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Policy Coordination and the Effectiveness of Fiscal Stimulus

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Abstract

This paper shows that government spending shocks in the U.S. has become ineffective due to lack of coordination between monetary and fiscal policies. Employing the post-war U.S. data, we report strong stimulus effects of fiscal policy during the pre-Volcker era, which rapidly dissipate as the sample period is shifted toward the post-Volcker era. We explain the causes of this phenomena via a sentiment channel. Employing the Survey of Professional Forecasters data, we show that forecasters tend to systematically overestimate real GDP growth in response to positive innovations in government spending when policies coordinate well with each other. On the other hand, they are likely to underestimate real GDP responses when the monetary authority maintains a hawkish stance that conflicts with the fiscal stimulus. The fiscal stimulus, under such circumstances, generates consumer pessimism, which reduces private spending and ultimately weakens the output effects of fiscal policy. We further report statistical test results that confirm our claims.

Keywords: Fiscal Policy; Time-varying Effectiveness; Policy Coordination; Sentiment; Survey of Professional Forecasters

JEL Classification: E32; E61; E62

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1 Introduction

We often observe sluggish economic recovery from recessions that naturally brought in heated debates on the effectiveness of fiscal policy in stimulating private activity. One group of researchers reports significantly positive output effects of fiscal stimulus, which can be consistent with the New Keynesian (NK) macroeconomic model. However, such effects could be replicated only in heavily restricted models. See, among others, Rotemberg and Woodford (1992), Devereux et al. (1996), Fatás and Mihov (2001), Blanchard and Perotti (2002), Perotti (2011), and Galí et al. (2007).

Many others, on the other hand, are skeptical about the effectiveness of fiscal policy. For instance, Ramey (2011) points out that expansionary government spending shocks tend to decrease consumption due to a negative wealth effect. See also, among others, Aiyagari et al. (1992), Hall (1986), Ramey and Shapiro (1998), Edelberg et al. (1999), Burnside et al. (2004), Cavallo (2005), Mountford and Uhlig (2009), Ramey (2012), and Owyang et al. (2013).

Another interesting question is whether fiscal stimulus spurs economic activity only during economic downturns. For example, Fazzari et al. (2015), Auerbach and Gorodnichenko (2012), Mittnik and Semmler (2012), Bachmann and Sims (2012), Bernardini et al. (2020), and Albertini et al. (2021) claim that fiscal policy tends to have a stronger output effect during times of slack, whereas Jia et al. (2021), Owyang et al. (2013), and Ramey and Zubairy (2018) find no such evidence. Barnichon et al. (2022) also propose a rather depressing nonlinear (sign-dependent) effect of government spending shocks, showing that unexpected decreases in government spending tend to generate more substantial (negative) impacts than positive shocks regardless of the state of the cycle.

Hall (2009) and Christiano et al. (2011) suggest that the government spending multiplier can be greater when the nominal interest rate is bounded at zero. Jo and Zubairy (2021) show that the government spending multiplier is larger in a low inflation recession than in a high inflation recession. Ghassibe and Zanetti (2020) linked the state dependence of fiscal multipliers to the source of the business cycle. Employing large panel of international data, Corsetti et al. (2012), Ilzetzki et al. (2013), and Born et al. (2019) report strong evidence of important roles of country characteristics such as the exchange rate regime and public indebtedness in determining the effectiveness of fiscal stimulus.

In their recent work, Leeper et al. (2017) proposed an interesting theoretical framework that generates substantially weaker responses of private spending to expansionary fiscal policy shocks in an active monetary/passive fiscal policy regime (Regime M) than in a passive monetary/active fiscal policy regime (Regime F).[1] This paper employs a similar New Keynesian DSGE model but with alternative identification strategies of the regimes. We note that the federal government had a budget deficit in 79 times out of 93 years from 1929 to 2021, which is about 85% odds.[2] In the absence of compelling empirical evidence of contractionary fiscal policy, we focus on the stance of monetary policy that tends to change over time given the expansionary stance of fiscal policy: (1) a dovish monetary policy coordinated well with an expansionary fiscal policy regime (Regime D); (2) a hawkish monetary policy conflicted with an expansionary fiscal policy regime (Regime H).[3]

In what follows, our simulation results demonstrate that private spending positively responds to the government spending shock in regime D, whereas it responds negatively in regime H, resulting in substantially weaker stimulating effects on the total output. Also, our benchmark model shows that the real interest rate plays a key role in generating qualitatively different output effects across the two regimes.

We are primarily interested in investigating the empirical validity of these theoretical predictions. Employing the post-war U.S. macroeconomic data, we report strong evidence of the time-varying effectiveness of fiscal stimulus with a possibility of structural breaks in the propagation mechanism of the government spending shock across time. Specifically, we observed strong effects of government spending on boosting private economic activity in earlier sample periods when the Fed stayed accommodative, while government spending shocks tend to discourage economic activity in the private sector when the Fed shifted to a hawkish stance which may conflict with expansionary fiscal policy. These findings are further justified via a Bayesian approach (time-varying coefficient VAR with stochastic volatility) and counterfactual simulation exercises.

Although these findings are overall consistent with the predictions of our proposed NK model, we notice a negligibly weak role of the real interest rate in propagating fiscal stimulus to economic activity. To resolve this issue, we propose an alternative explanation for the observed time-varying output effects of fiscal policy shocks using a sentiment channel.

It should be noted, however, that we are not the first to introduce the role of sentiment as one of potential drivers of macroeconomic fluctuations or as the instrument to evaluate

[1] An active monetary policy regime refers a case that the monetary authority responds to inflation aggressively. A passive fiscal policy regime means that dynamics of government spending has a strong feedback from rising government debt.

[2] See the federal surplus/deficit as percent of Gross Domestic Product (FYFSGDA188S) from the FRED.

