin Disinflation). Moreover, Romer and Romer [2004] illustrate that the Fed ran a very passive monetary policy throughout most of the 1970s, only interrupted in 1970 and in the winter of 1974/75. Subsequently, the Fed raised nominal interest rates in response to high inflation. This occurred around the time of Paul Volcker’s appointment in August 1979 (Volcker Disinflation). In addition, Romer and Romer [2004] characterize the chairmanships of Paul Volcker (1979-1987) and Alan Greenspan (1987-2006) as periods in which the Fed responded actively to inflation. These periods were only interrupted during the first half of the 1990s and 2000s, when the Fed lowered nominal interest rates in response to the 1990-91 and 2000-01 recessions. Finally, Taylor [2007] argues that the Fed deviated substantially from the Taylor principle between 2002 and 2005. Our estimates in Figure 1 support this argument as well. Between 2002 and

Figure 1: Evolution of Monetary Policy between 1954 and 2008



Note: The figure shows the pointwise-posterior median estimate of $1 - G(z)$, along with the 68 percent credible bands. $1 - G(z)$ can be interpreted as a measure for monetary policy activism. Grey bars represent recessions as defined by the National Bureau of Economic Research (NBER). The figure provides evidence that monetary policy behaves differently from how theory models it. Monetary policy evolves continuously over time, changes smoothly, and responds several times to both inflation and recessions.

2005, $1 - G(z_{t-1})$ reaches its lowest value of the entire sample period. Appendix D contains a more detailed timeline of monetary policy in the United States.

The posterior of $1 - G(z_t)$, shown over time in Figure 1, also sheds some light on the theoretical frameworks in the literature. These frameworks typically only distinguish between active and passive monetary policy, and assume that monetary policy rules do not change in response to economic conditions such as inflation or recessions – or to government spending shocks (e.g., in Kim [2003], Zubairy [2014] or Leeper et al. [2017]). Others use the Markov-switching approach to model policy changes (e.g., Davig and Leeper [2011]). That method implies that the central bank jumps randomly from one regime directly to another.

Figure 1 provides evidence that the monetary policy regime evolves differently from these models. First, monetary policy is not well described by the two extreme regimes, but rather by a process that evolves continuously over time. For example, during the second half of the 1950s and in the late 1960s, the monetary policy regime reflects two different degrees of passiveness. Similarly, in 1970 and throughout the 1980s, the policy regime exhibits two different degrees of activeness. This observation implies that even within the active and the

passive regimes themselves, differences occur.

Second, Figure 1 suggests that when monetary policy changes, it changes smoothly rather than abruptly. For example, the change in the monetary policy regime around 1980 took three years. The data do not seem to prefer abrupt policy changes (e.g., an extremely large γ parameter, casting doubt on the Markov-switching approaches). This result supports Bianchi and Ilut [2017] and Chang and Kwak [2017], who reach similar conclusions that monetary policy evolves smoothly.

Finally, our estimates of monetary policy and the narrative evidence given by Romer and Romer [2004] demonstrate that the central bank has changed its policies several times to economic conditions during our sample period. For example, the Fed increased nominal interest rates aggressively (and became extremely active) in response to high inflation in 1959 and 1979-1982. Similarly, the Fed lowered nominal interest rates (and became more passive) in response to the recessions of 1957-58, 1990-91 and 2000-01. These observations suggest that impulse response functions that fix the Taylor rule after the government spending shock might be misplaced, rendering their conclusions suspect.

Taken together, these observations guide our approach to estimate the government spending multiplier conditional on the monetary policy regime in what follows. Most importantly, we observe that the central bank *does* change its policy in response to economic conditions during our sample period. Next, we ask whether the evolution of monetary policy can be affected by government spending shocks. Our main model allows for the possibility that the central bank adjusts its policy regime in response to future economic conditions after the shock (e.g., the central bank can switch to a more active policy when inflation grows large after a government spending increase). In the next section, we explore how the assumption of a constant policy rule may influence how our modeled economy evolves after government spending shocks.

4 Results

This section presents our main results. We estimate our model using quarterly data for the U.S. economy from 1954Q3 to 2007Q4. The sample excludes the period in which monetary policy in the United States and other countries was constrained by the zero lower bound. This cutoff is often applied in the literature to avoid contamination through the effects of the Great Recession and unconventional monetary policy (e.g., Leeper et al. [2017] or Arias et al. [2019]). More importantly, though, the Fed kept nominal interest rates at zero between 2008Q4 and 2015Q4. This says that despite massive fiscal policy interventions such as the American Recovery and Reinvestment Act (ARRA) or the Troubled Asset Relief Program

(TARP), the Fed did not respond to economic conditions during this period.¹²

We estimate our model with fully Bayesian methods that also include the marginal posterior distributions of the transition parameter γ and the threshold parameter c . We then employ generalized impulse response functions to estimate the dynamic effects of a government spending shock. The generalized impulse response functions require an initial condition that allows us to estimate the multiplier for each initial monetary policy regime of our sample period. For comparison, we divide the initial regimes into quintiles and compare the multipliers when monetary policy is initially “very active,” “weakly active,” “neutral,” “weakly passive,” or “very passive” according to the value of $G(z_{t-1})$ during the impact period. If the conventional wisdom is correct, we expect to see increasing multiplier estimates as we move smoothly from the most active quintile toward the most passive quintile. Finally, the generalized impulse response functions allow the central bank to change its policy regime in response to the government spending shock. For instance, the central bank can switch to a more active policy if inflation grows large after the shock.

To estimate the multiplier, we follow Auerbach and Gorodnichenko [2012], and use the formula for the *sum* multiplier

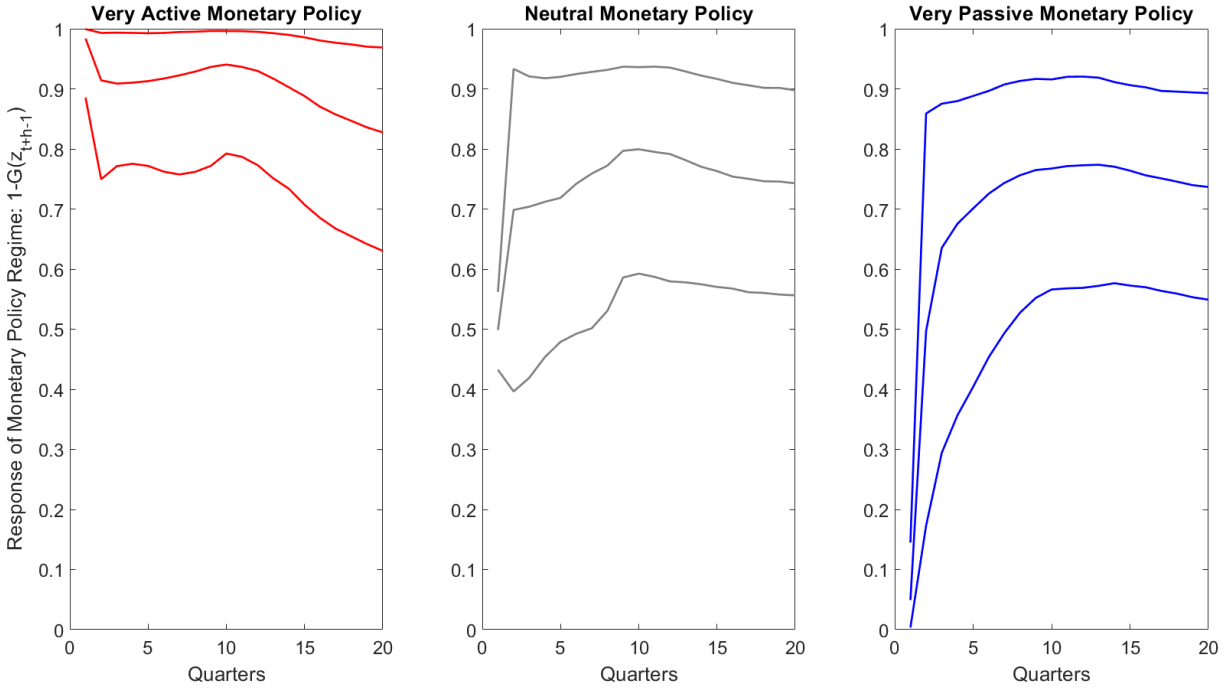
$$\Phi_h = \frac{\sum_h^{j=0} y_j}{\sum_h^{j=0} g_j} \times \frac{\bar{Y}}{\bar{G}}, \quad (12)$$

where y_j and g_j are the responses of output and government spending, respectively, in period j after the government spending shock. $\frac{\bar{Y}}{\bar{G}}$ is the sample mean of the output-to-government-spending ratio. Using the sum formula, we estimate the multiplier for different time horizons after the shock.

Figure 2 displays the response of the monetary policy regime after the government spending shock if the monetary policy regime is initially “very active,” “neutral,” and “very passive” using the evolution of $1 - G(z_{t-1})$. The figure reveals interesting results: when the monetary policy regime is initially “very active” (left subplot), the central bank does not adjust its responsiveness to inflation very much. The bank remains in the active sphere of the monetary policy spectrum for the entire horizon after the shock. In contrast, if the monetary policy regime is initially “neutral” or “very passive,” then the central bank responds quickly and transitions fast to a more active regime. Thus, shortly after the shock, the central bank – regardless of its initial regime – conducts policy in more or less the same way and responds actively to inflation. This result implies that the common practice of keeping the monetary policy regime constant after the shock is misplaced, *especially* for passive regimes. A natural question to ask is whether the response of the monetary policy regime affects the government

¹²We analyze the government spending multiplier at the zero lower bound in Section 6.

Figure 2: Response of the Monetary Policy Regime to a Government Spending Shock



Note: The figure shows the pointwise-posterior median evolution of $1 - G(z)$, along with the 68 percent credible bands, after a government spending shock when monetary policy is initially “very active,” “neutral,” or “very passive.” The figure provides evidence that the central bank responds quickly after the shock. Shortly after the shock and regardless of its initial condition, the central bank responds actively to inflation.

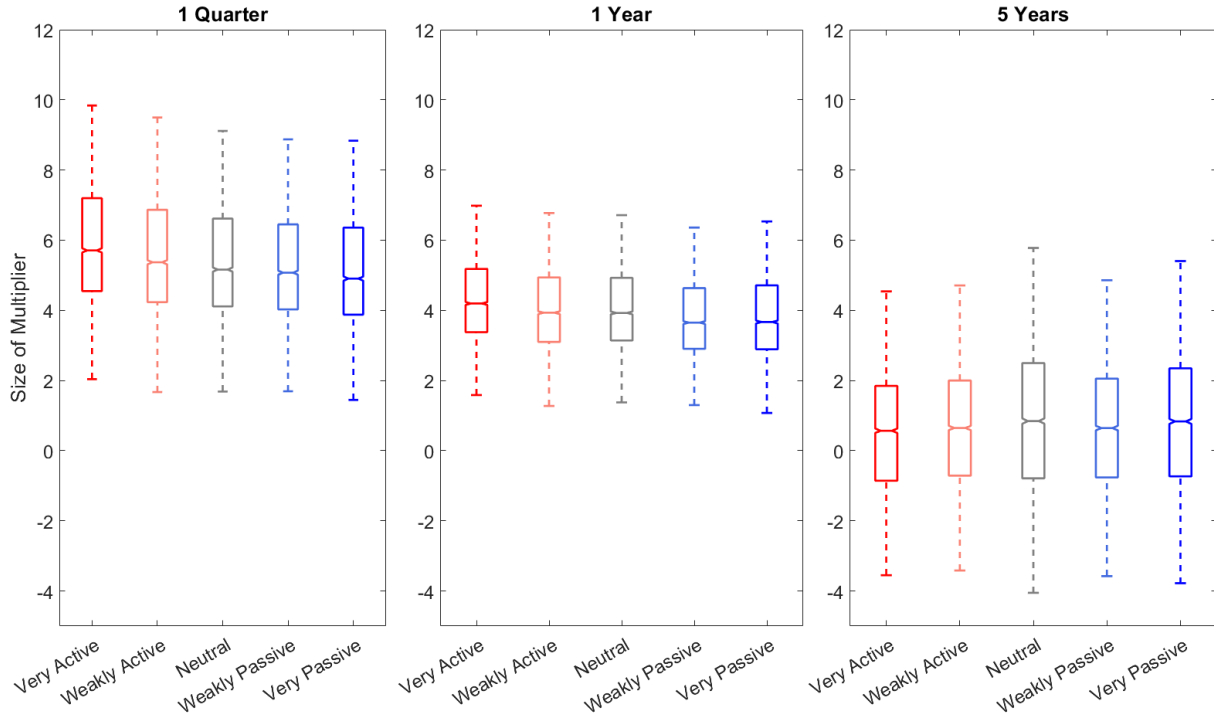
spending multiplier. It does. Figure 3 shows the estimated multipliers as a function of the initial regime.¹³

Figure 3 compares the distributions of the government spending multiplier across the initial monetary policy regimes using boxplots. If the consensus in the literature holds, then, after accounting for the response of the monetary policy regime, the boxplots *should* shift upwards for the more passive regimes. However, Figure 3 shows that there is little variation in the boxplots in a given time period after the shock. One quarter after the shock, the posterior median is around 5 regardless of the initial regime. One year after the shock, the posterior median is around 4. Five years after the shock, the posterior median is around 1 while the distributions are not entirely positive. Thus, we find that the multiplier decreases in magnitude over time, but this decrease does not seem to depend on the initial monetary policy regime as the literature has concurred.

The results of this exercise contradict the conventional wisdom that the government

¹³The results in the figure differ from those in theory, such as in Leeper et al. [2017]. But we can replicate the theoretical consensus by restricting monetary policy to remain constant after a fiscal shock. This result appears in detail in Section 5.

Figure 3: Estimated Multipliers when Monetary Policy is Fully Responsive



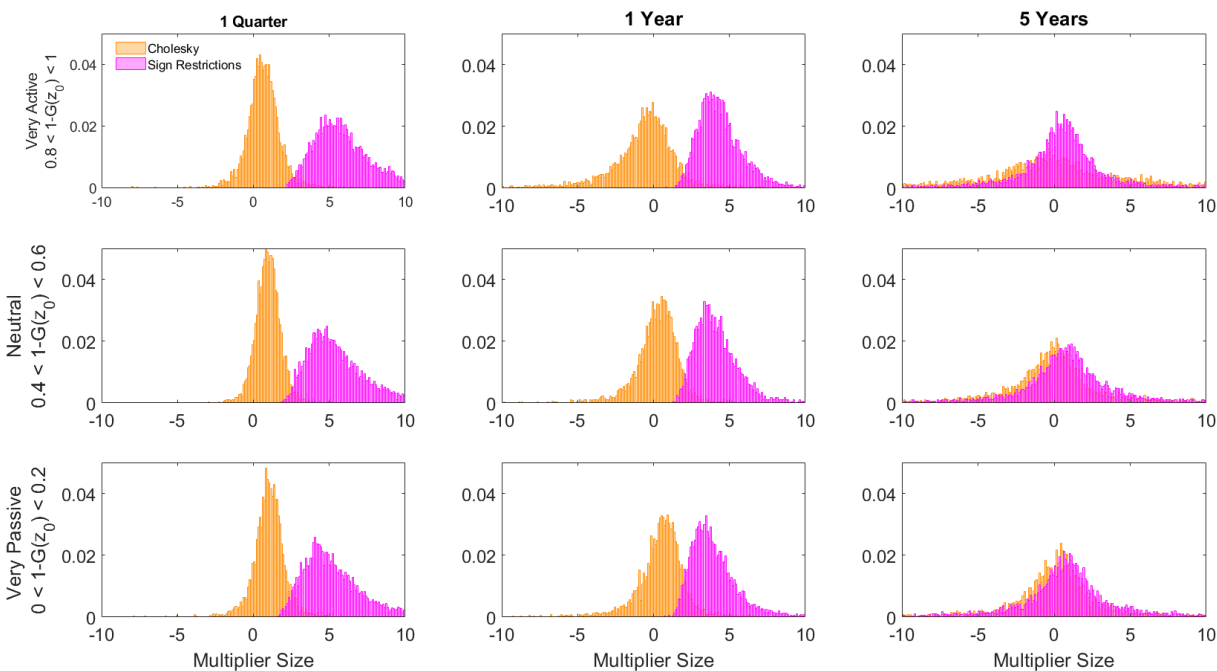
Note: The figure uses boxplots to show the estimated multiplier posterior distributions one quarter, one year, and five years after the shock across initial monetary policy regimes. The middle line and the box present the posterior median and the corresponding 50 percent credible bands. The upper and lower lines correspond to the highest and lowest value of the distribution that is not considered an outlier. The Figure suggests that the multiplier decreases over time but is almost completely unaffected by the initial monetary policy regime.

spending multiplier is larger when monetary policy is passive. Our findings suggest that the multiplier does not depend on monetary policy, either in the short run or in the long run. This largely reflects the fact that the central bank responds quickly after a shock, and that it transitions rapidly to an active regime even if the initial regime had not previously been “very active.” The consensus in the literature does not account for this response of the monetary policy regime to the government spending shock. The consensus result hinges on a model assumption that lacks empirical support, and vanishes once that assumption is relaxed. Thus, we must conclude that the consensus – that government spending multipliers are larger when monetary policy is passive – is artificial, rather than a feature of the data.

Furthermore, the results given in Figure 3 suggest that multiplier estimates in the short run are considerably larger than in the long run. The empirical government-spending literature largely agrees that the multiplier lies between 0.3 and 2.1 [Ramey, 2016, Ramey and Zubairy, 2018]. However, the multiplier can also be negative [Perotti, 2014] and as high as 3.5 [Edelberg et al., 1999]. Our estimated short- and medium-run multiplier estimates

are considerably larger. However, the literature has either focused only on longer-run multipliers, and/or has employed identification strategies related to the Cholesky approach, in which government spending is ordered first in a recursive VAR. This method imposes zero restrictions, which require one to assume that governments adjust their spending only with a delay when responding to shocks other than to government spending itself. This can be a strong assumption. Recent events such as the Troubled Asset Relief Program (TARP) of 2008 or the CARES Act of 2020 have shown that governments can and do react fast to changes in the business cycle, at least during times of crises.

Figure 4: Multiplier Comparison: Recursive-Identification vs. Sign-Restriction Approach



Note: Comparison of recursive (orange) and sign-restricted (violet) multiplier estimates for the “very active,” “neutral,” and “very passive” monetary policy regime. The figure shows that estimated multipliers are similar after five years but differ in the short and medium runs.

In Figure 4, we compare the multiplier estimates from our main exercise (violet histograms) to those using the Cholesky approach (orange histograms). We find smaller estimates in the short and medium runs for the Cholesky approach. The estimates are similar after five years. The comparison of multiplier estimates using different identification approaches is an important exercise. The findings indicate that once we consider multipliers in the short and medium runs, and replace the zero restrictions related to the recursive approach with our mostly uncontroversial sign restrictions, the multiplier may be larger than previously conjectured in the literature. We next provide additional evidence that the

constant-regime assumption is the key factor underlying the conventional wisdom.