[3] We assume that the central bank maintains a dovish policy stance that coordinates well with expansionary fiscal policy in regime D. In regime H, however, monetary policy makers respond aggressively to inflationary pressure, conflicting with fiscal stimulus. That is, we assume that fiscal and monetary policies are well coordinated only in regime D.

the stimulus effects of government policies. There is a large and fast-growing literature. Hall (1993), Blanchard (1993), Cochrane (1994), Christopher D. Carroll and Wilcox (1994), Ludvigson (2004), Beaudry and Portier (2006, 2007), Akerlof and Shiller (2010), and Barsky and Sims (2012) studied the relationship between consumer confidence and consumption expenditures, while Konstantinou and Tagkalakis (2011), Bachmann and Sims (2012), and Jia et al. (2021) focused on how consumer confidence affects the effectiveness of fiscal policy.

Using the recent experience during the Covid-19 pandemic crisis, Georgarakos and Kenny (2022) find that consumers perceptions that are related to the effectiveness of fiscal polices have a causal effect on consumer spending. An array of recent researches by George-Marios Angeletos are closely related with such views, both theoretically and empirically, in the sense that aggregate demand shocks may cause business cycle fluctuations under the bounded rationality. See among others, Angeletos and Lian (2018), Angeletos et al. (2020), Angeletos and Lian (2022a) and Angeletos and Lian (2022b). For example, Angeletos and Lian (2022a) introduce a feedback loop between real economic activity and consumer expectations that amplifies the business cycle fluctuations that are caused by demand shocks, suggesting the concept of the confidence multiplier. See also Gabaix (2020) and Farhi and Werning (2019) for the policy propagation channel analysis under the New Keynesian models with the bounded rationality.

We address the responses of sentiment to changes in government spending by investigating how market participants revise their economic prospects when they receive new information on the stance of fiscal policy through the lens of the Survey of Professional Forecasters (SPF) data. We show that forecasters tend to over-predict GDP growth when monetary policy coordinates well with fiscal policy, while systemic under-predictions are likely to occur when the Fed adopts a hawkish stance. We view persistent over-predictions as a sign of optimism, while under-predictions reflect pessimistic economic prospects in the market.

We further investigate this conjecture by regressing five-quarter (longer-run) ahead forecasts of real GDP growth on one-quarter (short-run) ahead forecasts of real government spending growth employing a fixed-size rolling window scheme. Results reveal strong positive correlations (optimism) for the pre-Volcker era, while we observe negative correlations (pessimism) when the stance of monetary policy became hawkish. These findings imply that time-varying responses in sentiment may explain the time-varying effectiveness of fiscal policy on private spending. In regime D, fiscal stimulus generates consumer optimism, which boosts economic activity in the private sector. In regime H, however, it generates consumer pessimism, resulting in subsequent decreases in private spending, which ultimately weakens the effectiveness of fiscal policy. We also provide statistical evidence in favor of such views employing structural break tests by Hansen (1997, 2001) and Bai and Perron (2003).

Leeper et al. (2017) also show that fiscal policy can be less effective when the monetary authority stays hawkish. However, their contributions are mostly theoretical because their major findings are based on counterfactual analyses using the full sample period data. On the contrary, we provide historical evidence of the time-varying effects of fiscal stimulus for the post-war U.S. data. Furthermore, we suggest a sentiment channel as an alternative propagation mechanism to the real interest rate channel to explain the output effects of fiscal policy under different policy regimes.

Perotti (2005) suggests similar evidence that fiscal policy became less effective in more recent sample periods using macroeconomic data from 5 OECD countries including the U.S. However, he fails to provide convincing explanations what caused such changes. Bilbiie et al. (2008) also report time-varying effects of fiscal stimulus, but they focus more on the role of different feedback rules of government spending similar to the work of Leeper et al. (2017). They suggest that financial market deregulation made it possible for households to smooth consumption, which makes fiscal policy less effective.[4] It seems, however, that these arguments are at odds with the data. In fact, saving rates have substantially declined since the 1980's when deregulation began in the U.S.

The remainder of this paper is organized as follows. Section 2 reports simulation results that highlight qualitatively different output effects of fiscal stimulus across regimes based on our New Keynesian model in the Appendix. Section 3 presents our empirical models along with data descriptions. We demonstrate time-varying responses of our key economic variables to government spending shocks via both the benchmark frequentist models as well as a Bayesian approach. We report evidence against an important role of the real interest rate in propagating fiscal stimulus over time. We then discuss a possibility of the existence of a sentiment channel as an alternative explanation of which the importance of its role is confirmed by counterfactual simulation exercise. Employing the SPF data, section 4 provides a novel statistical approach that extracts useful information on how market participants revise their economic prospects when they receive new information on government spending. We show market agents become more optimistic in regime D in response to the government spending shock, while they become pessimistic in regime H, which helps explain weaker output effects of fiscal policy in regime H. Section 5 concludes.

[4]They argue that more savings instruments became available due to financial market deregulation, which helped households to act in line with the permanent income hypothesis.

2 Model Simulations

This section reports simulation results, obtained from a New Keynesian model that features external habit formation in consumption, variable capacity utilization, investment adjustment costs, and monopolistic competition in the production. Sticky prices and sticky wages are modeled using the framework of Calvo (1983) and Yun (1996). Government spending directly enters the household's utility function as a complement to private consumption. As shown by Linnemann and Schabert (2004) and Leeper et al. (2017), this specification helps reconcile theory and empirical evidence regarding the consumption response to government spending shocks. Monetary authority is assumed to follow a Taylor rule, while fiscal rules are specified with a feedback to government debt similar to those of Leeper et al. (2017).

It should be noted that our paper is largely empirical, and the model presented in this section is mainly to clearly introduce our definitions of policy regimes. For this purpose, we present a standard New Keynesian DSGE model, leaving a full description and derivations of the model in the Appendix including calibrated parameter values in Table B1.

As briefly explained in the introduction, we define the following two regimes that highlight the implications of policy coordination of monetary and fiscal policies. In regime D, policy makers stay accommodative in the stance of both monetary and fiscal policies. The dovish central bank puts greater emphasis on output stabilization, thus responds only weakly to inflation to keep the balance between output and inflation stability. In regime H, however, the hawkish central bank prioritizes keeping inflationary pressure in check, which results in more aggressive responses to the inflation gap, conflicting with the fiscal stimulus of the government. See the Appendix for more detailed description of our model.

Figure 1 displays simulated impulse-response functions (IRFs) of key macroeconomic variables to a 1% government spending shock under these two regimes: regime D (solid) and regime H (dashed). We observe persistently positive output effects of fiscal policy only in regime D, where the monetary authority maintains a dovish monetary policy stance in collaboration with the fiscal stimulus. Output and inflation both rise in response to the government spending shock, but the central bank raises the interest rate at a slower rate than inflation, resulting in a decrease in the real interest rate for about two years. Responses of the private GDP also stay positive in the first two years until persistently positive consumption responses are dominated by the negative response of investment. The total GDP exhibits persistent, solid positive responses even when the private GDP responds negatively after the first two years, which implies that responses of public (government) spending dominate those of the private GDP.

On the other hand, we obtained substantially weaker output effects of fiscal policy in regime H, which sharply contrast with those in regime D. In response to the government spending shock, the nominal interest rate rises faster than inflation as the central bank raises the interest rate aggressively to curb inflation, maintaining a hawkish policy stance. Consequently, the real interest rate rises, crowding out investment, which results in immediate decreases in the private GDP. Private consumption responds positively, reflecting the complementarity between government spending and consumption. However, its positive responses are dominated by decreases in investment, resulting in negative responses of the private GDP. The total GDP rises in the short-run, driven by increases in government spending, but eventually falls below zero due to substantial negative responses of the private GDP.

Overall, our simulation results demonstrate that the effectiveness of fiscal policy greatly hinges upon the coordination of monetary and fiscal policies. In the next section, we report strong empirical evidence of time-varying output effects of fiscal policy in the private sector using the post-war U.S. macroeconomic data. We found a very limited role of the real interest rate in the propagation mechanism of fiscal policy, which is at odds with the simulation results from our benchmark model presented in this section. In what follows, we suggest a sentiment channel as an alternative to the real interest rate channel.

Figure 1 around here

3 The Empirics

This section presents our baseline empirical model for the U.S. post-war macroeconomic data. We report solid empirical evidence that supports time-varying output effects of fiscal policy.

3.1 The Baseline Empirical Model

We employ the following vector autoregressive (VAR) process of order p.

$$\mathbf{x}_t = \gamma' \mathbf{d}_t + \sum_{j=1}^{p} \mathbf{A}_j \mathbf{x}_{t-j} + \mathbf{C}\varepsilon_t, \tag{1}$$

where

$$\mathbf{x}_t = [g_t \quad y_t \quad \mathbf{z}_t]',$$

$\mathbf{d}_t$ is a vector of deterministic terms that includes an intercept and up to quadratic time trend. $\mathbf{C}$ denotes a lower-triangular matrix and ε_t is a vector of mutually orthonormal structural shocks, that is, $E\varepsilon_t\varepsilon_t' = \mathbf{I}$. We are particularly interested in the j-period ahead orthogonalized impulse-response function (IRF) defined as follows.

$$IRF_{k,j} = E\left(\mathbf{x}_{t+j}|\varepsilon_{k,t} = 1, \Omega_{t-1}\right) - E\left(\mathbf{x}_{t+j}|\Omega_{t-1}\right), \tag{2}$$

where $\varepsilon_{k,t}$ is the structural shock to the k^{th} variable in (1) that occurs at time t. Ω_{t-1} is the adaptive information set at time $t-1$, that is, $\Omega_j \supseteq \Omega_{j-1}, \forall j$.

g_t denotes (log) federal government spending, which is used to identify the fiscal policy shock. Following Blanchard and Perotti (2002), we employ *discretionary* components of government spending, which led g_t to be ordered first in $\mathbf{x}_t$, meaning that g_t is not contemporaneously affected by innovations in other variables within one quarter. This assumption is frequently employed in the current literature, because implementations of discretionary fiscal policy actions require Congressional approvals, which normally take more than one quarter.[5]

y_t is the (log) real per capita gross domestic product (gdp_t).[6] $\mathbf{x}_t$ includes a vector of control variables from the money market $\mathbf{z}_t = [int_t \quad mon_t]'$, where int_t is the effective federal funds rate, which can be used to identify the monetary policy shock, and mon_t is the (log) monetary base.[7] These variables are ordered in the last block, because the Federal Open Market Committee (FOMC) can revise the stance of monetary policy by holding regular and emergency meetings whenever policy-makers deem it necessary. Note that int_t is ordered before mon_t, because the Fed targets the interest rate, while the monetary base changes endogenously.

It is well known that econometric inferences from recursively identified VAR models may not be robust to alternative VAR ordering. It turns out that our empirical findings are

[5]Kim and Jia (2017) employed the government total expenditures that includes transfer payments in addition to the discretionary government consumption and investment spending. Since transfer payments have automatic stabilizers, they put g_t next to y_t.

[6]We also consider the private real GDP per capita $pgdp_t$ for y_t to measure the stimulus effects of fiscal policy on private activity. In addition, we directly employ private spending variables for y_t such as private consumption ($conm_t$) and private investment ($invt_t$).