5 Counterfactuals

Because our main results contradict the theoretical consensus, we conduct a counterfactual analysis. First, we analyze what would happen to the government spending multiplier if the central bank were to keep its monetary policy regime temporarily constant after the shock. Second, we fully replicate the framework that underlies the consensus in the literature with our empirical model. To do this, we distinguish only between the purely active and the purely passive monetary policy regimes, and we keep the monetary policy regimes constant for the entire horizon after the shock. The findings of these exercises suggest that the constant-regime assumption drives the result in the theoretical literature.

Because the policymaker *can* commit to a policy rule for an extended period of time, these contingencies represent hypothetical policy scenarios.

5.1 Government Spending Multiplier under temporary constant Monetary Policy

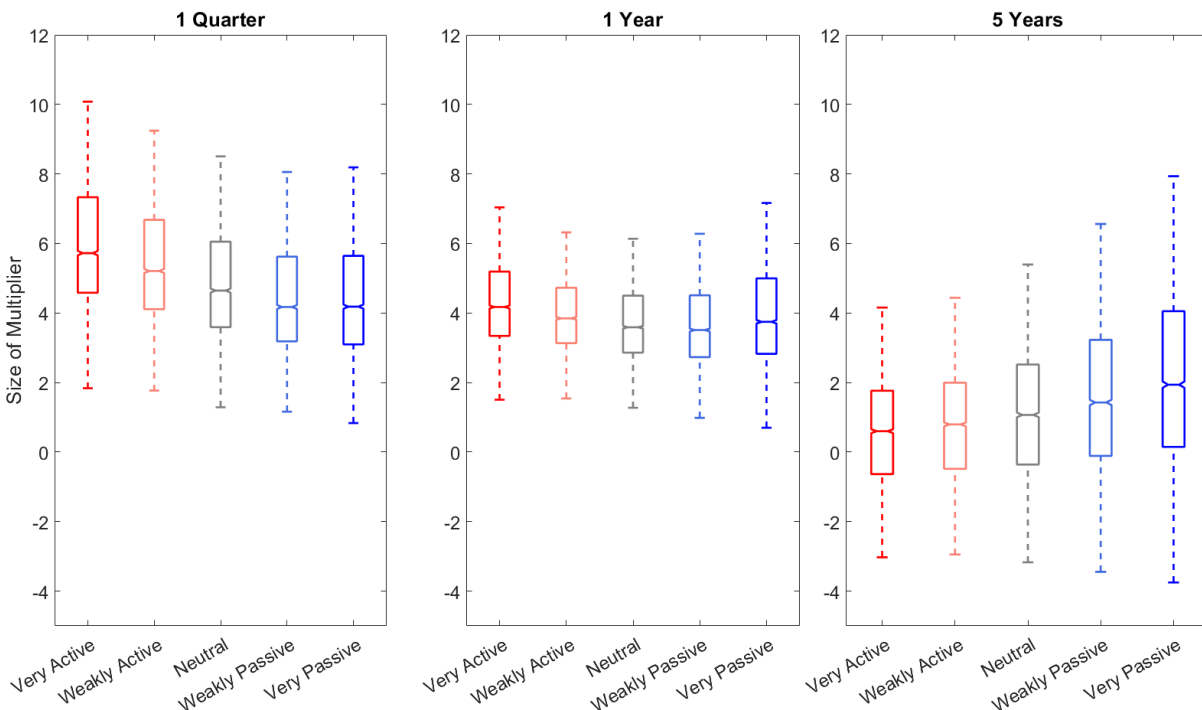
In Section 4, we found that the central bank adjusted its monetary policy regime immediately after the government spending shock when the regime was not already “very active.” However, the central bank can also choose to maintain its policy of the impact period for some specific time after the shock. In fact, the monetary policy regime *is* nearly constant during certain subperiods of our sample period. For instance, in Figure 1, the monetary policy regime is purely active throughout the 1980s and the second half of the 1990s. By contrast, the monetary policy regime is purely passive in the second half of the 1970s and the first half of the 2000s. Sometimes the central bank signals its intentions to hold policy constant in the future. For example, on April 29th, 2020, the Federal Reserve announced:

“The ongoing public health crisis will weigh heavily on economic activity, employment, and inflation in the near term, and poses considerable risks to the economic outlook over the medium term. In light of these developments, the Committee decided to maintain the target range for the federal funds rate at 0 to 1/4 percent. The Committee expects to maintain this target range until it is confident that the economy has weathered recent events and is on track to achieve its maximum employment and price stability goals.”¹⁴

¹⁴Source: <https://www.federalreserve.gov/newsevents/pressreleases/monetary20200429a.htm>

Hence, we address this possibility directly. First, we analyze what would happen to the government spending multiplier if the central bank were to keep its policy temporarily constant after the government spending shock. To do this, we employ the generalized impulse response functions but keep the inflation parameter $\phi_{\pi,t+h}$ constant for one year, two years and five years. After these periods expire, we continue as before by updating $\phi_{\pi,t+h}$ via the “rolling” Kalman filter for each period h of the forecast horizon. Figures 5 and 6 display the corresponding multiplier estimates when the monetary policy regime is restricted to remain in its initial regime for one year and five years after the shock, respectively.

Figure 5: Estimated Multipliers when Monetary Policy is Constant for One Year

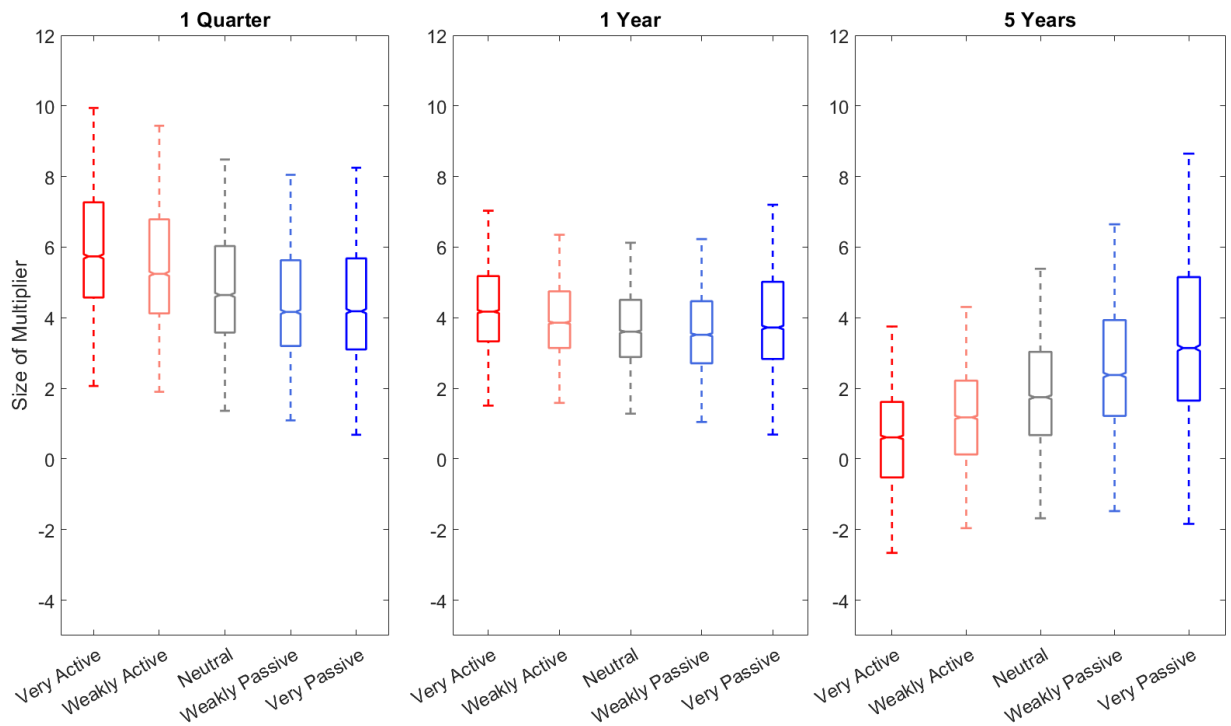


Note: The figure uses boxplots to show the estimated multiplier distributions across initial monetary policy regimes when the regime is kept constant for one year after the shock. The middle line and the box present the posterior median and 50 percent credible bands of the corresponding distribution. The upper and lower lines correspond to the highest and lowest values of the distribution that are not considered to be outliers. The figure suggests that the multiplier decreases over time but is almost completely unaffected the initial monetary policy regime even if the regime is held constant for one year.

If the central bank keeps its policy regime constant for one year after the shock, the results are similar to those in Section 4. The government spending multiplier decreases in magnitude over time, but it does not vary significantly with the initial monetary policy regime. We find similar results if we keep the monetary policy regime constant for two

years.¹⁵ By contrast, when the central bank keeps its policy regime constant for five years after the shock, we do observe a higher multiplier estimate in more passive regimes after five years. This is in line with the conventional wisdom. While this finding indicates the possibility that the government spending multiplier may depend on monetary policy, the difference in the multiplier hinges crucially on the central bank’s willingness to maintain the monetary policy regime for extended periods of time. According to our analysis, the central bank must maintain its initial policy rule for more than two years before one can discern any noticeable difference in the government spending multiplier with respect to monetary policy.

Figure 6: Estimated Multipliers when Monetary Policy is Constant for Five Years



Note: The figure uses boxplots to show the estimated multiplier distributions across initial monetary policy regimes when the regime is kept constant for five years after the shock. The middle line and the box present the posterior median and 50 percent credible bands of the corresponding distribution. The upper and lower lines correspond to the highest and lowest values that are not considered to be outliers of the distributions. The figure suggests that there multiplier is larger for the more passive regimes in the long run but only if the monetary policy regime is held constant for five years.

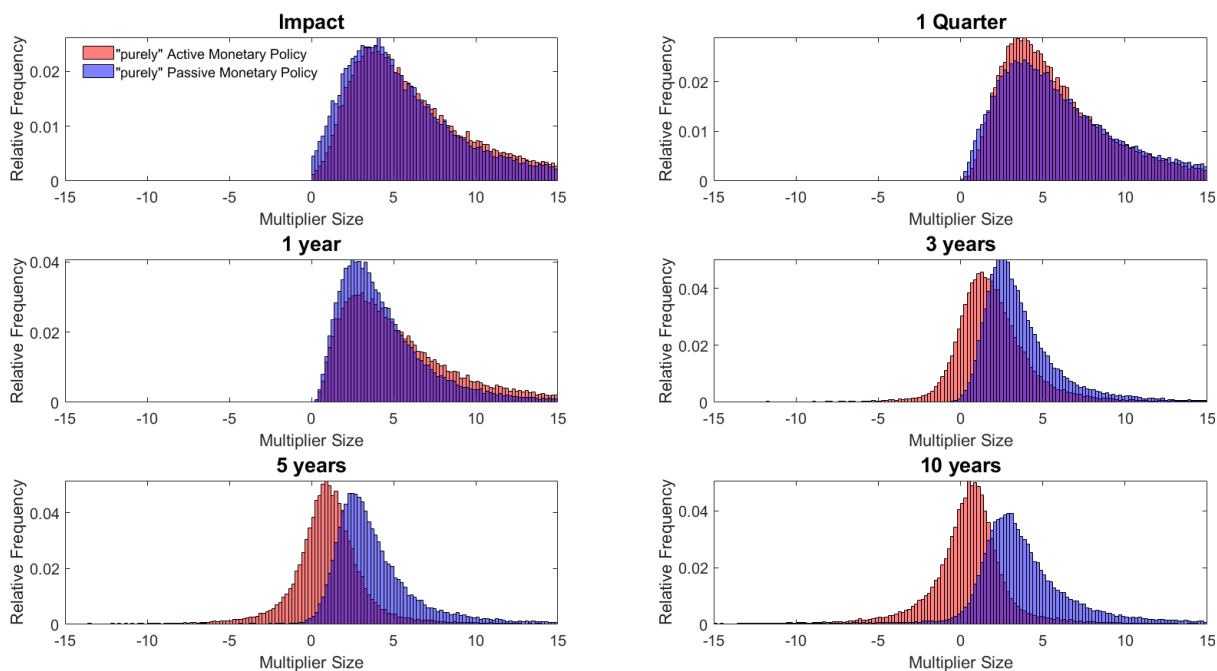
¹⁵Results are available upon request.

5.2 Government Spending Multiplier under Fully Constant Monetary Policy Regimes

We now “replicate” the theoretical framework with our empirical model. Recall that theory interprets monetary policy as binary, and only distinguishes between “active” and “passive” monetary policy regimes. In addition, some studies use the Markov-switching approach to model policy changes, which implies that the central bank jumps from one monetary policy regime directly to another. Lastly, theory treats the monetary policy regime constant for the entire time after the shock when they compute impulse response functions.

To implement the same framework, we estimate our model with an exogenous γ that we calibrate to be very large. This choice ensures that the central bank jumps from one regime directly to another. We then use traditional impulse response functions to estimate the dynamic effects of the shock. The traditional impulse response functions estimate the effects for the purely active (*AM*) and the purely passive (*PM*) regime, rather than an interior combination. Finally, the traditional impulse response functions keep the regimes constant for the entire forecasted horizon after the shock. Figure 7 presents the results.

Figure 7: Estimated Multipliers when Monetary Policy is Fully Constant



Note: The figure shows the estimated multiplier distributions for the most active (red) and most passive (blue) monetary policy regimes when the regimes are kept constant for the entire time after the shock. The figure suggests that the multiplier is larger when monetary policy is and remains “purely” passive.

Figure 7 illustrates that if we replicate the theoretical framework with our empirical

model, then we can replicate the theoretical consensus. Up to one year after the shock, the distributions of the multiplier highly overlap so that there is no meaningful difference in the estimated multiplier between the two most extreme monetary policy regimes. One year after the shock, the estimated multiplier distributions start to separate. In the long run, the multiplier is estimated to be higher when monetary policy is and remains purely passive. This evidence matches the theoretical consensus, and mirrors the results in Leeper et al. [2017], who find comparable multipliers in the short run but a higher multiplier under passive monetary policy in the long run. We also obtain similar results when we conduct the same analysis, but estimate the full posterior distribution of our model including an estimated γ , instead of setting γ exogenously equal to a large number (see Figure 9).

Because we find regime-dependent multipliers *only when we restrict the monetary policy regime to remain unchanged* as in these two exercises, we conclude the conventional wisdom in the literature is largely driven by that constant-regime assumption. Regardless of whether we interpret monetary policy as a continuous or a binary process, or whether we model smooth or abrupt regime changes, the government spending multiplier diverges in the long run only if we keep the monetary policy regime constant for at least two years. When we relax this assumption and allow the central bank to respond freely after the shock, the multipliers ceased to diverge. These results indicate that to boost the impact of government spending on the economy, the central bank would need to accommodate inflation for a surprisingly long period of time after a government spending shock. However, this would conflict with the central bank's main mandate to maintain price stability, and it would violate the Taylor principle.

6 Government Spending Multiplier at the Zero Lower Bound

The most conspicuous exception to the active/passive characterization of monetary policy is the zero lower bound, which is a defining feature of the modern economy. Following the 2008 financial crisis, the Fed kept nominal interest rates at zero between 2008Q4 and 2015Q4. During that period, the federal funds rate was unresponsive despite large fiscal policy interventions such as the Troubled Asset Relief Program (TARP) and the American Recovery and Reinvestment Act (ARRA). In 2020, nominal interest rates fell back toward zero, and, as of this writing, many countries are at the zero lower bound. In addition, the Federal Reserve has announced its policy to keep interest rates low even if inflation exceeds its fixed target of two percent. This suggests that the zero lower bound will remain in place

for the foreseeable future, at least in the United States, despite the \$2.2 trillion Coronavirus Aid, Relief and Economic Security (CARES) Act. These characteristics of the zero lower bound signal a major departure from the Fed’s historical behavior. In this section, we estimate the multiplier when nominal interest rates are stuck at zero. To do this we consider impulse response functions where the monetary policy regime remains fixed for the entire forecast horizon after the shock.

Our previous analysis truncated the data sample in 2007Q4 to avoid contaminating findings with the Great Recession and its unconventional monetary policy, but there is substantial interest in the size of the government spending multiplier when monetary policy is constrained by the zero lower bound. This question has been at the center of policy and academic debates since the 2008 financial crisis. However, the size of the multiplier when nominal interest rate are stuck at zero remains an open question. For example, different theories provide opposing predictions. On the one hand, Woodford [2011], Christiano et al. [2011], and Eggertsson [2011] predict that the government spending multiplier is higher when monetary policy accommodates inflation, and that it is especially high when monetary policy is constrained by the zero lower bound. This belief has been challenged by Aruoba and Schorfheide [2013], Mertens and Ravn [2014], Kiley [2016], and Wieland [2018], who argue that the multiplier at the zero lower bound can be below one. The disagreement leaves the question to empiricists for adjudication.

However, empirical studies reach different conclusions as well. Almunia et al. [2010] and Gordon and Krenn [2010] find multipliers above two for the zero lower bound period during the Great Depression. Miyamoto et al. [2018] estimate the multiplier at the modern zero lower bound in Japan to lie between 1.5 and 2. By contrast, Ramey and Zubairy [2018] provide evidence that the government spending multiplier at the (modern) zero lower bound is below one. Similar estimates can be found in Ramey [2011] and Crafts and Mills [2013] for the zero lower bound epochs from the first half of 20th century in the U.S. and the U.K., respectively.

In this section, we conduct several exercises to analyze the government spending multiplier at the zero lower bound. First, we estimate our model for an extended sample period up to 2015Q4 in order to include the zero lower bound. The federal funds rate is stuck at zero between 2008Q4 and 2015Q4, which prevents us from extracting useful information during that period. Hence, we follow Wu and Xia [2016], and replace the federal funds rate during the zero lower bound by the shadow rate.¹⁶ Second, we estimate a linear VAR for the period between 2008Q4 and 2015Q4, and compare the multiplier estimates with those from

¹⁶Wu and Xia [2016] estimate the shadow rate using a shadow rate term structure model in which the shadow rate is a linear combination of latent factors. See Wu and Xia [2016] for an overview.

a version of our replication exercise in Section 5 where we estimate the full posterior of our model, including the posterior of γ and c .