[7]We observed no evidence of structural breaks in the output effects of monetary policy. Results are available upon requests.

not subject to such criticism as long as we are interested in the IRFs to the government spending (g_t) shock, $IRF_{1,j}$. Given the location of g_t, $IRF_{1,j}$ is *numerically identical* even if all variables next to g_t are randomly re-shuffled. See Christiano et al. (1999) for details.[8]

3.2 Data Descriptions

We obtained most data from the Federal Reserve Economic Data (FRED) website. Observations are quarterly frequency and span from 1960Q1 to 2017Q3.

g_t is federal consumption expenditures and gross investment (FGCE), which constitutes *discretionary* components of federal government expenditures. The private GDP ($pgdp_t$) is the total GDP (gdp_t; NGDP) minus total (federal and state & local) government consumption expenditures and gross investment (GCE). Consumption ($conm_t$) is total personal consumption expenditures on nondurables (PCND) and services (PCESV). Investment ($invt_t$) denotes private nonresidential fixed investment (PNFI). All spending variables are expressed in real per capita terms. That is, they are divided by the GDP deflator (GDPDEF) and by the civilian noninstitutional population (CNP16OV), then log-transformed.

The nominal interest rate (int_t) is the effective federal funds rate (FEDFUNDS) divided by 100, which can be used to identify the monetary policy shock.[9] mon_t is the monetary base (BOGMBASE), expressed in natural logarithm. We also employ the *ex post* real interest rate in our VAR models, which equals int_t minus the consumer price index (CPIAUCSL) based inflation.

Sentiment ($sent_t$) is not directly observable. Thus, we later augment our benchmark VAR model (1) with a sentiment index obtained from the University of Michigan's Survey of Consumers. Our VAR model mainly utilizes the Index of Consumer Expectations (ICE), which provides information on the level of consumer confidence about economic conditions in the near future. In addition to this forward-looking sentiment index, we also employ the Current Conditions Index and the Index of Consumer Sentiment (combined index), obtained from the same source. All three indices are highly correlated with each other, thus yield qualitatively similar empirical results.[10]

One noble feature of this paper is to infer the changes in unobservable sentiment using the Survey of Professional Forecasters (SPF) data. Unlike the ICE that reflects consumers'

[8]Similarly, all response functions to monetary policy shocks are robust to alternative ordering given the location of monetary variables, int_t and mon_t.

[9]We observed no significant evidence of structural breaks in the output effects of monetary policy. Results are available upon requests.

[10]Results are available from authors upon request.

expectations about the overall economic conditions, SPF data shows market participants' forecasts of individual macroeconomic variable. This allows us to extract revisions of people's economic prospects given new information on fiscal policy. The median SPF forecasts data for relevant variables were obtained from the Philadelphia Fed website for the period between 1968$Q4$ and 2017$Q3$.[11] There were 9 changes in the base year in the National Income and Product Account (NIPA) during this sample period. Some authors (Ramey (2011); Forni and Gambetti (2016)) use growth rates of the SPF forecasts without adjusting for these changes in the base year, which generates 9 outlier observations in the data. To prevent this, we re-scaled all relevant forecast data so that they are expressed in 2009 dollar terms.

It should be noted that forecasters were asked to predict *nominal* defense spending until 1981$Q2$.[12] Since then, they were asked to predict *real* federal consumption expenditures and gross investment. Following Ramey (2011), we used the GDP deflator median forecasts to convert nominal defense spending forecasts to real defense spending forecasts. We combine the real defense spending forecasts with the real federal spending growth forecasts in order to acquire the data for reasonably long sample period. This seems to be a fairly good approximation for the growth rate forecasts, because they tend to exhibit high degree comovements.[13] Ramey (2011) also employed a similar approach. In what follows, we study how market participants' revisions of the forecasts of government spending growth, $g_{t+j}^{SPF} - g_{t-1}^{SPF}$ are associated with their forecasts of the output growth in the future, $y_{t+j+k}^{SPF} - y_{t-1}^{SPF}, k > 0$.[14]

3.3 Empirical Findings

3.3.1 The Weakening Effectiveness of Fiscal Stimulus

This section reports an array of the impulse-response function ($IRF_{1,j}$) estimates to a positive 1% structural shock to government spending (g_t) as described in (1) and (2). We also report 90% confidence intervals obtained from 500 nonparametric bootstrap simulations using the empirical distribution.[15] Our findings below demonstrate that the output effects of fiscal

[11] The mean SPF forecasts yielded qualitatively similar results.

[12] We thank Tom Stark at the Philadelphia Fed who kindly provided us nominal defense spending data from 1968$Q4$ to 1981$Q2$, which are not available on the SPF website.

[13] Results are available upon requests from authors.

[14] We assume these forecasts are formulated utilizing the information set at time $t-1$, since the current period data such as y_t and g_t are not known at time t. Note that forecasters are asked to predict, or nowcast, y_t and g_t. Note also that forecasters are asked to predict the values at time $t-1$ (previous period) because these values are subject to revisions, although their predictions normally stay the same from the previous period.

[15] The 5^{th} and 95^{th} percentiles of the 500 response function estimates constitute the 90% confidence interval.

stimulus have become substantially weaker over time.

Figures 2 and 3 present the responses of the GDP (gdp_t) and the private GDP ($pgdp_t$), respectively, to the expansionary government spending shock from a quad-variate VAR with $\mathbf{x}_t = [g_t \quad gdp_t(pgdp_t) \quad int_t \quad mon_t]'$. Specifically, figures in the panel (a) are based on the first 30-year sample period (SP1), 1960Q1 to 1989Q4, while the last 30-year sample period (SP2), 1987Q4 to 2017Q3, was used to generate the IRFs in the panel (b).

It should be noted that the output responses from these sub-sample periods are *qualitatively* different. The IRF point estimates of the total GDP and the private GDP to the government spending shock are well above zero in SP1 ($1960Q1 - 1989Q4$), whereas their responses have become substantially muted when we employ data in SP2 ($1987Q4 - 2017Q3$). Putting it differently, both output responses remain positive for a prolonged period of time in SP1, but their responses become overall negative in SP2. We also note that the private GDP never respond positively to the shock in SP2, implying that initial positive responses of the total GDP simply reflect increases in government spending.

These IRFs imply the possibility of the time-varying effectiveness of fiscal policy in stimulating private activity ($pgdp_t$). In other words, the government spending shock seems to have promoted private spending in SP1 but not in SP2.

Motivated by these findings, we further investigate such possibility via repeated VAR model estimations with a fixed-size rolling window scheme described as follows. We use the rolling-window scheme instead of a recursive scheme because we are interested in detecting structural changes in the data generating process of $\mathbf{x}_t$.

We begin with an estimation of the VAR model using the first $T_0(< T)$ observations, $\{\mathbf{x}_t\}_{t=1}^{T_0}$. After obtaining the first round set of IRF estimates, we move the sample period window forward by one. That is, new observations at time $T_0 + 1$ ($\mathbf{x}_{T_0+1}$) are added to the sample, but we drop the oldest ones at time $t = 1$ ($\mathbf{x}_1$) to maintain the same size of the sample window. Using $\{\mathbf{x}_t\}_{t=2}^{T_0+1}$, we estimate the second round set of IRFs. We repeat until we obtain the last round IRFs using $\{\mathbf{x}_t\}_{t=T-T_0+1}^{T}$, totalling $T - T_0 + 1$ sets of the IRF estimates.

We report our estimates with a 30-year ($T_0 = 120$ quarters) fixed-size rolling window in the lower panel of Figures 2 and 3 for the GDP variables.[16] The range of the x-axis (Date) is from 1989Q4 to 2017Q3, where the fine grid points indicate the ending period of each rolling window. The y-axis (Year) is the time horizon (j) of the response function indexed from 0 to 5 years. The z-axis is the response ($IRF_{1,j}$) of each variable to a 1% government

[16]We also implemented the same analysis with a 20-year window as well as a 40-year window scheme. Results are overall similar and are available upon requests.

spending shock.

The surface graphs in the panel (c) of Figures 2 and 3 reveal dramatic decreases in the responses of gdp_t and $pgdp_t$ over time, respectively. Strong positive responses of the GDP variables are rapidly dragged down as observations are shifted toward later sample periods.

It should be noted that the responses of the private GDP become substantially negative, pushing the total GDP responses toward a negative region, which implies that the weakening stimulus effects of fiscal policy are mainly driven by time-varying responses of private spending.

To highlight these transitions over time, in panel (d) of Figures 2 and 3, we report the responses of the output variables in the short-run to the long-run by dissecting the surface graphs at $y = 0,\ 2,\ 5$ (years) of the y-axis from the right to the left. Contemporaneous responses (impact; $y = 0$) of gdp_t and $pgdp_t$ do not exhibit substantial variations over time, while the responses in 2 years and in 5 years clearly show a downward trend, implying the substantially diminished effects of fiscal stimulus over time. It is also interesting to see that positive responses of gdp_t on impact ($y = 0$) are due to increases in g_t itself because $pgdp_t$ barely responds when the shock occurs.

Figures 2 and 3 around here

Observing these remarkably dramatic changes in private GDP responses over time, we further look into the source of these transitions by investigating the IRFs of the two private spending variables, consumption ($conm_t$) and investment ($invt_t$). In Figures 4 and 5, we report the IRF estimates with $\mathbf{x}_t = [g_t \ \ conm_t(invt_t) \ \ int_t \ \ mon_t]'$.

We note a close resemblance between consumption ($conm_t$) responses and those of the private GDP ($pgdp_t$) as can be seen in Figure 4. Consumption increases greatly and significantly over time when g_t shock occurs in SP1. In SP2, however, consumption responses continue to fall to a negative region, although we still observe a weak but positive responses in the very short-run. The surface graph in the panel (c) confirms rapid deteriorations of the consumption responses over time. The panel (d) graph also shows a clear downward trend of the responses of consumption to the spending shock in the medium-run and in the long-run.

On the other hand, investment ($invt_t$) responds overall negatively to the government spending shock as can be seen in Figure 5, although negative responses of $invt_t$ tend to go deeper as the sample period moves forward.[17] These findings are confirmed by the downward trend in the IRFs in the medium-run as well as in the long-run, whereas initial responses are overall negligible as can be seen in Figure 5(d).

These IRF analyses provide strong evidence that fiscal stimulus has become less effective in stimulating private spending. The positive responses of the private GDP in earlier sample periods are mostly driven by rising consumption given overall negative responses of investment to the shock. On the contrary, fiscal policy has become dramatically ineffective over time. The private GDP ($pgdp_t$) responds mostly negatively to the government spending shock when more recent sample periods are employed, generating completely ineffective stimulus effects of fiscal policy.

Figures 4 and 5 around here

3.3.2 Assessing the Role of the Interest Rate under Different Regimes

This subsection empirically assesses the role of the interest rate channel of expansionary fiscal policy shocks under different policy regimes described earlier in our theoretical models. Section 2 demonstrates that government spending shocks generate persistent stimulus effects on private spending only in regime D when the monetary policy stance stays accommodative. The nominal interest rate rises more slowly than the inflation rate, resulting in decreases in the real interest rate, which stimulate private investment as well as consumption.

On the other hand, the nominal interest rate rises faster than the inflation rate in regime H as the central bank maintains its hawkish policy stance to suppress inflationary pressure. The real interest rate rises, which decreases private investment substantially, dominating positive responses of consumption in the short-run. Therefore, fiscal policy fails to stimulate private activity in regime H.