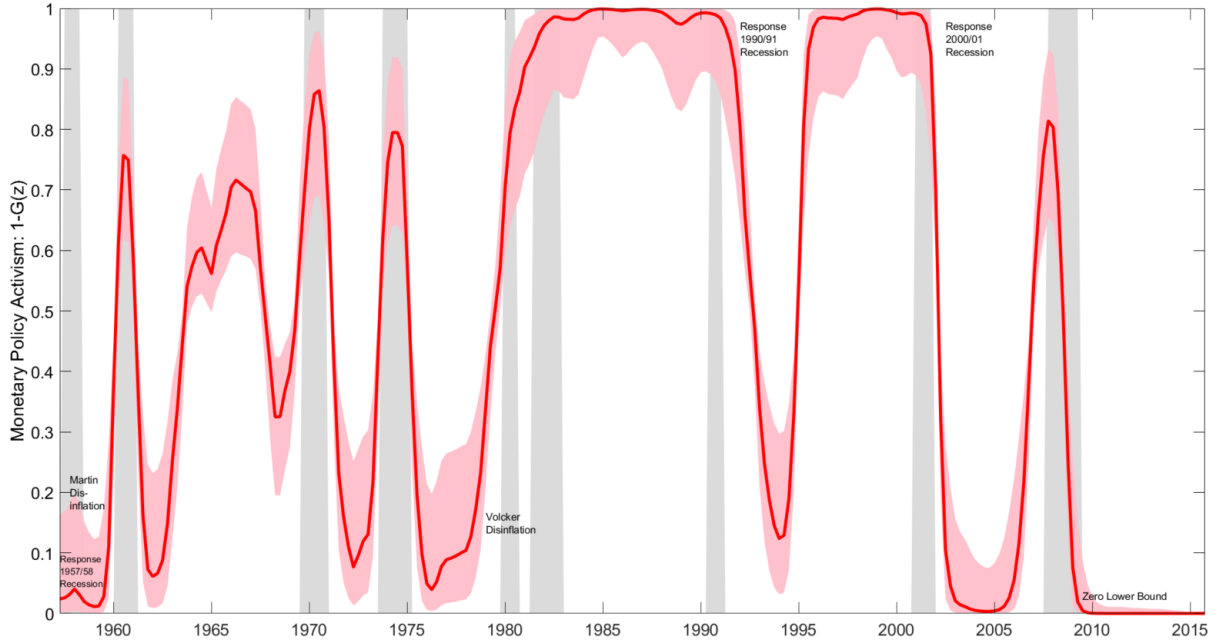
Figure 8 displays the evolution of the monetary policy regime for the extended sample period. We can see that $1 - G(z_{t-1})$ is zero between 2009 and the end of the sample with very little uncertainty. This observation suggests that the zero lower bound is a very passive monetary policy regime. This assessment supports the “crowding-in” argument proposed by Woodford [2011], Christiano et al. [2011] and Eggertsson [2011] that the government spending multiplier should be particularly large at the zero lower bound because the central bank does not respond to inflation, which lowers the real interest rate, and leads households to increase consumption. However, if we compare the multiplier between the purely active and the purely passive regimes, we find that the distributions highly overlap in the short and long runs; this contradicts the divergence result from Figure 7.¹⁷ Were the zero lower bound truly a purely passive regime, we would expect the corresponding multiplier to be larger than the multiplier from the purely active regime. Hence, we conclude the modern zero lower bound in the United States does not correspond to a purely passive regime. It is likely characterized by other unexplored factors.

To illustrate further, we estimate a linear VAR model for the zero lower bound period, and compare the multiplier to those from the purely active and purely passive regimes when we estimate the full posterior of our baseline model. Figure 9 displays the results. The related literature uses a variety of identification strategies, so we report both Cholesky and sign-restricted multipliers. In the left column of Figure 9, we identify the government spending shock using the Cholesky method. Following Blanchard and Perotti [2002], we order government spending first in a recursive VAR. In the right column, we apply our sign-restriction approach. When we identify the government spending shock using the recursive approach, we find that the government spending multiplier at the zero lower bound is very small, similar to the multiplier under the purely active monetary policy regime. Regardless of the forecast horizon, the distribution is centered around zero. Hence, the multiplier at the zero lower bound may not even be positive. This result is in line with that of Ramey and Zubairy [2018], who also find small multipliers at the zero lower bound. By contrast, when we identify the government spending shock using our sign-restriction approach, we find that the multiplier is large and similar to the multiplier under the purely passive monetary policy regime.

These comparisons between the multipliers at the zero lower bound using different identification strategies yield interesting insights. Using the recursive approach, the multiplier is smaller than the estimate when we employ our sign-restriction approach (even after taking

¹⁷Results are available upon request.

Figure 8: Evolution of Monetary Policy Between 1954 and 2016

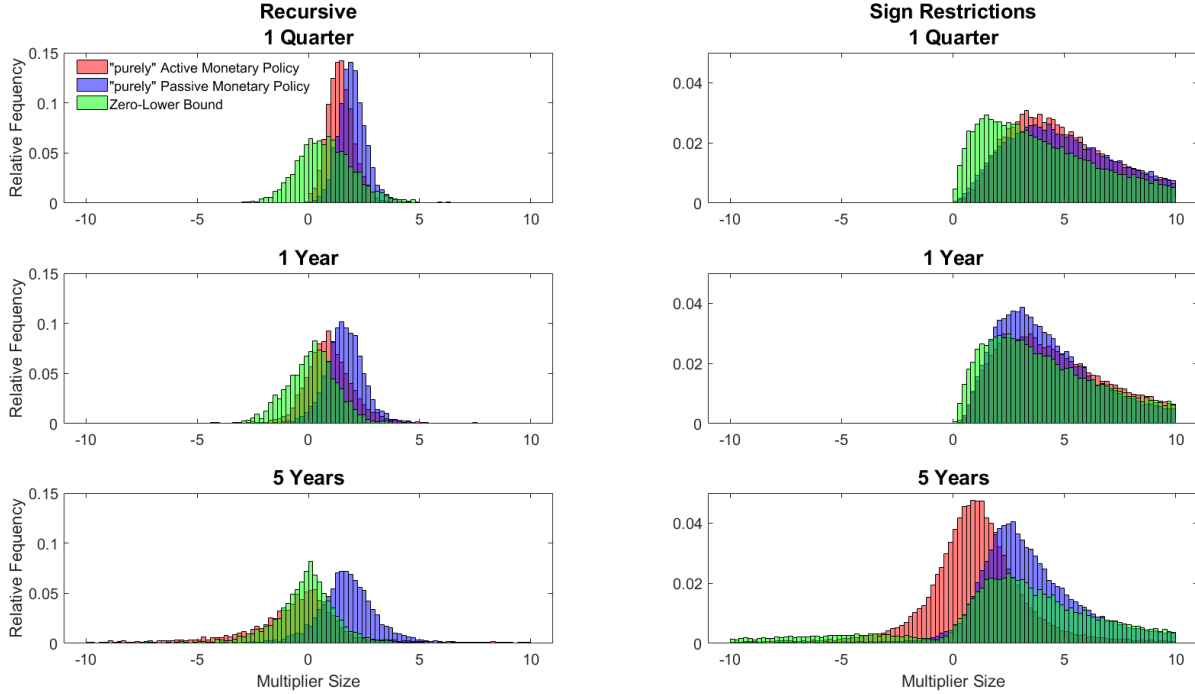


Note: The figure shows the pointwise-posterior median estimate of $1 - G(z)$, along with the 68 percent credible bands. $1 - G(z)$ can be interpreted as a measure for monetary policy activism. Grey bars represent recessions as defined by the National Bureau of Economic Research. This Figure represents an extension of Figure 1. During the zero lower bound era (2008Q4-2015Q4), $1 - G(z)$ is consistently zero with very little uncertainty indicating that the the zero lower bound era is an “extremely passive” monetary policy regime.

the estimation uncertainty into account). The Cholesky approach, as used in Ramey and Zubairy [2018], is related to timing restrictions which may be violated, especially during periods in which the central bank is constrained by the zero lower bound. Ordering government spending first in a recursive VAR implies that governments can change their spending plans only with a delay when responding to shocks other than to government spending itself. This was once a popular approach. The CARES Act of 2020 represents the largest economic stimulus package of any kind in U.S. history, and it was debated and passed into law in only a few weeks. The Paycheck Protection Program (PPP) attached to that package disbursed \$349 billion in less than two weeks. The TARP of 2008 similarly disbursed hundreds of billions in the same quarter in which Congress passed the bill. This demonstrates that governments *can and do* react quite fast during times of crises, and that practices in the current era violate the timing restrictions of the recursive approach.

By contrast, our set of sign restrictions is consistent with a large class of neoclassical and new Keynesian models, and hence, has theoretical foundations. These sign restrictions also allow government spending to respond quickly, as was the case in the CARES, PPP,

Figure 9: Multiplier Estimates at the Zero Lower Bound



Note: The figure shows the distribution of the estimated multipliers at the zero lower bound (green) and when the monetary policy regime is and remains “purely active” (red) and “purely passive” (blue) after the shock. The figure compares estimates from the standard Cholesky approach (left) and our sign-restriction approach (right). Results suggest that the strategy for identifying government spending shocks matters when nominal interest rates are stuck at zero.

and TARP examples. Given this additional evidence, we conclude that the multiplier at the modern zero lower bound in the United States is larger than had been shown by previous estimates in the literature (e.g., Ramey and Zubairy [2018]). For the period between 2008Q4 and 2015Q4, our median multiplier estimate is 4.5 on impact and decreases to 3 after five years.

7 Conclusion

This paper develops a flexible, nonlinear, structural vector autoregression model to investigate the relationship between the government spending multiplier and monetary policy. Conventional wisdom suggests that the multiplier is larger when nominal interest rates respond less than one for one to inflation. However, models supporting this consensus keep the monetary policy regime constant after the government spending shock when they estimate multipliers. As a result, the literature ignores how the *central bank* adjusts its policy

regime in response to the economic conditions after a change in government spending. Our approach relaxes this assumption, and allows the central bank to update its policy regime after a government spending intervention.

Our analysis shows that the central bank responds quickly after the shock, and that it transitions rapidly to an active regime even if the initial regime had not previously been active. The response of the monetary policy regime to the government spending shock has vital implications for the multiplier. Once we account for the reaction of the policy regime, the relationship between the multiplier and monetary policy vanishes. By contrast, when we keep the monetary policy regime counterfactually constant after the shock, we find multiplier estimates that support the conventional wisdom. This leads us to conclude that the key driver of the conventional wisdom is the constant-regime assumption — not the data.

An exception to the active/passive characterization of monetary policy occurs when nominal interest rates are stuck at zero. We analyze the multiplier at the zero lower bound by keeping the monetary policy regime constant after the government spending increase. We then compare multiplier estimates employing different strategies for identifying government spending shocks. When we use the standard Cholesky approach, we find multiplier estimates near zero. On the contrary, when we apply our sign restriction method, the multiplier estimates exceed one. The timing restrictions related to the Cholesky approach may be violated, especially at the zero lower bound. This scheme requires us to assume that governments react to changes in the business cycle only with a delay. However, both the Troubled Asset Relief Program of 2008 and the Coronavirus Aid, Relief and Economic Security (CARES) Act of 2020 in the United States have demonstrated in a dramatic way that governments can and do react quite fast during times of crises (e.g., when nominal interest rates are cut to zero in response to a severe economic downturn).

Our analysis highlights the necessity of accounting for the reaction of monetary policy to the government spending shock to properly study the relationship between monetary policy and the multiplier. Failure to do so ignores the central bank's ability to respond to the shock. This leads to a misrepresentation of how the multiplier depends on monetary policy. Our results indicate that to boost the impact of government spending on the economy, the central bank would need to tolerate inflation for a long period of time — for longer than two years, according to our estimation. However, this creates a dilemma because it would require the central bank to violate its main mandate of price stability, and would conflict with the Taylor principle. In addition, we show that the identification of government spending shocks matters when monetary policy is constrained by the zero lower bound. Once we employ our sign restrictions on impulse response functions for identification, the multiplier at the modern zero lower bound in the United States is larger than has previously been found (e.g.,

Ramey and Zubairy [2018]). This result suggests that recent fiscal stimulus packages, such as the American Recovery and Reinvestment Act of 2009 and the CARES Act of 2020, may have larger economic effects than the literature has formerly concluded.

References

- S. R. Aiyagari, L. J. Christiano, and M. Eichenbaum. The output, employment, and interest rate effects of government consumption. *Journal of Monetary Economics*, 30(1):73–86, 1992.
- M. Almunia, A. Benetrix, B. Eichengreen, K. H. O’Rourke, and G. Rua. From great depression to great credit crisis: similarities, differences and lessons. *Economic policy*, 25(62): 219–265, 2010.
- J. E. Arias, D. Caldara, and J. F. Rubio-Ramirez. The systematic component of monetary policy in svars: An agnostic identification procedure. *Journal of Monetary Economics*, 101:1–13, 2019.
- S. B. Aruoba and F. Schorfheide. Macroeconomic dynamics near the zlb: A tale of two equilibria. 2013.
- A. J. Auerbach and Y. Gorodnichenko. Measuring the output responses to fiscal policy. *American Economic Journal: Economic Policy*, 4(2):1–27, 2012.
- C. Baumeister and J. D. Hamilton. Inference in structural vector autoregressions when the identifying assumptions are not fully believed: Re-evaluating the role of monetary policy in economic fluctuations. *Journal of Monetary Economics*, 100:48–65, 2018.
- M. Baxter and R. G. King. Fiscal policy in general equilibrium. *The American Economic Review*, pages 315–334, 1993.
- N. Ben Zeev and E. Pappa. Chronicle of a war foretold: The macroeconomic effects of anticipated defence spending shocks. *The Economic Journal*, 127(603):1568–1597, 2017.
- F. Bianchi and C. Ilut. Monetary/fiscal policy mix and agents’ beliefs. *Review of economic Dynamics*, 26:113–139, 2017.
- O. Blanchard and R. Perotti. An empirical characterization of the dynamic effects of changes in government spending and taxes on output. *The Quarterly Journal of Economics*, 117(4):1329–1368, 2002. doi: 10.1162/003355302320935043.
- J. Boivin. Has us monetary policy changed? evidence from drifting coefficients and real-time data. *Journal of Money, Credit and Banking*, 38(5 (August 2006)):1149–1174, 2006.
- G. Caggiano, E. Castelnuovo, V. Colombo, and G. Nodari. Estimating fiscal multipliers: News from a non-linear world. *The Economic Journal*, 125(584):746–776, 2015.
- F. Canova and E. Pappa. Fiscal policy, pricing frictions and monetary accommodation. *Economic Policy*, 26(68):555–598, 2011.
- C. K. Carter and R. Kohn. On gibbs sampling for state space models. *Biometrika*, 81(3): 541–553, 1994.

- C. Carvalho, F. Nechio, and T. Tristao. Taylor rule estimation by ols. *Available at SSRN 3265449*, 2019.
- Y. Chang and B. Kwak. Us monetary-fiscal regime changes in the presence of endogenous feedback in policy rules. 2017.
- S. Chib and E. Greenberg. Understanding the metropolis-hastings algorithm. *The american statistician*, 49(4):327–335, 1995.
- L. Christiano, M. Eichenbaum, and S. Rebelo. When is the government spending multiplier large? *Journal of Political Economy*, 119(1):78–121, 2011.
- R. Clarida, J. Gali, and M. Gertler. Monetary policy rules and macroeconomic stability: evidence and some theory. *The Quarterly journal of economics*, 115(1):147–180, 2000.
- J. S. Cloyne, Ò. Jordà, and A. M. Taylor. Decomposing the fiscal multiplier. Technical report, National Bureau of Economic Research, 2020.
- T. Cogley and T. J. Sargent. Evolving post-world war ii us inflation dynamics. *NBER macroeconomics annual*, 16:331–373, 2001.
- T. Cogley and T. J. Sargent. Drifts and volatilities: monetary policies and outcomes in the post wwii us. *Review of Economic dynamics*, 8(2):262–302, 2005.
- O. Coibion and Y. Gorodnichenko. Monetary policy, trend inflation, and the great moderation: An alternative interpretation. *American Economic Review*, 101(1):341–70, 2011.
- N. Crafts and T. C. Mills. Rearmament to the rescue? new estimates of the impact of “keynesian” policies in 1930s’ britain. *The Journal of Economic History*, 73(4):1077–1104, 2013.
- T. Davig and E. M. Leeper. Monetary–fiscal policy interactions and fiscal stimulus. *European Economic Review*, 55(2):211–227, 2011.
- C. de Vries and W. Li. Estimating monetary policy rules with serially correlated monetary policy shocks. *Unpublished manuscript*, 2013.
- B. Dupor and R. Li. The expected inflation channel of government spending in the postwar us. *European Economic Review*, 74:36–56, 2015.
- W. Edelberg, M. Eichenbaum, and J. D. Fisher. Understanding the effects of a shock to government purchases. *Review of Economic Dynamics*, 2(1):166–206, 1999.
- G. B. Eggertsson. What fiscal policy is effective at zero interest rates? *NBER Macroeconomics Annual*, 25(1):59–112, 2011.
- A. Ellahie and G. Ricco. Government purchases reloaded: Informational insufficiency and heterogeneity in fiscal VARs. *Journal of Monetary Economics*, 90:13–27, 2017. doi: 10.1016/j.jmoneco.2017.06.002.

- A. Fatás and I. Mihov. The effects of fiscal policy on consumption and employment: theory and evidence. 2001.
- M. Forni and L. Gambetti. Sufficient information in structural vars. *Journal of Monetary Economics*, 66:124–136, 2014.
- S. Frühwirth-Schnatter. Data augmentation and dynamic linear models. *Journal of time series analysis*, 15(2):183–202, 1994.
- J. Galí, J. D. López-Salido, and J. Vallés. Understanding the effects of government spending on consumption. *Journal of the european economic association*, 5(1):227–270, 2007.
- A. B. Galvão and M. T. Owyang. Financial stress regimes and the macroeconomy. *Journal of Money, Credit and Banking*, 50(7):1479–1505, 2018.
- D. Gefang and R. Strachan. Nonlinear Impacts of International Business Cycles on the UK—A Bayesian Smooth Transition VAR Approach. *Studies in Nonlinear Dynamics and Econometrics*, 14(1), 2010.
- R. J. Gordon and R. Krenn. The end of the great depression 1939-41: Policy contributions and fiscal multipliers. Technical report, National Bureau of Economic Research, 2010.
- M. T. Kiley. Policy paradoxes in the new keynesian model. *Review of Economic Dynamics*, 21:1–15, 2016.
- S. Kim. Structural shocks and the fiscal theory of the price level in the sticky price model. *Macroeconomic Dynamics*, 7(5):759–782, 2003.
- S. Kim, N. Shephard, and S. Chib. Stochastic volatility: likelihood inference and comparison with arch models. *The review of economic studies*, 65(3):361–393, 1998.
- G. Koop, M. H. Pesaran, and S. M. Potter. Impulse response analysis in nonlinear multivariate models. *Journal of econometrics*, 74(1):119–147, 1996.
- S. Laumer. Government spending and heterogeneous consumption dynamics. *Journal of Economic Dynamics and Control*, page 103868, 2020.
- S. Laumer and C. Philipps. Does the government spending multiplier depend on the business cycle? 2020.
- E. M. Leeper, T. B. Walker, and S.-C. S. Yang. Fiscal foresight and information flows. *Econometrica*, 81(3):1115–1145, 2013.
- E. M. Leeper, N. Traum, and T. B. Walker. Clearing up the fiscal multiplier morass. *American Economic Review*, 107(8):2409–54, 2017.
- L. Linnemann and A. Schabert. Productive government expenditure in monetary business cycle models. *Scottish Journal of Political Economy*, 53(1):28–46, 2006.