The U.S. post-war data seems to be overall consistent with the theoretical predictions on the output effects of fiscal policy. However, the data shows a very limited role of the interest rate in the transmission mechanism of fiscal policy. For this purpose, we consider the VAR model (1) with $\mathbf{z}_t = [rffr_t \;\; mon_t]'$, where $rffr_t$ is the *ex-post* real federal funds

[17] We obtain negligible responses of $invt_t$ from the first 30-year sample period, $1960Q1$˜$1989Q4$. These seem to be outliers because we obtained qualitatively similar negative responses by shifting the window by just a few years such as $1962Q1$˜$1991Q4$.

rate (FEDFUNDS) accompanied by the log-transformed monetary base. We used the CPI-based inflation rate to obtain the *ex-post* inflation.

Clarida et al. (2000) demonstrated that the Fed had remained dovish (accommodative) during the pre-Volcker era ($1960Q1-1979Q2$), while it had switched to a hawkish monetary policy stance after Paul Volcker's tenure began in the third quarter of 1979. With the Taylor rule parameter estimates from their work, the New Keynesian model predicts the real interest rate to rise in response to the government spending shock in regime H (post-Volcker era), while it is expected to decline in regime D (pre-Volcker era).

We report empirical evidence that is at odds with these predictions. As can be seen in Figure 6, $rffr_t$ positively responds to the fiscal shock in SP1 ($1960Q1 - 1989Q4$), while it responds negatively in SP2 ($1987Q4 - 2017Q3$).[18] Also, the IRFs of $rfft_t$ from the rolling window scheme in the panel (c) and (d) clearly demonstrate a downward trend in all horizons. Recall private investment tends to decline in response to the government spending shock in both regimes. That is, $rffr_t$ and $invt_t$ both decline in response to the fiscal shock in SP2. This implies that private investment must have *shifted* to the left by *exogenous* factors rather than endogenously responding to changes in the real interest rate. We introduce a sentiment channel to explain this possibility of exogenous factors in the next section.

Figure 6 around here

3.3.3 Fiscal Policy Effects on Sentiment

The role of sentiment as one of potential drivers of macroeconomic fluctuations has long been discussed in the current literature. Hall (1993) and Blanchard (1993), among others, emphasize the causal effects of animal spirit on economic activity, whereas Cochrane (1994) claims that consumer confidence reflects news about changes in economic productivity in the future, which creates a close link between innovations in consumer confidence and subsequent variations in economic activity.

Using a nonlinear state-dependent VAR model, Bachmann and Sims (2012) suggest that the government spending shock can trigger consumer optimism during times of slack, which results in a high fiscal multiplier during recessions. On the other hand, Jia et al. (2021)

[18]We used 30-year windows to obtain sharp estimates. Similar results are obtained using the pre-Volcker era with a 20-year window.

demonstrate that the shock is likely to generate consumer *pessimism* in all phases of business cycle when properly detrended data are used.

Recognizing a potentially important role of sentiment, we shift our attention from state-dependent nonlinearity to a time-dependent stochastic process because sentiment responses seem to change over time. For this purpose, we estimate and report the time-varying dynamic adjustments of sentiment in response to the government spending shock, utilizing the VAR model (1) with $\mathbf{z}_t = [int_t \;\; mon_t \;\; sent_t]'$. Recall that the location of $sent_t$ in the VAR does not matter for the fiscal policy effects as long as $sent_t$ is placed next to g_t. See Christiano et al. (1999) for detailed explanations on this property.

Figure 7 clearly shows qualitatively different responses of sentiment over time. Sentiment ($sent_t$) responds positively to the government spending shock in SP1 ($1960Q1 - 1989Q4$), while the shock generates consumer pessimism in SP2 ($1987Q4 - 2017Q3$). The figures in the lower panel exhibit a downward trend especially in the two-year and in the five-year sentiment responses, while a long swing is observed in the contemporaneous responses on impact.

We consider these changes in the response function of $sent_t$ as a clue to understand why the output effects of fiscal stimulus have become weaker over time. Significant stimulating effects of fiscal policy during earlier sample periods are consistent with consumer optimism that results from the government spending shock. On the other hand, it tends to generate consumer pessimism with later observations, decreasing not only investment but also consumption.[19]

One possible criticism against this view is the following. Sentiment may simply reflect changes in consumption rather than leading it. This doesn't seem to be the case especially in SP2. As can be seen in Figure 4(b), consumption initially responds positively for a while when the government spending shock occurs, whereas sentiment starts deteriorating immediately in Figure 7(b). That is, pessimism goes deeper since the impact of the shock. If $sent_t$ simply reflects the changes in $conm_t$, sentiment must have risen at least in the short-run because consumption rises in the short-run. Therefore, sentiment seems to be leading the innovations in consumption rather than passively reflecting innovations in consumption.

In what follows, we implement counterfactual simulation exercises to provide some insights on the importance of this sentiment channel in propagating government spending shocks to private activity.

[19] It might be the case that large sudden increases in government spending are perceived as a confirmation of an incoming recession in near future, generating consumer pessimism, which then results in a decrease in private spending.

Figure 7 around here

3.4 Robustness Check Analysis: Bayesian Evidence

3.4.1 The Time-Varying Parameter VAR Model with Stochastic Volatility

This subsection implements robustness check analysis via the time-varying parameter VAR with stochastic volatility (TVPVAR-SV) model framework following the work of Primiceri (2005) and Negro and Primiceri (2015). For this purpose, we rewrite our baseline empirical model (1) as follows.

$$\mathbf{x}_t = \gamma_t^{'} \mathbf{d}_t + \sum_{j=1}^{p} \mathbf{B}_{j,t} \mathbf{x}_{t-j} + \mathbf{u}_t, \tag{3}$$

where $\mathbf{x}_t$ is a $k \times 1$ vector of endogenous variables $[g_t \quad y_t \quad \mathbf{z}_t]'$ and $\gamma_t = [\alpha_t \quad \delta]'$ denotes a vector of coefficients on $\mathbf{d}_t = [1 \quad t]'$.[20] $\mathbf{B}_{j,t}$ denote matrices of time-varying coefficients, and $\mathbf{u}_t$ is a vector of heteroskedastic unobservable shocks with variance-covariance matrix $\mathbf{\Omega}_t$. We follow the conventional setup as follows.

$$\mathbf{u}_t = \mathbf{C}_t \Sigma_t \varepsilon_t, \tag{4}$$

where $\mathbf{C}_t$ is a time-varying lower-triangular matrix and Σ_t is the diagonal matrix with $\sigma_{i,t}, i = 1, ..., k$, the standard deviations of orthonormal structural shocks ε_t, $E\varepsilon_t\varepsilon_t^{'} = \mathbf{I}$ and $\mathbf{\Omega}_t = \mathbf{C}_t \Sigma_t \Sigma_t^{'} \mathbf{C}_t^{'}$.

θ_t denotes the vector of time-varying coefficients, $\left[\alpha_t \; vec\left(\mathbf{B}_{1,t}\right)' ... vec\left(\mathbf{B}_{p,t}\right)'\right]'$, stacked by rows, while $\mathbf{c}_t$ is the vector of non-zero elements (stacked by rows) of $\mathbf{C}_t$. The model's dynamics are described as follows.

$$\begin{aligned} \theta_t &= \theta_{t-1} + \nu_t, \\ \mathbf{c}_t &= \mathbf{c}_{t-1} + \xi_t, \\ \ln \sigma_{i,t} &= \ln \sigma_{i,t-1} + \eta_{i,t} \end{aligned} \tag{5}$$

[20] As we introduce the time-varying intercept (α_t), we assume that the coefficient on linear time trend δ is time-invariant to reduce the parameter dimension.

All the innovations are assumed to be jointly normally distributed. That is,

$$\begin{bmatrix} \varepsilon_t & \nu_t & \xi_t & \eta_t \end{bmatrix}' \overset{iid}{\sim} N(\mathbf{0}, \mathbf{V}), \tag{6}$$

$$\mathbf{V} = \begin{bmatrix} \mathbf{I} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{Q} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \mathbf{S} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{W} \end{bmatrix},$$

where diagonal elements including the identify matrix are positive definite matrices.

3.4.2 Bayesian Priors and Sample Splitting

We simulate the Bayesian posterior distribution to estimate the parameters. The priors for the coefficients are jointly normal while the covariance matrices (Q, W, S) are Inverse-Wishart. c_t and $\ln \sigma_{i,t}$ are also normally distributed.

We split our sample into two periods. The first period is from 1960 to 1988 and the second period is from 1989 to 2017. We divide the two periods up to have roughly equal number of observations. Then we follow Primiceri (2005) to choose the hyperparameters and estimate a separate model in each period. For example, we use the OLS estimates, from the first forty observations, as the hyperparameters of the coefficients' prior distribution in each period. For the first (second) period, the OLS estimates from 1960 to 1969 (from 1989 to 1998) forms the prior and the posterior estimates start from 1970 to 1988 (from 1999 to 2017).

We use sample splitting because our full-sample estimates seems sensitive to the prior values - showing signs of over-regularization. The full-sample's prior uses the OLS estimates from 1960 to 1969. In the presence of structural changes, 1960-1969 OLS estimates can be misleading priors for posterior estimates in 1999 to 2017. Under over-regularization, misleading priors can lead to substantial bias in posterior estimates. In Appendix F, we find the 1989-1998 OLS estimates have better 1999-2017 predictive performance than 1960-1969 OLS estimates. Sample splitting allows us to use 1989-1998 OLS estimates instead as priors for posterior estimates in 1999 to 2017.

In Appendix F, we also attempted to diffuse the prior distribution, but the prior dominance seems persistent even then. We suspect the prior dominance is caused by our number of parameters being larger than usual for the TVPVAR-SV model. For example, Primiceri (2005) is a VAR(3) system with no linear trend. While we have a VAR(4) system with a linear trend. For the sentiment VAR model, we removed mon_t to have the system as VAR(4).

3.4.3 Posterior Distribution and IRF Estimations

We employ Gibbs sampling steps to stimulate the quarterly posterior distributions following Negro and Primiceri (2015). After discarding the first simulated $3,000$ draws as burn-ins, we simulate the posterior distribution based on the $5,000$ draws that follow after.

We use the posterior draws of the parameters and forward iteration to stimulate the posterior distribution of the impulse responses. We simulate an independent shock by orthogonalizing a posterior draw of Ω_t. And we use the posterior distribution's median as our point estimate to be robust against outliers. And the credible set (similar to the confidence interval) is generated from the posterior distribution's 5% and 95% percentiles.

Figure 8 reports our time-varying Bayesian IRF estimates for three key variables, private GDP ($pgdp_t$), consumption ($conm_t$), and Sentiment ($sent_t$), which corresponds to previous time-varying IRFs with a rolling window scheme. We used initial 10-year observations to form the prior for each sub-sample period. We obtained qualitatively similar results, confirming our major empirical findings.[21]

As can be seen in the first row figures for $pgdp_t$ responses, the Bayesian IRFs from the first subsamples remain positive on impact through 5-year horizons. On the other hand, the IRFs from the second subsamples drop to negative areas after the shock occurs. Note that the IRF estimates from our TVPVAR-SV models demonstrate stable responses over time in each sub-sample, although their responses are qualitatively and quantitatively different across the subsamples. We view these results are consistent with the previous findings from the rolling-window scheme that exhibit gradual changes as the estimation window slowly moving towards different regimes. IRF estimates for $conm_t$ and $sent_t$ also confirm such assessment.

Figure 8 around here

3.4.4 Diagnostics

We find 3000 burn-in draws to be sufficient for simulating draws as approximately mutually independent. Appendix F provides graphs showing the average autocorrelation of coefficient

[21]The results for other variables are also qualitatively similar as those from the benchmark models. Full estimation results are available upon requests.

draws for the years 1985 and 2010. We see a sharp decay in the autocorrelation coefficients as evidence of approximate independence.