- M. Lippi and L. Reichlin. The dynamic effects of aggregate demand and supply disturbances: Comment. *The American Economic Review*, 83(3):644–652, 1993.
- K. R. Mertens and M. O. Ravn. Fiscal policy in an expectations-driven liquidity trap. *The Review of Economic Studies*, 81(4):1637–1667, 2014.
- W. Miyamoto, T. L. Nguyen, and D. Sergeyev. Government spending multipliers under the zero lower bound: Evidence from japan. *American Economic Journal: Macroeconomics*, 10(3):247–77, 2018.
- A. Mountford and H. Uhlig. What are the effects of fiscal policy shocks? *Journal of Applied Econometrics*, 24(6):960–992, sep 2009. doi: 10.1002/jae.1079.
- A. Orphanides. Monetary policy rules based on real-time data. *American Economic Review*, 91(4):964–985, 2001.
- A. Orphanides. Monetary-policy rules and the great inflation. *American economic review*, 92(2):115–120, 2002.
- A. Orphanides. Monetary policy rules, macroeconomic stability, and inflation: A view from the trenches. *Journal of Money, Credit and Banking*, pages 151–175, 2004.
- A. Orphanides and J. C. Williams. Robust monetary policy rules with unknown natural rates. *Brookings Papers on Economic Activity*, 2002(2):63–145, 2002.
- R. Perotti. Defense government spending is contractionary, civilian government spending is expansionary. Technical report, National Bureau of Economic Research, 2014.
- G. E. Primiceri. Time varying structural vector autoregressions and monetary policy. *The Review of Economic Studies*, 72(3):821–852, 2005.
- V. A. Ramey. Identifying government spending shocks: It’s all in the timing. *Quarterly Journal of Economics*, 126(1):1–50, 2011. doi: 10.1093/qje/qjq008.
- V. A. Ramey. Macroeconomic shocks and their propagation. In *Handbook of macroeconomics*, volume 2, pages 71–162. Elsevier, 2016.
- V. A. Ramey and M. D. Shapiro. Costly capital reallocation and the effects of government spending. *Carnegie-Rochester Conference Series on Public Policy*, 48:145–194, 1998. doi: 10.1016/S0167-2231(98)00020-7.
- V. A. Ramey and S. Zubairy. Government spending multipliers in good times and in bad: evidence from US historical data. *Journal of Political Economy*, 126(2):850–901, 2018.
- C. Ritter and M. A. Tanner. Facilitating the Gibbs sampler: the Gibbs stopper and the griddy-Gibbs sampler. *Journal of the American Statistical Association*, 87(419):861–868, 1992.
- C. D. Romer and D. H. Romer. Choosing the federal reserve chair: lessons from history. *Journal of Economic Perspectives*, 18(1):129–162, 2004.

- C. D. Romer and D. H. Romer. The most dangerous idea in federal reserve history: Monetary policy doesn't matter. *American Economic Review*, 103(3):55–60, 2013.
- J. F. Rubio-Ramirez, D. F. Waggoner, and T. Zha. Structural vector autoregressions: Theory of identification and algorithms for inference. *The Review of Economic Studies*, 77(2):665–696, 2010.
- C. Sims. Discussion of cogley and sargent, 'evolving post-world war ii inflation dynamics'. *NBER Macroeconomics Annual*, 16, 2001.
- J. B. Taylor. Discretion versus policy rules in practice. In *Carnegie-Rochester conference series on public policy*, volume 39, pages 195–214. Elsevier, 1993.
- J. B. Taylor. A historical analysis of monetary policy rules. In *Monetary policy rules*, pages 319–348. University of Chicago Press, 1999.
- J. B. Taylor. Housing and monetary policy. Technical report, National Bureau of Economic Research, 2007.
- H. Uhlig. What are the effects of monetary policy on output? Results from an agnostic identification procedure. *Journal of Monetary Economics*, 52(2):381–419, 2005. doi: 10.1016/j.jmoneco.2004.05.007.
- C. Walsh. Speed limit policies: the output gap and optimal monetary policy. *American Economic Review*, 93(1):265–278, 2003.
- J. Wieland. State-dependence of the zero lower bound government spending multiplier. *Unpublished manuscript*, 2018.
- M. H. Wolters. Estimating monetary policy reaction functions using quantile regressions. *Journal of Macroeconomics*, 34(2):342–361, 2012.
- M. Woodford. Simple analytics of the government expenditure multiplier. *American Economic Journal: Macroeconomics*, 3(1):1–35, 2011.
- J. C. Wu and F. D. Xia. Measuring the macroeconomic impact of monetary policy at the zero lower bound. *Journal of Money, Credit and Banking*, 48(2-3):253–291, 2016.
- S. Zubairy. On fiscal multipliers: Estimates from a medium scale dsge model. *International Economic Review*, 55(1):169–195, 2014.

A Mathematical Appendix

This section provides a more detailed overview about the new TVP-ST-VAR model used in the main text. The model is a hybrid between the smooth-transition VAR model, popularized by Auerbach and Gorodnichenko [2012], and a univariate regression with time-varying parameters and stochastic volatility. The latter part is used to estimate the state determining variable z_t which is then employed to distinguish between different monetary policy regimes in the main model. As argued in the main text, this extension is necessary because monetary policy is not well explained by a single observable variable so that the standard approach in the ST-VAR literature is not the optimal choice. In the following, we explain the model in greater details and lay out each step of the estimation procedure. First, the model is characterized via the following equations

$$X_t = (1 - G(z_{t-1}))\Pi_{AM}X_{t-1} + G(z_{t-1})\Pi_{PM}X_{t-1} + u_t \quad (\text{A.1.})$$

$$u_t \sim N(0, \Omega_t) \quad (\text{A.2.})$$

$$\Omega_t = (1 - G(z_{t-1}))\Omega_{AM} + G(z_{t-1})\Omega_{PM} \quad (\text{A.3.})$$

$$G(z_t) = \frac{1}{1 + \exp(\gamma(z_t - c))} \quad (\text{A.4.})$$

X_t is a vector of endogenous variables and represents the underlying economy. The state of the economy evolves continuously over time. The model A.1. - A.4. approximates the true evolution of this state via an evolving mix of two extreme states. Here, monetary policy is purely active in the *AM* and purely passive in the *PM* regime. The relative weights assigned to the purely passive regime is given by the transition function G . G is continuous, monotonically decreasing and bounded between 0 and 1. If z_t increases, the value of G decreases and the model assigns relatively more weight towards the *AM* regime. Here, γ governs the speed with which this transition occurs. If $\gamma \approx 0$, $G \approx 0.5 \forall t$, the model becomes a linear model and no transition occurs. If $\gamma \approx \infty$, the model jumps from one extreme regime to another as soon as z_t the threshold parameter c . In this case, the model becomes a threshold VAR model, thus nesting the Markov-type models.¹⁸

To characterize monetary policy as the evolving, potential nonlinear, state of the economy, we follow the DSGE literature and estimate an interest-rate rule, i.e., a version of the Taylor rule as introduced by Taylor [1993]. Because there is strong evidence that the coefficients of the Taylor rule and the variance of the corresponding residuals vary over time, we estimate a Taylor rule with time-varying parameters and stochastic volatility. More formally, we estimate

$$y_t = \mathbf{z}'_t \phi_t + e_t, \quad e_t \sim N(0, \sigma_t^2) \quad (\text{A.5.})$$

¹⁸We exploit this feature in our replication exercise. See Section 4.

$$\phi_t = \phi_{t-1} + v_t, v_t \sim N(0, q) \tag{A.6.}$$

$$\log(\sigma_t) = \log(\sigma_{t-1}) + \eta_t, \eta_t \sim N(0, w) \tag{A.7.}$$

where y_t is the federal funds rate, \mathbf{z}_t consists of a constant, inflation and output. ϕ_t is a vector of time-varying coefficients. ϕ_t follows a random walk process which allows for both temporary and permanent shifts in the parameters. We use the time-varying inflation parameter, $\phi_{\pi,t}$ as z_t in the model A.1. - A.4.. The variance of the residuals can also vary over time. We assume that $\log(\sigma_t)$ follows a random walk.

We estimate A.1.-A.4. and A.5.-A.7. separately in two steps using Bayesian estimation techniques. Bayesian methods allow the data to become informative about the model structure. This feature is particularly important because the data are also informative about the evolution of monetary policy which is key to understanding the effect of government spending depends on monetary policy. We next lay out each step of the estimation procedure.

A.1 Estimation of Taylor rule with time-varying coefficients and stochastic volatility

A.5. and A.6. form a linear-Gaussian state space model. This structure allows us to sample the latent states ϕ_t via standard (Kalman) filtering techniques. In contrast, A.5. and A.7. form a Gaussian but nonlinear state space model. We follow Kim et al. [1998] and Primiceri [2005] and transform this structure into a linear-Gaussian state space model and then use standard (Kalman) filtering techniques to sample σ_t . To sample the time-varying coefficients ϕ^T , the stochastic volatility σ^T and the hyperparameter q and w , we employ the following algorithm

Algorithm 1 Multi-Move Gibbs Sampler for TVP Taylor rule

1. Initialize: ϕ_0, q_0, σ_0 and w_0 ;
 2. Draw $\phi^T | X^T, q, \sigma^T, w$ using standard (Kalman) filtering methods;
 3. Draw $q | \phi^T$;
 4. Draw $\sigma^T | X^T, \phi^T, q, w$ using standard (Kalman) filtering methods;
 5. Draw $w | \sigma^T$
 6. Repeat steps 2 through 5 and keep the desired number of draws after a burn-in phase.
-

A.1.1 Time-varying coefficients

Conditional on σ^T and q , A.5. is linear and Gaussian with a known variance. Following Frühwirth-Schnatter [1994] and Carter and Kohn [1994], the density of α^T , $p(\alpha^T | X^T, \sigma^T, q)$, can be factored as

$$p(\phi^T | X^T, \sigma^T, q) = p(\phi_T | X_T, \sigma_T, q) \prod_{t=1}^{T-1} p(\phi_t | \phi_{t+1}, X^t, \sigma^T, q) \quad (\text{A.8.})$$

where

$$p(\phi_t | \phi_{t+1}, X^t, \sigma^T, q) \sim N(\phi_t | \phi_{t+1}, P_{t|t+1}) \quad (\text{A.9.})$$

$$\phi_t | \phi_{t+1} = E(\phi_t | \phi_{t+1}, X^t, \sigma^T, q) \quad (\text{A.10.})$$

$$P_{t|t+1} = \text{Var}(\phi_t | \phi_{t+1}, X^t, \sigma^T, q). \quad (\text{A.11.})$$

We use the (forward) Kalman filter to estimate $\phi_{T|T}$ and $P_{T|T}$, the mean and variance of the distribution of ϕ_T . A draw of this distribution is then used in the (backwards) Kalman smoother to estimate $\phi_{t|t+1}$ and $P_{t|t+1}$. Draws from the corresponding distributions yield the whole sequence of ϕ_t for $t = \{1, \dots, T-1\}$. A general overview about Kalman filtering techniques is given in Appendix A.1.4.

A.1.2 Stochastic Volatility

Consider the following model

$$y_t - \mathbf{z}'_t \phi_t = y_t^* = \sigma_t \xi_t, \quad \xi_t \sim N(0, 1) \quad (\text{A.12.})$$

$$\log(\sigma_t) = \log(\sigma_{t-1}) + \eta_t, \quad \eta_t \sim N(0, w) \quad (\text{A.13.})$$

The model A.12. and A.13. forms a Gaussian, but nonlinear state space model. However, the model can be transformed into a linear model by squaring and taking the log of A.12.

$$\log(y_t^{*2}) = 2\log(\sigma_t) + \log(\xi_t^2) \quad (\text{A.14.})$$

or

$$y_t^{**} = 2h_t + \nu_t. \quad (\text{A.15.})$$

(A.15.) and (A.13.) form a linear model, but $\nu_t \sim \log \chi^2(1)$. Following Kim et al. [1998] and Primiceri [2005], we approximate the $\log \chi^2(1)$ distribution by a mixture of seven normal distributions with weights q_i , means $m_i - 1.2704$ and variance v_i^2 . The constants $\{q_i, m_i, v_i^2\}$ are known and provided in Table A.1. Following Kim et al. [1998] and Primiceri [2005], we define $s^T = [s_1, \dots, s_T]$ where s_t represents the indicator variable, i , that belongs to the normal distribution that is used to approximate the distribution of ν in period t , i.e., $\nu_t | s_t \sim N(m_i - 1.2704, v_i^2)$ and $\text{Prob}(s_t = i) = q_i$. Then, conditional on ϕ^T, w, s^T and the data (A.15.) and (A.13.) is approximately linear and Gaussian. This structure implies that h_t can be sampled using standard (Kalman) filtering. Similar to the step that samples ϕ^T , the distribution of h^T can be factored as

$$p(h^T | X^T, \phi^T, w, s^T) = p(h_T | X_T, \phi_T, w, s^T) \prod_{t=1}^{T-1} p(h_t | h_{t+1}, X^t, \phi^T, w, s^T) \quad (\text{A.16.})$$

with

$$p(h_t|h_{t+1}, X^t, \phi^T, w, s^T) \sim N(h_{t|t+1}, B_{t|t+1}) \quad (\text{A.17.})$$

$$h_{t|t+1} = E(h_t|h_{t+1}, X^t, \phi^T, w, s^T) \quad (\text{A.18.})$$

$$B_{t|t+1} = \text{Var}(h_t|h_{t+1}, X^t, \phi^T, w, s^T). \quad (\text{A.19.})$$

Table A.1.: Mixture of Normal Distribution

w	q_i	m_i	v_i^2
1	0.00730	-10.12999	5.79596
2	0.10556	-3.97281	2.61369
3	0.00002	-8.566686	5.17950
4	0.044395	2.77786	0.16735
5	0.34001	0.61942	0.64009
6	0.24566	1.79518	0.34023
7	0.25750	-1.08819	1.26261

Note: See Kim et al. [1998] for details.

A.1.3 Hyperparameters for q and w

Finally, we discuss the hyperparameters for the variances q and w . Conditional on ϕ^T , σ^T and X^T , the residuals v_t and η_t are observable, and q and w both have an inverse Gamma distribution, i.e.

$$q \sim IG(\Psi_T, \psi_T) \quad (\text{A.20.})$$

$$w \sim IG(K_T, \kappa_T) \quad (\text{A.21.})$$

where $\Psi_T = \Psi_0 + \sum_{j=1}^{T-1} (\phi_j - \phi_{j-1})'(\phi_j - \phi_{j-1})$, $\psi_0 = \psi_T + T$, $K_T = K_0 + \sum_{j=1}^{T-1} (h_j - h_{j-1})^2$ and $\kappa_T = \kappa_0 + T$ with $\Psi_0 = 0.04 \times I$, $\psi_0 = 40$, $K_0 = 1$ and $\kappa_0 = 2$.

Figure A.1. shows the evolution of the inflation parameter and stochastic volatility. We then use the median estimate of the inflation parameter as z_t in the VAR part of the model.

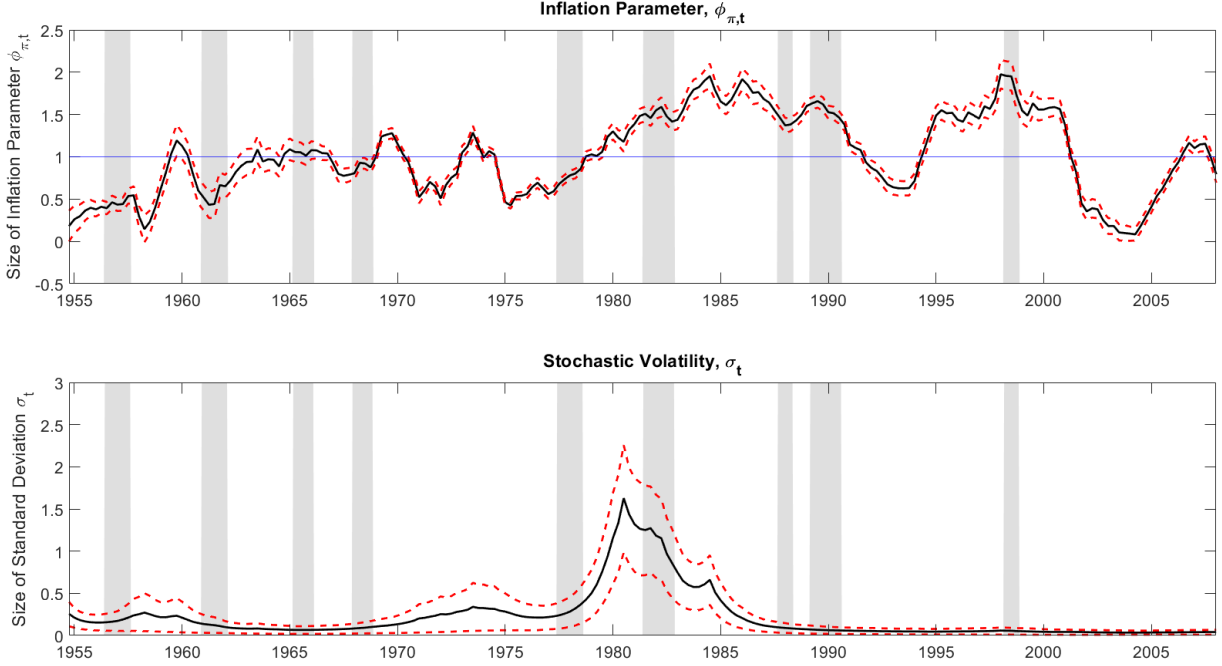
A.1.4 Kalman filter and smoother

In Sections A.1.1 and A.1.2, we used standard Kalman filtering methods to filter α^T and σ^T . This section provides a more detailed overview about these methods. Consider the following general state-space model

$$y_t = Z\alpha_t + e_t, e_t \sim N(0, H) \quad (\text{A.22.})$$

$$\alpha_t = T\alpha_{t-1} + R\eta_t, \eta_t \sim N(0, Q). \quad (\text{A.23.})$$

Figure A.1.: Inflation Parameter and Stochastic Volatility



Note: The figure shows the estimated sample paths of $\phi_{\pi,t}$ and σ_t . The black lines represent the pointwise-posterior median estimates and the red dotted lines correspond to the 68 percent credible bands. Both vary substantially during our sample period.

Kalman filtering techniques take Z, H, T, R, Q as given and provide an estimate of α_t via the following forecasting and updating steps

$$a_{t|t-1} = T a_{t-1} \quad (\text{A.24.})$$

$$P_{t|t-1} = T P_{t-1} T' + R Q R \quad (\text{A.25.})$$

$$a_t = a_{t|t-1} + P_{t|t-1} Z' F_t^{-1} (y_t - Z a_{t|t-1}) \quad (\text{A.26.})$$

$$P_t = P_{t|t-1} - P_{t|t-1} Z' F_t^{-1} Z P_{t|t-1} \quad (\text{A.27.})$$

where $F_t = Z P_{t|t-1} Z' + H$. The Kalman filter runs forward for $t = 1, \dots, T$. Then,

$$\alpha_T \sim N(a_T, P_T). \quad (\text{A.28.})$$

Conditional on α_T , the Kalman smoother runs backwards and samples α_t for $t = T - 1, T - 2, \dots, 0$ using the following steps

$$\alpha_{t|t+1} = \alpha_t + P_t T P_{t+1|t}^{-1} (\alpha_{t+1} - Z a_t) \quad (\text{A.29.})$$

$$P_{t|t+1} = P_t - P_t T P_{t+1|t}^{-1} T P_t \quad (\text{A.30.})$$

Then, $\alpha_t \sim N(\alpha_{t|t+1}, P_{t|t+1})$ for $t = T - 1, T - 2, \dots, 0$.

A.2 Estimation of the ST-VAR model

In this section, we describe the Bayesian estimation of the ST-VAR model. To sample the model parameters $\{\pi_{AM}, \pi_{PM}, \Omega_{AM}, \Omega_{PM}, c, \gamma\}$, we employ the following algorithm:

Algorithm 2 Multi-Move Gibbs Sampler for ST-VAR model

1. Initialize: Choose $\pi_{AM}^0, \pi_{PM}^0, \Omega_{AM}^0, \Omega_{PM}^0, \gamma^0, c^0$;
 2. Draw $\pi_{AM} | \Omega_{AM}, \Omega_{PM}, \gamma, c \sim N(m_{AM}, V_{AM})$ and $\pi_{PM} | \Omega_{AM}, \Omega_{PM}, \gamma, c \sim N(m_{PM}, V_{PM})$;
 3. Draw $\gamma | \pi_{AM}, \pi_{PM}, \Omega_{AM}, \Omega_{PM}, c$ using a Metropolis-Hastings step;
 4. Draw $c | \pi_{AM}, \pi_{PM}, \Omega_{AM}, \Omega_{PM}, \gamma$ from a Griddy Gibbs sampler;
 5. Draw $\Omega_{AM}, \Omega_{PM} | \pi_{AM}, \pi_{PM}, \gamma, c$ using a Metropolis-Hastings step;
 6. Repeat steps 2 through 5 and keep the desired number of draws after a burn-in phase.
-

A.2.1 Sample Π_E and Π_R

Define $\pi_r = \text{vec}(\Pi_r)$ for $r \in \{AM, PM\}$. For π_{AM} and π_{PM} , we choose the Minnesota prior, i.e.,

$$p(\pi_r) \sim N(m_0, V_0). \quad (\text{A.31.})$$

The Minnesota prior assumes that X_t follows a multivariate random walk process. π_0 is a vector equal to zero, except for the elements that correspond to a variable's own lags. In this way the Minnesota prior shrinks many parameters of π_r towards zero. In addition, V_0 has the following form

$$V(i, j)_0 = \begin{cases} \frac{\theta_1}{\ell^2} & \text{for parameter on own lags, } j = i \\ \frac{\theta_2 \sigma_{s,i}^2}{\ell^2 \sigma_{s,j}^2} & \text{for parameter on foreign lags, } j \neq i \end{cases} \quad (\text{A.32.})$$

where ℓ is the lag length and θ_1 and θ_2 are two hyperparameters. The Minnesota prior is a conjugate prior, which implies that the posterior distribution has the same form as the likelihood as the data. Therefore, the posterior of π_r is also normal, i.e.,

$$p(\pi_r | X^T, \Omega_r, \gamma, c) \sim N(m_r, V_r) \quad (\text{A.33.})$$

where

$$V_r = ((X'_{r,t-1} X_{r,t-1} \otimes \Omega_r^{-1}) + V_0^{-1})^{-1} \quad (\text{A.34.})$$

$$m_r = V_r \times ((X'_{r,t-1} X_{r,t-1} \otimes \Omega_r^{-1}) \hat{\pi}_r + V_0^{-1} \pi_0) \quad (\text{A.35.})$$

A.2.2 Sample γ

The posterior distribution of γ does not have a closed-form expression. Hence, we also rely on MCMC methods to approximate the marginal posterior distribution of γ . We follow Gefang and Strachan [2010] and Galvão and Owyang [2018] and use the Metropolis-Hasting algorithm to draw γ . Denote γ^* as a candidate draw from the proposal density $q(\gamma^*|\gamma^{k-1})$ where γ^{k-1} is the draw from the previous iteration. Then, we set $\gamma^k = \gamma^*$ with probability

$$\alpha(\gamma^*|\gamma^{k-1}) = \min\left\{\frac{p(\gamma^*|\cdot)/q(\gamma^*|\gamma^{k-1})}{p(\gamma^{k-1}|\cdot)/q(\gamma^{k-1}|\gamma^*)}, 1\right\} \quad (\text{A.36.})$$

$p(\gamma^*|\cdot)$ is the posterior distribution of γ^* while $p(\gamma^{k-1}|\cdot)$ is the posterior distribution of γ^{k-1} . Moreover, $q(\gamma^*|\gamma^{k-1})$ and $q(\gamma^{k-1}|\gamma^*)$ are the corresponding proposal densities. The proposal density is a gamma distribution. Under mild regularity conditions, the sequence of γ^k , after a sufficient burn-in phase, will approximate the true marginal posterior distribution of γ .

A.2.3 Sample c

The threshold parameter c does not have a closed-form for its posterior distribution. Again, we follow Gefang and Strachan [2010] and use the Griddy-Gibbs sampler proposed by Ritter and Tanner [1992] to draw c , where c is bounded between the minimum and maximum value of z_t . First, we select grid points to evaluate the density of z_t at these grid points and compute the corresponding (inverse) CDF. Second, we draw a random variable from a standard uniform distribution and plug it into the inverse CDF of z_t . This yields a draw from the marginal posterior distribution of c .¹⁹

A.2.4 Sample Ω_E and Ω_R

We assume that the purely *AM* and *PM* regimes are characterized not only by their own coefficient matrix Π_{AM} and Π_{PM} , but also by their own covariance matrix Ω_{AM} and Ω_{PM} . Under the assumption of heteroskedasticity, Ω_{AM} and Ω_{PM} are not conjugate. In addition, the joint posterior distribution of Ω_{AM} and Ω_{PM} cannot be written as the product of their marginal posterior distributions. Hence, they are also not independent and must be drawn jointly. We follow Galvão and Owyang [2018] and use the Metropolis-Hastings algorithm proposed by Chib and Greenberg [1995] to sample Ω_{AM} and Ω_{PM} . Denote Ω_{AM}^* and Ω_{PM}^* as a pair of candidate draws from the proposal densities $q(\Omega_{AM}^*|\Omega_{AM}^{k-1}, \Omega_{PM}^{k-1})$ and $q(\Omega_{PM}^*|\Omega_{AM}^{k-1}, \Omega_{PM}^{k-1})$, respectively. Ω_{AM}^{k-1} and Ω_{PM}^{k-1} are draws from the previous iteration. Then, we set $\{\Omega_{AM}^k, \Omega_{PM}^k\} = \{\Omega_{AM}^*, \Omega_{PM}^*\}$ with probability

$$\alpha(\Omega_{AM}^*, \Omega_{PM}^*|\Omega_{AM}^{k-1}, \Omega_{PM}^{k-1}) = \min\left\{\frac{p(\Omega_E^*, \Omega_R^*|\cdot)/q(\Omega_E^*, \Omega_R^*|\cdot)}{p(\Omega_E^{k-1}, \Omega_R^{k-1}|\cdot)/q(\Omega_E^{k-1}, \Omega_R^{k-1}|\cdot)}, 1\right\} \quad (\text{A.37.})$$

$p(\Omega_{AM}^*, \Omega_{PM}^*|\cdot)$ and $p(\Omega_{AM}^{k-1}, \Omega_{PM}^{k-1}|\cdot)$ are the corresponding posterior distributions of Ω_{AM}^* , Ω_{PM}^* and Ω_{AM}^{k-1} , Ω_{PM}^{k-1} . Moreover, $q(\Omega_{AM}^*, \Omega_{PM}^*|\cdot)$ and $q(\Omega_{AM}^{k-1}, \Omega_{PM}^{k-1}|\cdot)$ are their respective proposal

¹⁹This property is known as the inversion method. Suppose X is random variable with an unknown distribution and u is standard uniformly distributed. Then $F^{-1}(u)$ can be used to generate draws of X with the specified CDF F .

densities. The proposal density for both Ω_{AM} and Ω_{PM} is a inverse-Wishart distribution. Under relative mild regularity conditions, the sequences of $\{\Omega_{AM}^k, \Omega_{PM}^k\}$, after a sufficient burn-in phase, will approximate the true posterior distribution of Ω_{AM} and Ω_{PM} .

A.3 Generalized Impulse Response Functions

In nonlinear models, the impulse response parameters depend on the shock sign, the shock size and the timing of the shock. Furthermore, the state of the economy can change over time and after the shock. To incorporate these features, we follow Koop et al. [1996] and use generalized impulse response functions to estimate the dynamic effects of the shock. The generalized impulse responses defined as the difference of two simulated paths of the economies, i.e.,

$$\begin{aligned} GIRF(h) = & (1 - G(z_{t+h-1})) \Pi_{AM} X_{t+h-1} + G(z_{t+h-1}) \Pi_{PM} X_{t+h-1} + u_{t+h}^\epsilon \\ & - (1 - G(z_{t+h-1})) \Pi_{AM} X_{t+h-1} + G(z_{t+h-1}) \Pi_{PM} X_{t+h-1} + u_{t+h} \end{aligned} \quad (\text{A.38.})$$

First, we draw a set of reduced form parameters $\Omega_{AM}, \Omega_{PM}, \Pi_{AM}$ and Π_{PM} . Second, the generalized impulse response functions require an initial condition. Hence, we draw an initial condition that comes with the lags for the VAR X_{t-1} , the value of the state variable z_{t-1} and a sequence of the reduced-form residuals of H periods u_t^H . Third, we transform the sequence of reduced form residuals into structural shocks and “back”

$$\epsilon_{t+h} = (Chol(\Omega_t)Q)^{-1} u_{t+h} \quad (\text{A.39.})$$

$$\tilde{\epsilon}_{t+h} = \epsilon_{t+h} + \delta \quad (\text{A.40.})$$

$$u_{t+h}^\epsilon = Chol(\Omega_t)Q\tilde{\epsilon}_{t+h} \quad (\text{A.41.})$$

where δ represents the shock size, $Chol(\Omega_t)$ is the Cholesky decomposition of Ω_t and Q is an orthogonal matrix, called *rotation* matrix. We follow Rubio-Ramirez et al. [2010] and draw Q from the Haar measure via the QR decomposition of a matrix of standard normal random variables. Fourth, conditional on the starting period, we roll the model forward using Equation (1) of our main model and add the value of u_{t+h} to the path of the economy without the government spending shock and u_{t+h}^ϵ to the path of the economy with the government spending shock. Fifth, we take the difference between the simulated economies to obtain one candidate draw of the generalized impulse response functions. If the candidate satisfies the sign restrictions, we store the draw. If not, we discard the candidate, draw a new rotation matrix and check again if the imposed sign restrictions are satisfied. We repeat this procedure until a sufficient number of candidates are accepted and move to the next initial condition. Then, we average over the accepted candidates and initial conditions to obtain one realization of the generalized impulse response functions. Finally, we repeat the whole procedure for the next draw of the reduced form parameters, $\Omega_{AM}, \Omega_{PM}, \Pi_{AM}$ and Π_{PM} . This procedure yields a distribution of generalized impulse response functions that are consistent with our sign restrictions.

B Taylor Rules

This section discusses some of the issues related to the estimation of Taylor rules. This is an important section because the inflation parameter of the Taylor is used in our main model to distinguish between different monetary policy regimes. The estimation of Taylor rules has spurred a large amount of research that needs to be reviewed in the context of our study.

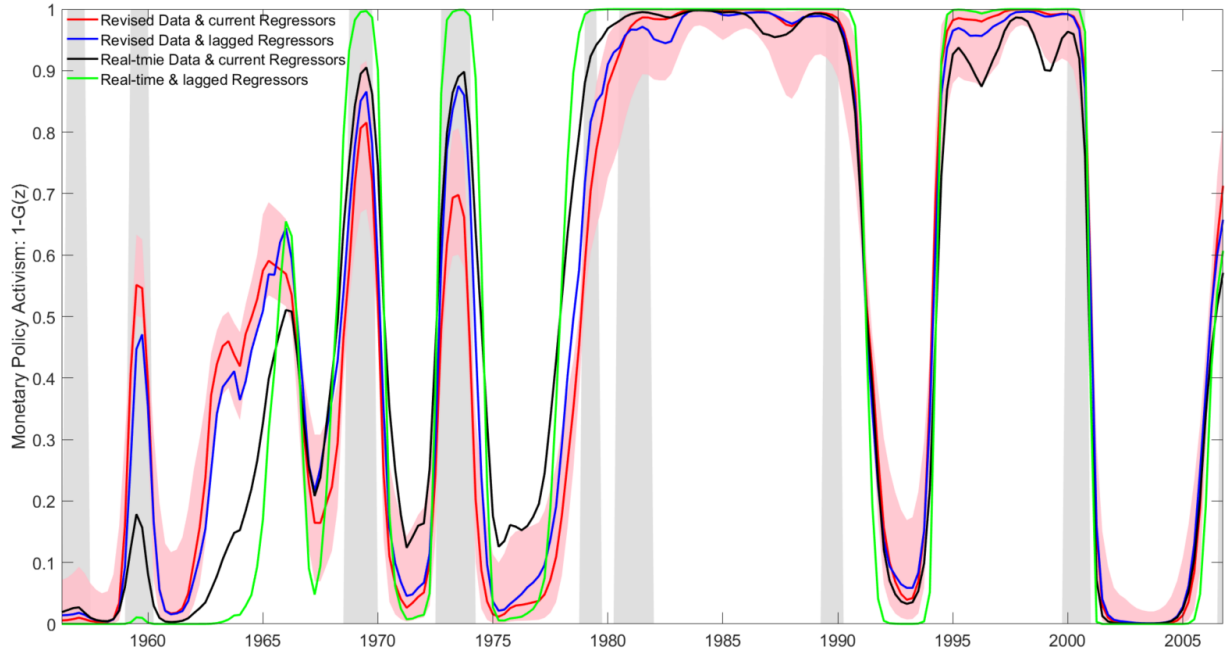
Taylor [1999] estimates a simple Taylor rule by regressing the federal funds rate as the monetary policy instrument on real output and the inflation rate using least squares. Clarida et al. [2000] also estimates a simple Taylor rule. However, they exploit a large set of instruments and employ general methods of moments. Both studies find that monetary policy changed considerably before and after 1980. However, both studies use revised data. Orphanides [2001, 2002, 2004] and Boivin [2006] suggest real time data should be used because revised data may contain information that were unavailable to policy makers during the time they faced policy decisions. These studies find that monetary policy may also have been active in the 1970s, which contradicts the conclusion drawn by Taylor [1999] and Clarida et al. [2000].

Moreover, Cogley and Sargent [2001] find that monetary policy has varied substantially over time as well. Sims [2001] argues that the change in monetary policy is due to time-varying heteroskedasticity in the Taylor-rule residuals. However, Cogley and Sargent [2005] demonstrate that their 2001-results are robust when after accounting for stochastic volatility in their model.

de Vries and Li [2013] illustrate that the Taylor rule residuals are serially correlated. In this case, instruments such as lagged regressors, real time data or the Clarida et al. [2000]'s instruments are not necessarily valid instruments. However, the resulting endogeneity problem does not cause a substantial bias. Finally, Carvalho et al. [2019] provide evidence that the bias of a Taylor rule estimated with least squares is small.

To sum up, the estimation of Taylor rules has spurred extensive research. The key question is whether monetary policy has changed over time. According to the literature, the choice of the data (revised vs. real time data) and assumptions regarding the residuals (constant vs. stochastic volatility) might be important. Recent evidence, however, suggests that differences in the estimators are small. In this section, we estimate the Taylor rule in four different specifications: (i) revised data with current inflation and real GDP growth rates (main specification used in the paper), (ii) revised data with lagged inflation and real GDP growth rates, (iii) real-time data with current inflation and real GDP growth rates, and (iv) real-time data data with lagged inflation and real GDP growth rates. Figure B.1. presents the corresponding evolutions of $1 - G(z_{t-1})$.

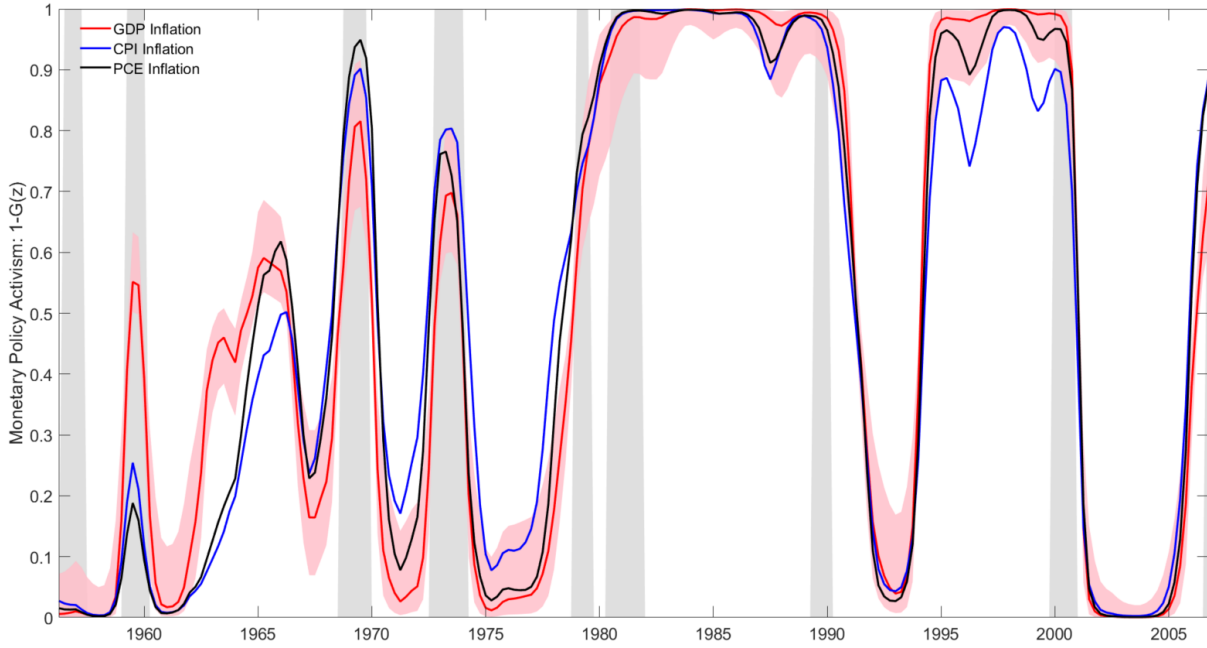
Figure B.1.: Evolution of Monetary Policy: Revised vs. Real-Time Data



Note: The Figure compares the estimated evolutions of $1 - G(z_{t-1})$ using current and lagged regressors from revised and real-time data, respectively. The results indicate that if differences exist, they are small and appear in the pre-Volcker period.

Figure B.1. presents the evolution of $1 - G(z_{t-1})$ rather than the evolution of the inflation parameter because $1 - G(z_{t-1})$ determines the dynamics of our model in the main text while the inflation parameter is just an ingredient of $G(z_{t-1})$. Figure B.1. shows that there exist differences between revised vs real time data and between current and lagged regressors. However, these differences are small and mostly arise in the pre-Volcker era. For example, not all estimations “recognize” the Martin disinflation in 1959/60. However, all estimated evolutions capture the change in monetary policy around 1980, the responses to the 1957/58, 1990/91 and 2000/01 recessions and the substantial deviation from the Taylor principle between 2002 and 2005.

Figure B.2.: Evolution of Monetary Policy: different Price Indices



Note: The Figure compares the estimated evolutions of $1 - G(z_{t-1})$ using different price indices to compute inflation. The results indicate that if differences exist, they are small and appear mostly in the pre-Volcker period and in the second-half of the 1990s.

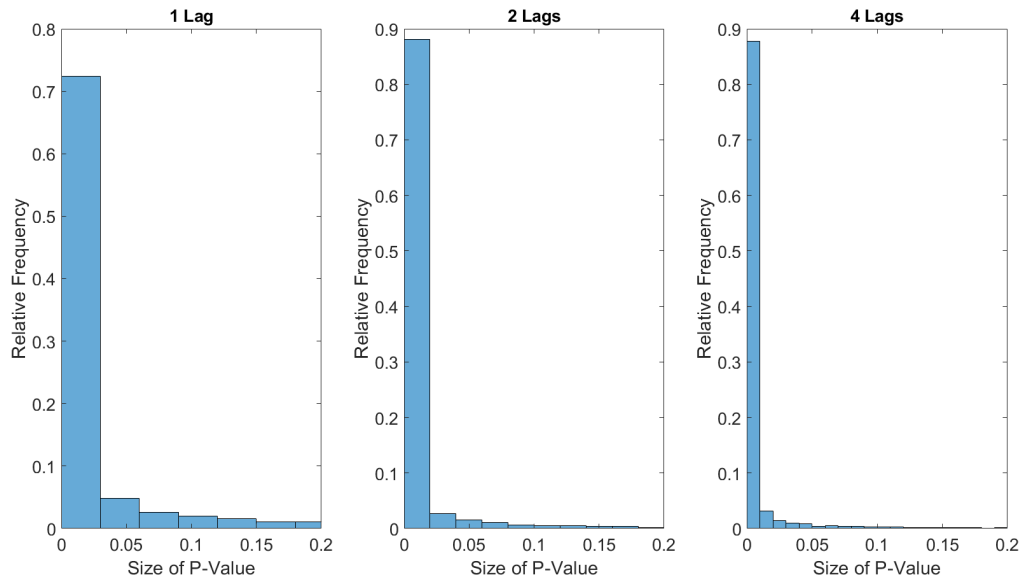
Another way to test the robustness of our results is to use different price indices to compute inflation. In Figure B.2., we show that $1 - G(z_{t-1})$ displays a similar path regardless of whether we compute inflation based on (i) the GDP deflator, (ii) CPI or (iii) PCE. As in Figure B.1., if differences exist, they appear in the pre-Volcker era. For example, the Martin disinflation is not as pronounced for CPI or PCE as it is for the GDP deflator. In contrast, all three specifications display the changes in monetary policy in 1957/58, around 1980, 1990/91, 2000/01 or between 2002-2005.

A possible explanation for the similar estimates of $1 - G(z_{t-1})$, regardless of whether we use (i) current or lagged regressors, (ii) revised or real-time data or (iii) different price indices to compute inflation, is serial correlation in the residuals of the different Taylor rule specifications. To investigate this feature of the residuals, we conduct the Ljung-Box Q-Test.

We use Bayesian methods to estimate the different Taylor rules. Hence, we can also compute the posterior distribution of the residuals. We then conduct the Ljung-Box Q-Test on each draw of the residuals. The null of the Ljung-Box Q-Test is that the residuals are not serially correlated. In Figure B.3., we show the distributions of p-values of this test with one, two and five lags for our baseline specification. Figure B.3. provides strong evidence for serial correlation in the residuals. For example, the share of p-values that is below 0.05 for the test

with one lag is above 70 percent. This implies that more than 70 percent of the p-values are below 0.05. For the tests with two and four lags, the share of small p-values increases to above 90 percent. This test for our baseline specification provides strong evidence that the residuals are indeed serially correlated. According to de Vries and Li [2013], if the residuals of the Taylor rule are serially correlated, the estimation using standard instruments such as lagged regressors or the Clarida et al. [2000]’s instruments face endogeneity problems, but the resulting bias is small.

Figure B.3.: Ljung-Box Q-Test



Note: The figure shows the distribution of p-values for a Ljung-Box Q-Test. The Figure provides evidence that the residuals of the estimated Taylor rule exhibit serial correlation. For example over 70 percent of the p-values for the Test with one lag is below 0.05. For the tests with two and four lags, the percentage is larger.

C Additional Results

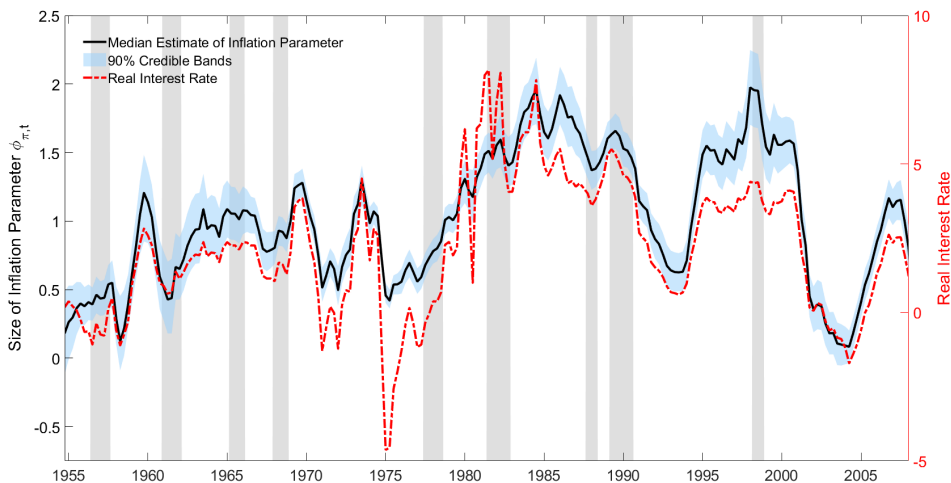
This appendix provides supplemental results. We (i) elaborate on the relationship between the real interest rate and our estimated evolution of monetary policy, (ii) include a different measure for agents' expectations regarding future government spending plans, (iii) examine the consequences of the binary interpretation of monetary policy, (iv) compare different updating rules for the state variable after the government spending shock, and (v) consider different shock signs and different shock sizes.

C.1 Monetary Policy and the real interest rate

When we study the evolution of monetary policy, we face the difficulty that we do not have a natural comparison. For example, Auerbach and Gorodnichenko [2012] show that their estimated evolution of the business cycle is well aligned with the NBER recessions. For monetary policy, there does not exist such a natural benchmark. However, when theorists make their predictions about whether the government spending multiplier also depends on monetary policy, they also refer to the real interest rate. Hence, in Figures C.1. and C.2., we compare the estimated evolution of the inflation parameter $\phi_{\pi,t}$ and the transition function $1 - G(z_{t-1})$ with the real interest rate defined as the difference between the federal funds rate and annualized inflation based on the GDP deflator.

Figure C.1. illustrates that the estimated sample path of $\phi_{\pi,t}$ and the sample path of the real interest rate align well. During times of a responsive central bank (high $\phi_{\pi,t}$), real interest rates are high and vice versa. In addition, changes in $\phi_{\pi,t}$ and changes in the real interest rate occur at almost the same time. For example, in 1959, $\phi_{\pi,t}$ increases from almost zero to above one. The real interest rate change exactly at the same. Moreover, between 1975 and 1980, $\phi_{\pi,t}$ increases from around 0.5 to almost two. At the same time, the real interest rate displays its largest increase over our sample period. A similar relationship is demonstrated in Figure C.2.. When monetary policy was very passive e.g., during the late 1950s, second half of 1970s or the first half of the 2000, real interest rates were relatively low. In contrast, when monetary policy was very active e.g., during the 1980s or the second half of the 1990s, real interest rates were relatively high. Finally, when monetary policy changed and becomes more active, real interest rates increase and vice versa if monetary policy becomes more passive.

Figure C.1.: The Sample Paths of the estimated Inflation Parameter and the Real Interest Rate



Note: The red line shows the sample path of the real interest rate and the black line presents the pointwise-posterior median of the inflation parameter, $\phi_{\pi,t}$, along with the 68 percent credible bands in blue. The Figure establishes a close relationship between the two.

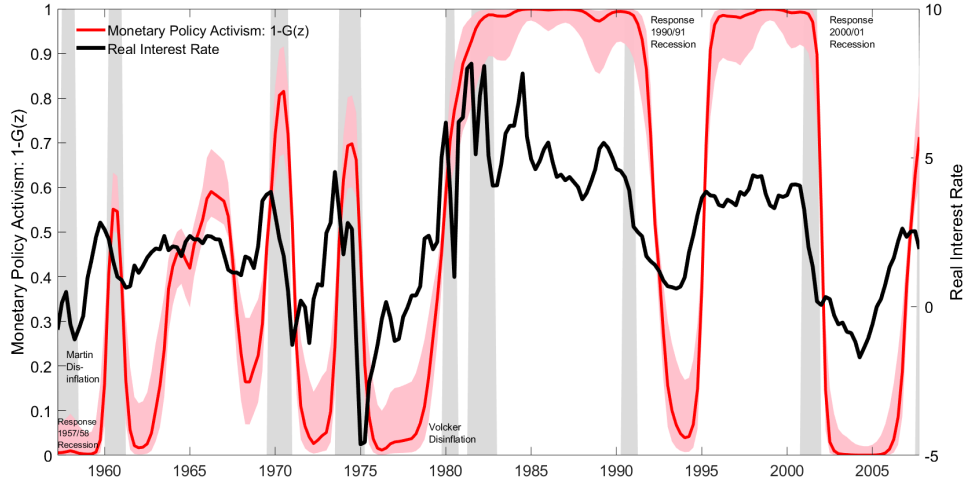
C.2 Alternative measures for agent’s expectations regarding future government spending plans

In our baseline specification, we augment our information set with Ramey’s news shocks to account for agent’s expectations regarding future government spending plans. This approach is one way to account for fiscal foresight. Fiscal foresight is an example of the *non-fundamentalness* problem that arises because the econometrician cannot observe all relevant information to recover structural shocks. In this case, a VMA model of the economy does not have a VAR representation, so a VAR cannot consistently estimate impulse response functions.²⁰

Ramey [2011] argues that her news shock series is not very informative for the post-Korean war period. Despite this concern, we include Ramey’s news shocks in our baseline specification as it has remained the standard expectation variable in the literature. In this section, we conduct a robustness check in which we re-estimate our replication exercise from Section 4 but replace Ramey’s news shocks with the shock series from Ben Zeev and Pappa [2017]. Ben Zeev and Pappa [2017] define a government spending shock as the shock that “best explains future movements in defense spending over a five-year horizon and is

²⁰See [Lippi and Reichlin, 1993, Leeper et al., 2013, Ramey, 2011]

Figure C.2.: The Sample Paths of $1 - G(z_{t-1})$ and the Real Interest Rate



Note: The red line shows the median estimate of $1 - G(z_{t-1})$ along with the 68 percent credible bands in pink. The black line corresponds to the real interest rate. The Figures demonstrates that during times of passive monetary policy, the real interest rate is relatively low, and relatively high during times of active monetary policy.

orthogonal to current defense spending”.

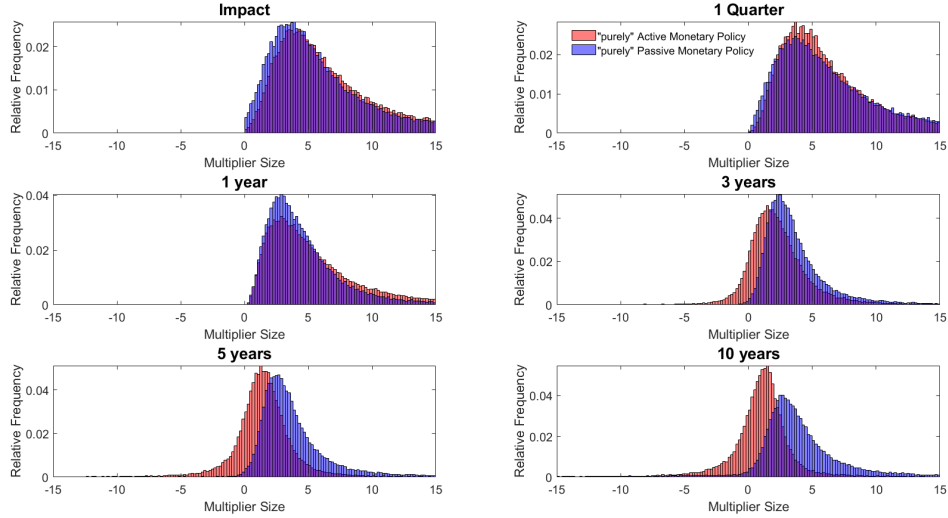
Figure C.3. shows that the results are very similar to our baseline specification in Figure 7. The multiplier estimates are comparable across monetary policy regimes up to one year after the shock. One year after the shock, the multiplier starts to diverge and is estimated to be higher in the long-run. These results mirror those in Leeper et al. [2017], who find that the multiplier is similar across monetary policy regimes in the short-run but diverges in the long-run.

C.3 Consequences of the binary interpretation of monetary policy

We now demonstrate the consequences of the binary interpretation of monetary policy. We repeat our main exercise but only distinguish between active and passive monetary policy. Using the generalized impulse response functions, we must employ a threshold value: If $G(z_{t-1}) \geq 0.5$, then the monetary policy regime in t is declared as “passive“ and as “active“ otherwise. We then estimate the generalized impulse response function for both regime drawing random initial conditions from the two subsets.

Figure C.4. shows the response of the monetary policy regime to the government spending shock. If monetary policy is initially “active”, the central bank changes its policy regime only slightly. In contrast, if the initial regime is “passive”, the central bank responds quickly and

Figure C.3.: Multiplier Estimates Ben Zeev and Pappa [2017] Shocks

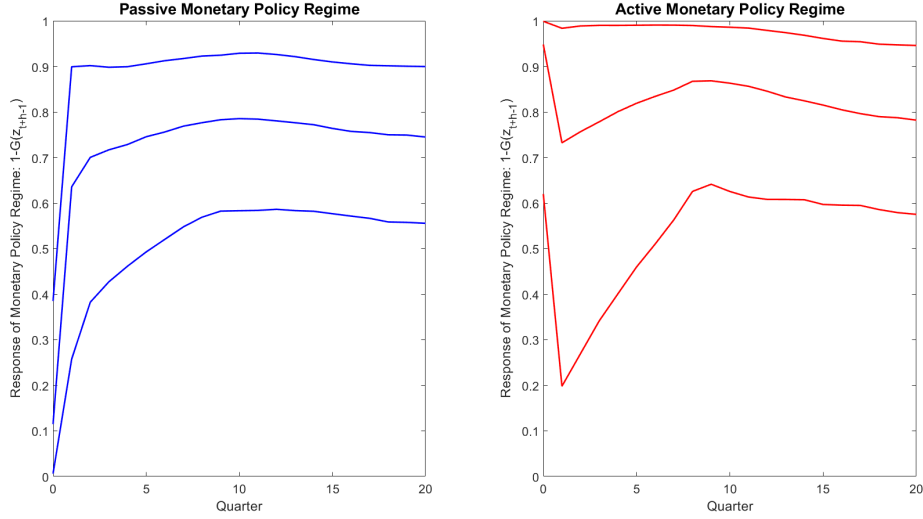


Note: The figure compares the estimated multipliers when monetary policy is and remains “purely active” (red) and “purely passive” (blue) after the shock. In this Figure, we replace the Ramey’s news shocks by the Ben Zeev and Pappa [2017] shocks to account for fiscal foresight. The result show that the multipliers diverge when monetary policy is and remains purely passive. This result is similar to that from Figure 7.

transitions fast to the “active” regime. Shortly after the shock, the central bank responds aggressively to inflation regardless of the initial regime. Figure C.5. illustrates that the government spending multiplier does not depend on the initial monetary policy regime – either in the short-run or in the long-run. These results mirror our main findings.

In section 3, we also find that monetary policy is not just active or passive but differences exists even within these broader regime in terms of how active or passive the monetary policy regime can be. In Figure C.6., we demonstrate the consequences of the binary interpretation of monetary policy when the regime is in fact continuous. Figure C.6. plots the distribution of $G(z_{t-1})$ to display the distribution of initial monetary policy regimes. There are two spikes close to zero and one. However, there are also many lying in the interior of the unit interval, many of them lie close to the threshold value. This casts doubt on the idea that these “active” and “passive” GIRFs are representative of the ends of the spectrum. Rather, on average, we compare initial monetary policy regimes of 0:14 and 0:84 to each other. These values are far from the monetary policy regimes of zero and one, but also incorporate many observations in the neighborhood of .5. The binary approach to GIRFs could fail to identify interesting variation. For example, Caggiano et al. [2015] and Ramey and Zubairy [2018] compare spending multipliers between expansions and recessions relying on the binary interpretation of the business cycle. Both studies conclude that there are

Figure C.4.: Response of Monetary Policy Regime when Monetary Policy is Binary



Note: The figure shows the pointwise-posterior median evolution of $1 - G(z)$, along with the 68 percent credible bands, when monetary policy is interpreted as binary: just active (red) and just passive (blue). The figure suggests that shortly after the shock and regardless of its initial condition, the central bank responds actively to inflation.

no significant differences in the estimated multipliers. However, results thus obtained are difficult to interpret because they could come from two sources: (i) multipliers that are indeed regime-independent; or (ii) underlying regimes that are not sufficiently different.

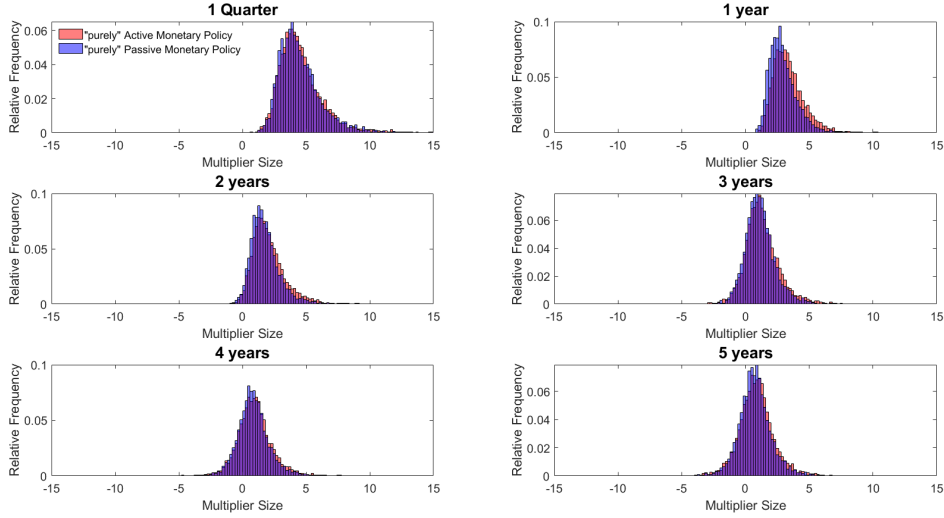
C.4 Alternative Updating Rules for z_t

In the main paper, we use the Kalman filter based on the forecasted values of the federal funds rate, inflation and output growth to forecast the inflation parameter as the state variable. Using the new value of the state variable, we update the weights $G(z_{t+h})$ and $1 - G(z_{t+h})$ on the purely passive and purely active monetary policy regimes. In this way, the government spending shock affects monetary policy, and changing monetary policy further impacts the transition of the government spending shock. However, other rules to update the inflation parameter of the Taylor rule are also conceivable. For example, we can run a regression based on forecasted values of inflation on the forecasted values of inflation and output growth, i.e.

$$\hat{i}_{t+h} = \hat{c}_{t+h} + \hat{\phi}_{\pi,t+h}\hat{\pi}_{t+h} + \hat{\phi}_{y,t+h}\hat{y}_{t+h} + \hat{\epsilon}_{mp,t}. \quad (\text{C.1.})$$

Then, we can use the point estimate of $\hat{\phi}_{\pi,t+h}$ as z_{t+h} to update the weights $G(z_{t+h})$ and $1 - G(z_{t+h})$. Alternatively, we can forecast the inflation parameter $\phi_{\pi,t+h}$ using a AR(1)

Figure C.5.: Multiplier Estimates when monetary policy is responsive but binary



Note: The figure displays the multiplier estimates for the scenario when the monetary policy regime is responsive but binary: just active (red) or just passive (blue). Regardless of the forecast horizon after the shock, there exists little evidence that the government spending multiplier depends on monetary policy.

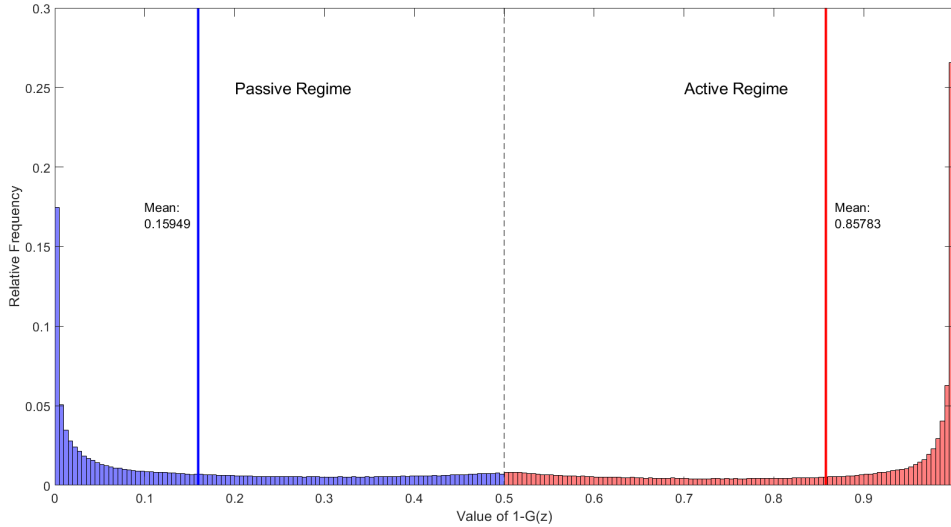
process with an exogenously given persistence parameter ρ , i.e.,

$$\phi_{\pi,t+h} = \rho\phi_{\pi,t+h-1}. \quad (\text{C.2.})$$

Finally, we can also include z_t in X_t , treat it as a “variable” and use the forecasts of z_{t+h} directly to update the weights $G(z_{t+h})$ and $1 - G(z_{t+h})$.

Figures C.7., C.8. and C.9. compare the responses of the monetary policy regime to the government spending shock. If we use the updating regression to forecast $\phi_{\pi,t+h}$ after the government spending shock (Figure C.7.), then regardless of its initial regime, the monetary policy regime reacts quickly and converges towards the sample mean (0.49). Ten quarters after the shock, the monetary policy regime is the similar regardless of its initial condition. In contrast, if we use an AR(1) process and set the persistence parameter equal to 0.95 (Figure C.8.), the monetary policy regime also responds after the shock, but only if the initial regime is “very active” or “very passive”. In addition, the response is much slower than that using the updating regression. Even 10 years after the shock, the monetary policy regime still has not fully converged. Finally, if we include z_t in X_t and treat it as a normal “variable”, the central bank also responds quickly but the behavior is somewhat different compared to Figures 2 and C.7. For the first ten quarters, the monetary policy regime converges towards the “weakly active” regime. Afterwards, the central bank changes it policy regime again: if

Figure C.6.: Distribution of initial Monetary Policy regimes



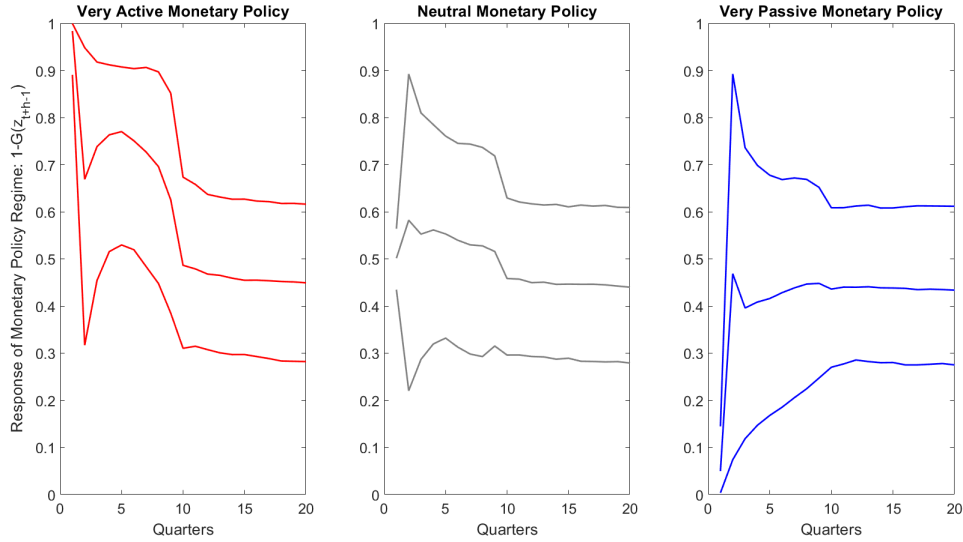
Note: Figure shows the distribution of the initial “passive” monetary policy regimes (blue) and the initial “active” monetary policy regimes (red). Bisecting the sample into “active” and “passive” regimes based on the initial value of $1 - G(z)$ lumps many of the central observations into one regime or the other. However, many initial monetary policy regime are close to the threshold value of 0.5 and the “distinct” regimes are relatively close to each other.

the regime was initially “very active”, the central bank tends to become more active, while if the initial regime was “very passive”, the policy regime becomes more passive.

Now, we compare multiplier estimates across the initial monetary policy regime using boxplots using the updating regression and the AR(1) process in Figures C.10. and C.11. as well as including z_t in X_t in Figure C.12.. In Figure C.10., the government spending multiplier does not depend on the monetary policy regime, either in the short- or long-run. This result is in line with our main conclusion that the government spending multiplier does not depend on the monetary policy regime once we account for the subsequent response of the policy regime. In contrast, if we apply the exogenous AR(1) process to update the inflation parameter (Figure C.11.), the government spending multiplier differs across initial monetary policy regimes and is estimated to be higher if monetary policy is initially “very passive”. Lastly, when we include z_t in X_t in Figure C.12., the multiplier estimates fluctuate around five one quarter and one year after the shock and decrease to around one five years after the shock. As in Figures 3 and C.10., the multiplier does not seem to depend on the initial monetary policy regime.

The findings of this exercise using alternative updating rules for $\phi_{\pi,t+h}$ is in line with the results from our main and counterfactual exercises: if the monetary policy regime is different

Figure C.7.: Response of Monetary Policy Regime using an Updating Regression



Note: The figure shows the pointwise-posterior median evolution of $1 - G(z)$, along with the 68 percent credible bands, when monetary policy is initially “very active,” “neutral,” and “very passive”. To update z_t , we use the regression in Equation (C.1). Regardless of the initial regime, monetary policy converges quickly to the “neutral” policy regime.

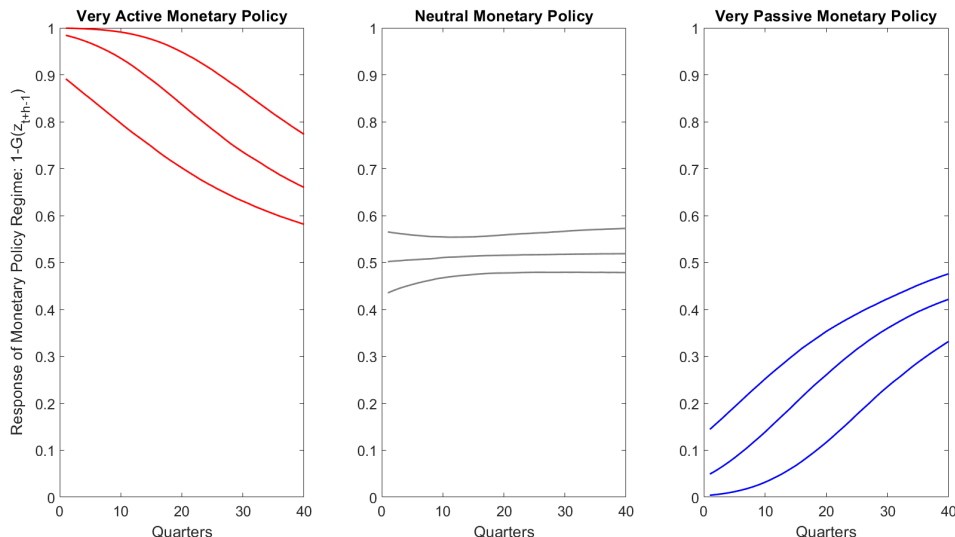
for a sufficiently long period of time, then the government spending multiplier depends on monetary policy as the conventional wisdom suggests. However, because the central bank typically responds quickly after the government spending shock and transitions fast to a similar regime regardless of its initial condition, we can infer that it is the constant-regime assumption that drives this consensus.

C.5 Different Shock Sizes and different Shock Signs

In this last section, we exploit a specific characteristic of nonlinear time series models such as ours. In nonlinear time series models, impulse response functions may depend on the shock sign or the shock size. Therefore, we now look at whether the government spending multipliers depends on monetary policy regimes if we vary the shock size and shock sign. First, Figures C.13. and C.14. compare the multiplier estimates across monetary policy regime when the government spending shock is large (two standard deviations) and small (half a standard deviation), respectively.

We find little evidence that the government spending multiplier depends on the shock size. Regardless of the shock size, the multiplier estimate is similar to our main exercise: the posterior median of the multiplier estimate decreases over time from above five after

Figure C.8.: Response of Monetary Policy Regime using an Updating AR(1) process with an Exogenous Persistence Parameter



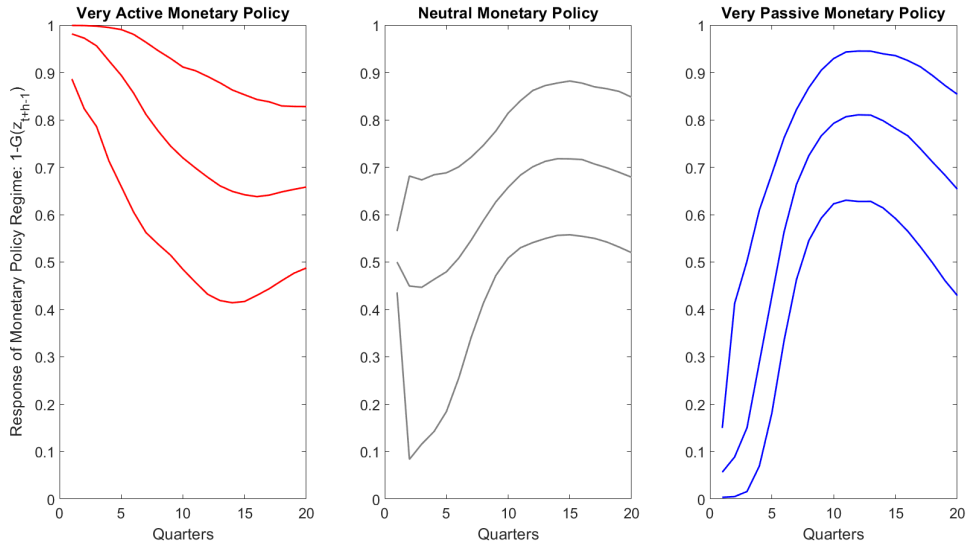
Note: The figure shows the pointwise-posterior median evolution of $1 - G(z)$, along with the 68 percent credible bands, when monetary policy is initially “very active,” “neutral,” and “very passive”. To update z_t , we use an updating AR(1) process with an exogenous persistence parameter set to 0.95 from Equation (C.2). Convergence in the policy regime is slow.

one quarter to around one after five years. We do not find any particular deviation in the response of the monetary policy regime after the shock to that from our main exercise shown in Section 4.²¹

Finally, we present the results of a contradictory government spending shock of size of one standard deviation. Figure C.15. shows that following a contractionary government spending shock, the central bank responds quickly. The response of the monetary policy regime after the shock is different from that in our main exercise. Here, the monetary policy regime does not converge to a “very active” policy but to a “weakly active” - “neutral” policy. If the monetary policy regime after five years is less active than in our main exercise and the conventional wisdom in the literature exhibit some truth about the multiplier, we should expect that the magnitude of the multiplier after five years is larger than in our main exercise, even if it does not depend on the initial monetary policy regime within the five-year period. However, Figure C.16. illustrates that our estimates do not support this claim. The posterior median after five years is around unity regardless of whether we consider an expansionary shock (leading to a “very active” policy after five years) or a contractionary shock (with a “weakly active” - “neutral” regime after five years).

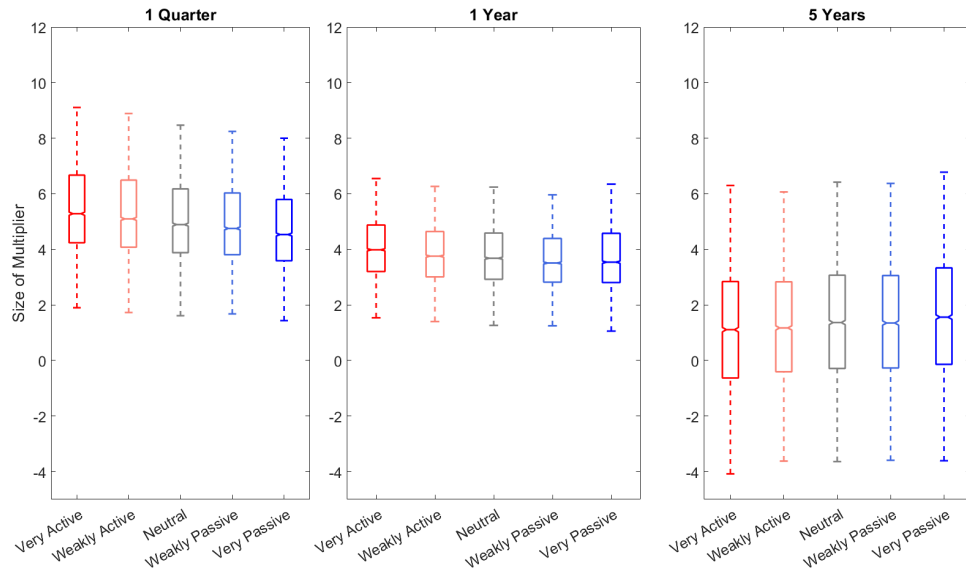
²¹Results are available upon request.

Figure C.9.: Response of Monetary Policy Regime when X_t includes z_t



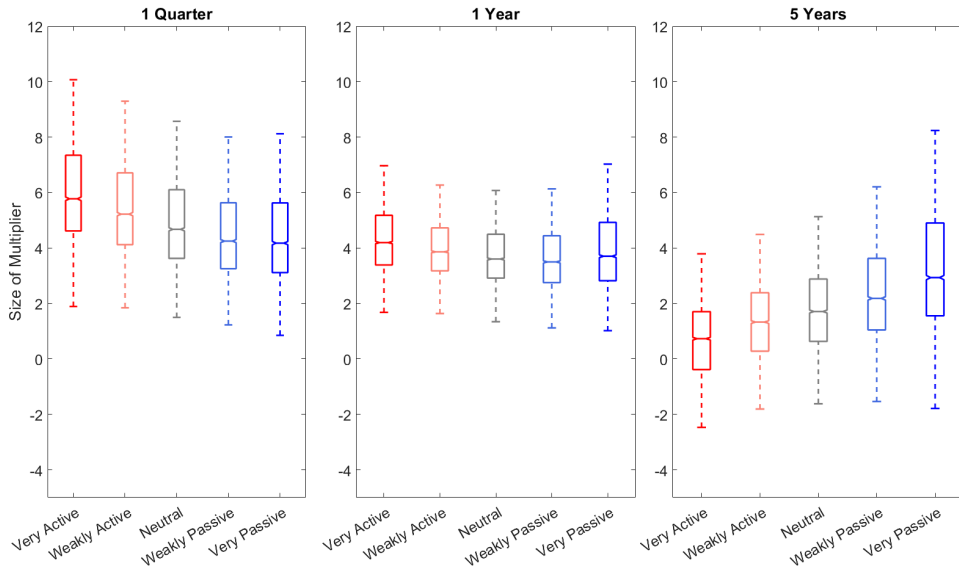
Note: The figure shows the pointwise-posterior median evolution of $1 - G(z)$, along with the 68 percent credible bands, when monetary policy is initially “very active,” “neutral,” and “very passive”. To update z_t , we include z_t in X_t . The monetary policy regime is converging in the short-run but starts to diverge again ten quarters after the shock.

Figure C.10.: Estimated Multipliers using an updating regression



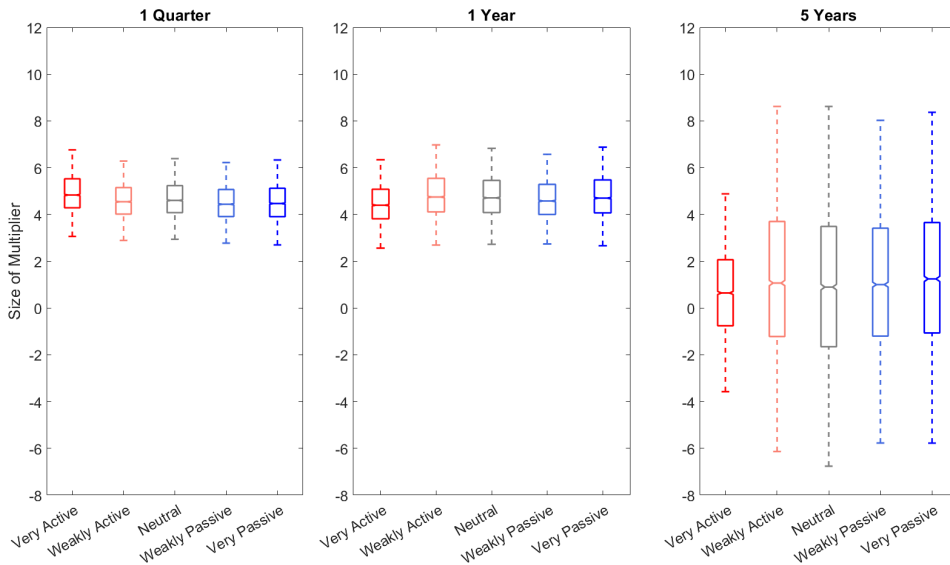
Note: The figure shows the estimated multiplier distributions one quarter, one year, and five years after the shock across initial monetary policy regimes using boxplots, and the regression in Equation (C.1.) to update the monetary policy regime. Estimated multipliers decrease with time but are almost completely unaffected by the initial policy regime.

Figure C.11.: Estimated Multipliers using an updating AR(1) process



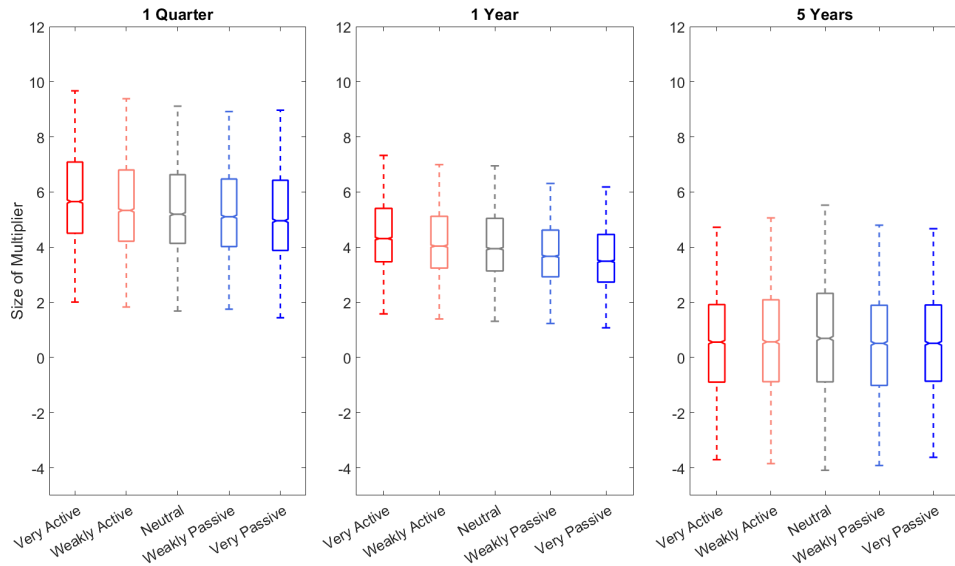
Note: The figure shows the estimated multiplier distributions one quarter, one year, and five years after the shock across initial monetary policy regimes using boxplots, and the AR(1) process in Equation (C.2.) to update the monetary policy regime. Estimated multipliers decrease with time and depend on the monetary policy regime after five years. This is the result of the slow convergence in Figure C.8..

Figure C.12.: Estimated Multipliers when X_t includes z_t



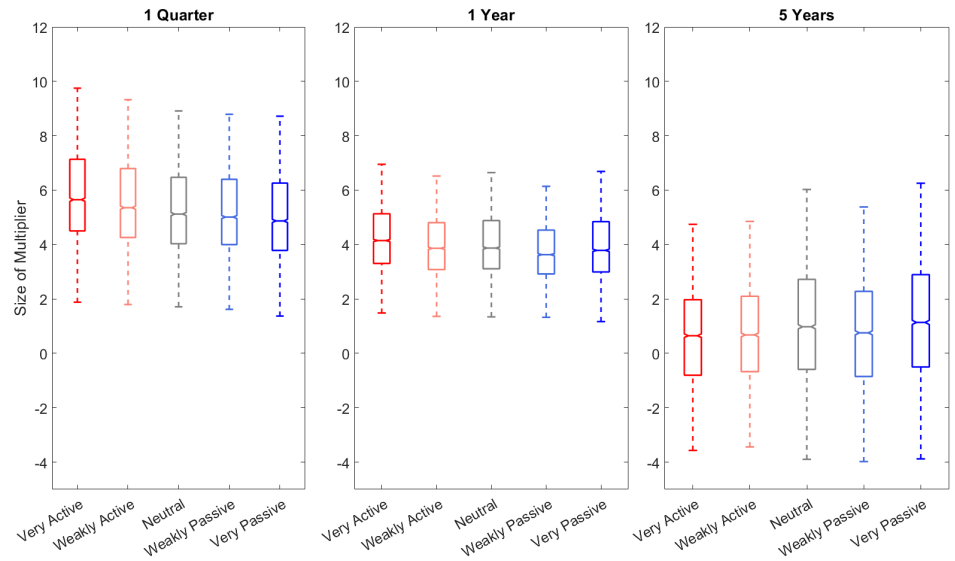
Note: The figure shows the estimated multiplier distributions one quarter, one year, and five years after the shock across initial monetary policy regimes using boxplots when z_t is included in X_t . Estimated multipliers decrease with time but do not depend on the initial monetary policy regime.

Figure C.13.: Estimated Multipliers for a large Government Spending Shock



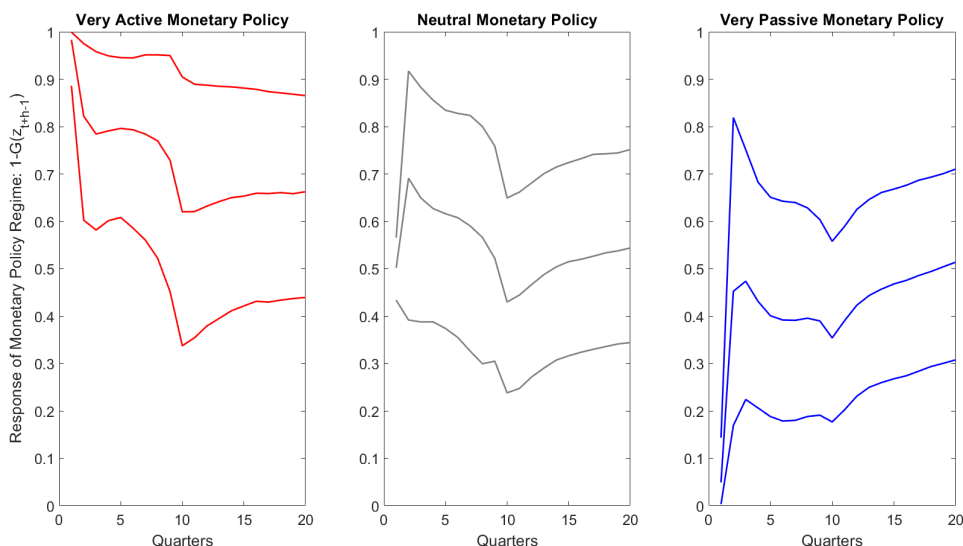
Note: The figure shows the estimated multiplier distributions one quarter, one year, and five years after a *large* government spending shock (two standard deviations) across initial monetary policy regimes using boxplots. To update the monetary policy regime, we use the “rolling” Kalman filter. Estimated multipliers decrease with time but do not depend on the initial policy regime.

Figure C.14.: Estimated Multipliers for a small Government Spending Shock



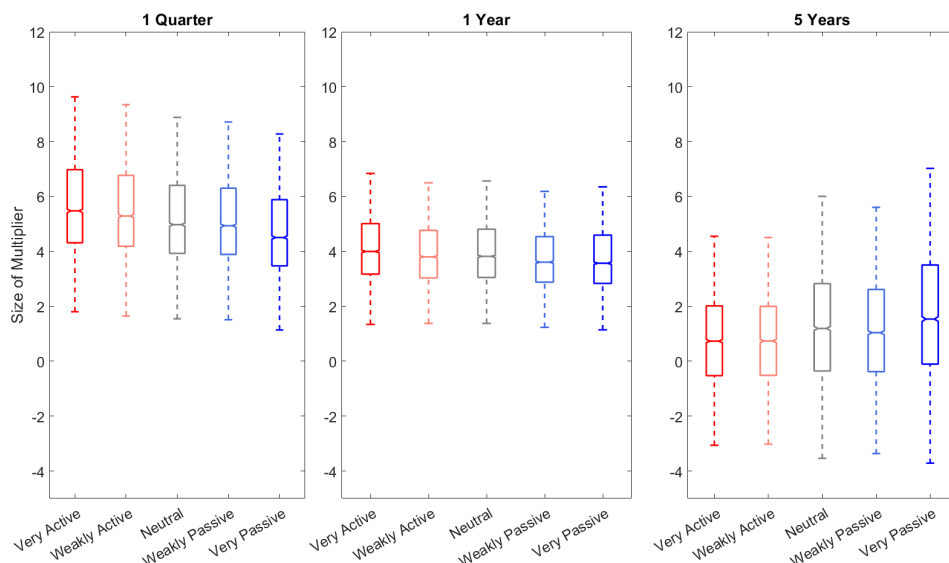
Note: The figure shows the estimated multiplier distributions one quarter, one year, and five years after a *small* government spending shock (0.5 standard deviations) across initial monetary policy regimes using boxplots. To update the monetary policy regime, we use the “rolling” Kalman filter. Estimated multipliers decrease with time but do not depend on the initial policy regime.

Figure C.15.: Response of Monetary Policy Regime for a contractionary Government Spending Shock



Note: The figure shows the pointwise-posterior median evolution of $1 - G(z)$, along with the 68 percent credible bands, when monetary policy is initially “very active,” “neutral,” and “very passive” for a contractionary government spending shock. Regardless of the initial regime, monetary policy converges to a “weakly active” or a “neutral” regime.

Figure C.16.: Estimated Multipliers for a contractionary Government Spending Shock



Note: The figure shows the estimated multiplier distributions one quarter, one year, and five years after a *contractionary* government spending shock across initial monetary policy regimes using boxplots. To update the monetary policy regime, we use the “rolling” Kalman filter. Estimated multipliers decrease with time but do not depend on the initial policy regime.

D A narrative History of Monetary Policy

In this Section, we provide a narrative perspective about the history of monetary policy since 1954. The key goal is to develop a better understanding of why monetary policy has changed over time and to support our empirical estimation of monetary policy in Figure 1. The section follows closely Romer and Romer [2004] and Romer and Romer [2013]. The authors analyze government reports such as Minutes and Transcripts of Federal Open Market Committee, Annual Reports of the Board of Governors of the Federal Reserve System and the Congressional testimony of the Fed's chairman. They conclude that changes in monetary policy mostly occurred because the chairman of the Fed and/or other FOMC members changed their beliefs about whether inflation has negative long-run consequences, their estimate of the natural rate of unemployment and/or the responsiveness of inflation to economic slack. In Figure D.1., we combine the narrative results with our estimated history of monetary policy.

- William McChesney Martin Jr. was the chairman of the Fed between April 1951 and January 1970. Romer and Romer [2004] describe Martin as having similar beliefs about monetary policy as Paul Volcker and Alan Greenspan. Martin believed that high inflation was harmful in the long-run and that monetary policy should respond to bad economic conditions.
- The Fed lowered nominal interest rates substantially in response to the 1953/54 and to the 1957/58 recessions. In Figure D.1., our measure of monetary policy activism takes a value close to zero at the beginning of our sample indicating that monetary policy has been very passive during the late 1950s.
- Following the 1957/58 recession, inflation started to raise. The Fed responded via extreme monetary tightening. Romer and Romer [2004] suggest calling this period the *Martin Disinflation*. $1 - G(z_{t-1})$ in Figure D.1. increases substantially in 1959.
- In the 1960s, the Fed changed their views and adopted “New Economics“. In particular, the Fed believed that there was a long-run trade off between inflation and unemployment and that inflation would be low even for low levels of unemployment. Consequently, the Fed loosened its policy and did not tighten monetary policy during the second half of the 1960s even when output growth was high and inflation was increasing.

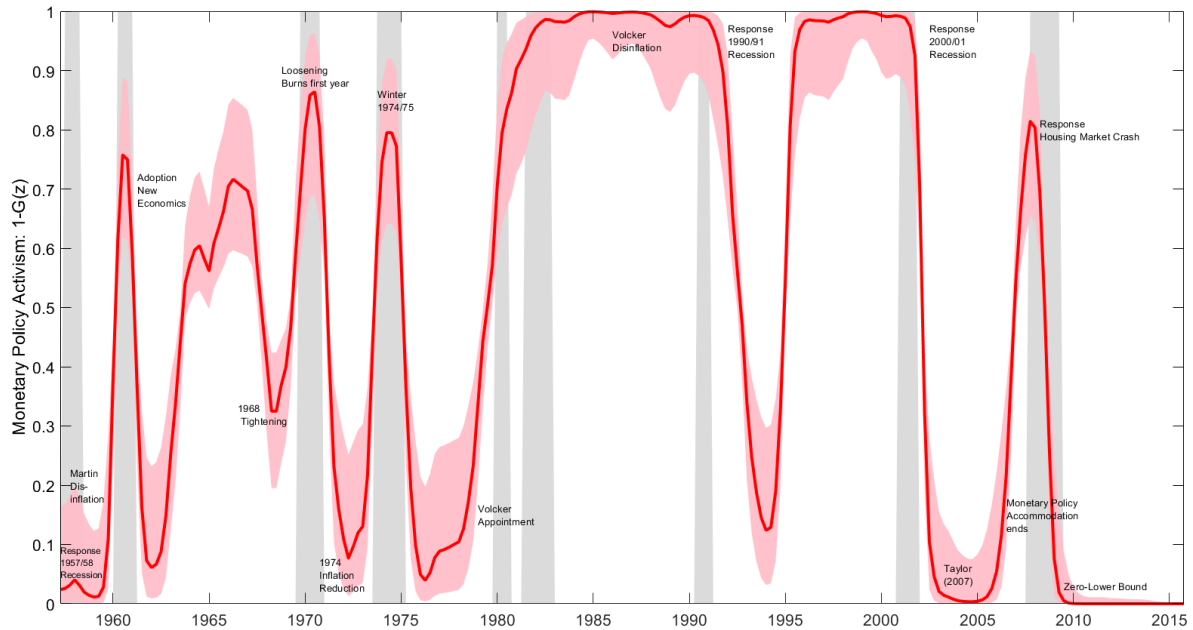
- At the end of Martin's tenure, the Fed returned to the natural rate framework but kept their optimistic view about the economy. In late 1968, the Fed tightened monetary policy substantially to combat increasing inflation. In 1968, $1 - G(z_{t-1})$ increases significantly.
- Arthur Burns followed William McChesney Martin Jr. as the chairman of the Fed in February 1970. Initially, he continued Martin's beliefs about monetary policy but became later pessimistic about the responsiveness of inflation to economic slack. At the end of his term, Burns would return to his initial views.
- During the first half Burns' first year, the Fed lowered nominal interest rate substantially. Their optimistic estimate about the natural rate of unemployment led the Fed to assume that there is significant slack in the economy. $1 - G(z_{t-1})$ in Figure D.1., displays this change as well: the monetary policy regime becomes more passive.
- When inflation rates did not fall as expected, the Fed changed its views and became overly pessimistic about the responsiveness of inflation to economic slack. According to Fed beliefs at that time, raising interest rates to cause economic slack would have been inefficient. As a consequence, the Fed imposed price controls in August 1971, which temporarily lowered inflation.
- Price controls were removed in January 1973 and inflation started to rise again. In addition, the Fed conducted an expansionary monetary policy.
- In the mid-1970s, the Fed's pessimism about the responsiveness of inflation to economic slack, restored the central bank's belief about the effectiveness of conventional monetary policy interventions. In addition, the Fed followed a higher estimate of the natural rate which in turn restored the belief that economic slack can lower inflation.
- In 1974, the Fed conducted contractionary monetary policy even though output was already decreasing. The Fed wanted to reduce inflation and was willing to accept a recession. Our estimate of monetary policy displays this change as well: $1 - G(z_{t-1})$ increases in the beginning of 1974.
- Due to increasing unemployment rates, the Fed lowered nominal interest rates in the winter of 1974/75. Monetary policy remained expansionary until the end of Burns' tenure in February 1978. Our model replicates this change: the value of $1 - G(z_{t-1})$ decreases in the winter of 1974/75 and remains high until 1979.

- William Miller would follow Arthur Burns to become the Fed's chairman between March 1978 and September 1979. In the following, the Fed became more optimistic about the natural rate of unemployment and believed monetary policy is ineffective to combat inflation: even though the Fed worried about high inflation during the late-1979s, the Fed's views on the responsiveness of inflation to slack and their high estimate of the natural rate prevented the Fed to tighten monetary policy.
- Paul Volcker became the chairman of the Fed in August 1979. The views about monetary policy changed fundamentally and remained the same during his chairmanship and under his successor Alan Greenspan: (i) high inflation has strong negative long-run consequences and little benefits, (ii) inflation responds to economic slack and (iii) their estimate of the natural rate was higher than previously assumed.
- In response to high inflation in the late 1970s, the Fed tightened monetary policy substantially which is believed to have caused the 1981/82 recession. Because of their high estimate of the natural rate, the unemployment rate necessary to reduce inflation was also high. According to our model, this change is the largest during our sample period. $1 - G(z_{t-1})$ increases from almost zero to almost one during a period of three years (1979 - 1982).
- Alan Greenspan replaced Paul Volcker as chairman of the Fed in August 1987 and served until January 2006.
- In response to the 1990/91 recession, the Fed lowered nominal interest rates. Our estimate of $1 - G(z_{t-1})$ changes from one to almost zero.
- Puzzle in the second half of the 1990s: Romer and Romer [2004] report that during the second half of the 1990s, inflation did not rise despite a long-lasting expansion. Greenspan argued that the economy became more competitive. Firms would rather cut costs than increase prices. In addition, Romer and Romer [2004] say that the Fed left real interest rates unchanged. According to our estimates, however, the Fed changed their responsiveness to inflation substantially. $1 - G(z_{t-1})$ decreases from almost one to zero and is very active until 2001.
- The Fed loosened monetary policy in response to the 2000/01 recession. We estimate that the policy regime changed from being very active to being very passive.
- Taylor [2007] argued that the Fed deviated substantially from the Taylor Principle between 2002 and 2005. He concludes that this deviation contributed enormously to

the housing bubble, which led to the Financial Crisis in 2008. Our model replicates this behavior as well. Between 2002 and 2005, our estimate of $1 - G(z)$ reaches its lowest level for the sample before the 2008 Financial crisis.

- In June 2004, the Fed raised the target rate for the federal funds rate for the first time since the 2000/01 recession. The Fed believed that monetary policy remained accommodative. The Fed kept raising the target rate for the federal funds rate to 5.25 in June 2006. However, the Fed stopped announcing that monetary policy is accommodative in August 2005. Our estimate of $1 - G(z)$ leaves its low level at zero at roughly the same time.
- In May 2005, the Fed first expressed concerns about a slowdown in economic growth, in part, because of a “gradual cooling” of the housing market.
- Ben Bernanke followed Alan Greenspan in February 2006.
- In September 2007, the Fed announced to reduce its target rate for the federal funds rate for the first time since its response to the 2000/01 recession.
- The Fed kept lowering nominal interest rates until the Fed finally cut them to zero in December 2008. Nominal interest rate remained at zero until December 2015.
- Janet Yellen replaced Ben Bernanke as the Fed’s chairman in February 2014.
- In December 2015, the Fed raised the target range for the federal funds rate to 1/4 to 1/2 percent.

Figure D.1.: Evolution of Monetary Policy between 1954 and 2016



Note: Figure shows the pointwise-posterior median estimate of $1 - G(z)$, along with the 68 percent credible bands. $1 - G(z)$ can be interpreted as a measure for monetary policy activism. Grey bars represent recessions as defined by the National Bureau of Economic Research (NBER). The Figure combines our estimated evolution of monetary policy with narrative evidence given by Romer and Romer [2004], and the author's reading of the FOMC press releases.