3.5 Counterfactual Simulation Exercises

This section implements counterfactual simulation exercises employing the 5-variate VAR model that is used in the previous section, $\mathbf{x}_t = [g_t \quad y_t \quad int_t \quad mon_t \quad sent_t]'$. The primary purpose of this exercise is to quantitatively measure the importance of the sentiment channel in the propagation of government spending shocks.

Following Bachmann and Sims (2012), we generate a hypothetical sequence of sentiment shocks that holds sentiment unchanged at all forecast horizons since the impact of the fiscal shock. Putting it differently, we eliminate the indirect output effects of the government spending shock through changes in sentiment that were triggered by the shock. In doing so, one can obtain the hypothetical direct fiscal spending shock effects on private activity.

Our model specification is more general than that of Bachmann and Sims (2012), $\mathbf{x}_t = [g_t \quad sent_t \quad y_t]'$, in the sense that our system comes with monetary control variables in $\mathbf{z}_t$. Furthermore, we allow sentiment to reflect innovations of all other variables by having $sent_t$ ordered last.[22]

Let $\bar{F}$ denotes the top-left 5 by 5 sub-matrix of the $5p$ by $5p$ companion matrix for the state-space representation. The h-period ahead impulse-response function of the i^{th} variable to the structural shock to the j^{th} variable is given by the following.

$$\psi_{i,j}(h) = s_i^{'} \bar{F}^{h-1} A_0^{-1} s_j, \tag{7}$$

where s_i is a 5 by 1 selection vector, for example, $s_5 = [0\ 0\ 0\ 0\ 1]^{'}$.

The contemporaneous sentiment response to a 1% government spending shock ($u_1^g = 1$) is given by $s_5^{'} A_0^{-1} s_1$. To cancel out this response, we generate the following hypothetical sentiment shock,

$$u_1^{sent} = -\frac{s_5^{'} A_0^{-1} s_1}{s_5^{'} A_0^{-1} s_5}, \tag{8}$$

We recursively calculate the sequence of sentiment shocks for the remaining period as follows.

$$u_h^{sent} = -\frac{s_5^{'} \bar{F}^{h-1} A_0^{-1} s_1 + \sum_{r=1}^{h-1} \left(s_5^{'} \bar{F}^{h-r} A_0^{-1} s_5 \right) u_r^{sent}}{s_5^{'} A_0^{-1} s_5}, \; h = 2, 3, ... \tag{9}$$

[22] However, we obtained qualitatively similar results with alternative ordering as in theirs. All results are available upon request.

Finally, we obtain the counterfactual impulse-response function of the i^{th} variable to the 1% fiscal spending shock by the following.

$$\hat{\psi}_{i,1}(h) = \psi_{i,1}(h) + \sum_{r=1}^{h} \left(s_i' \bar{F}^{h-r} A_0^{-1} s_5 \right) u_r^{sent}, \tag{10}$$

where $\psi_{i,1}(h)$ is the unrestricted h-period ahead impulse-response function.

Figure 9 reports estimated hypothetical response functions (solid lines) $\hat{\psi}_{i,1}(h)$. Dashed lines are unconstrained impulse-response functions $\psi_{i,1}(h)$ and their 90% confidence bands from 500 nonparametric bootstrap simulations. As can be seen at the bottom, sentiment responses are completely muted (solid lines) after we add the sequence of hypothetical sentiment shock in (8) and (9).

It should be noted that the structural shifts in the output responses we previously observed has disappeared after we control the endogenous responses of sentiment. Specifically,

1. The private GDP hardly responds to the government spending shock once we neutralize the indirect sentiment effect. That is, we observe virtually no direct effects of fiscal spending shocks on the private GDP ($pgdp_t$), implying that positive (or negative) output responses in the private sector have been driven by changes in sentiment.

2. The responses of the total GDP remain practically constant for all horizons, which means that qualitatively different responses of the total GDP over time were driven mainly by $pgdp_t$ responses. Furthermore, the total GDP responses reflect changes in the public sector components of the GDP only in the absence of meaningful responses of $pgdp_t$.

Putting it all together, we report substantial and important role of the sentiment channel in propagating fiscal policy shocks to private sector activity. And our simulation exercises demonstrate that changes in sentiment is the driving force of the time-varying effectiveness of fiscal policy.

Figure 9 around here

4 What Explains the Changes in Sentiment?

In this section, we provide statistical inferences about how market participants revise their economic prospects when they receive new information on fiscal actions. For this purpose, we investigate the time-varying relationship between GDP growth forecasts and government spending growth forecasts that are formulated by experts in the private sector, which helps understand the time-varying output effects of fiscal policy via the sentiment channel.

4.1 Understanding Dynamics of Sentiment through the Lens of the SPF

We first study how private agents revise their forecasts of real GDP growth when they update information on real government spending growth. For this purpose, we employ the Survey of Professional Forecasters (SPF) data for the period between $1968Q4$ and $2017Q3$. We are particularly interested in the relationship between the SPF forecasts of real GDP growth and those of real federal government spending growth.[23]

Let $\gamma_{x_t}^{SPF}(j+1) = x_{t+j}^{SPF} - x_{t-1}^{SPF}$ be the SPF growth rate *forecast* of (logged) x_t over $j+1$ quarters, while $\gamma_{x_t}(j+1) = x_{t+j} - x_{t-1}$ denotes the *realized* counterpart of $\gamma_{x_t}^{SPF}(j+1)$. We define the SPF forecast errors by $\hat{\gamma}_{x_t}(j+1) = \gamma_{x_t}^{SPF}(j+1) - \gamma_{x_t}(j+1)$. Note that we do not square or take absolute values of forecast errors as we do for loss functions, because the sign of the errors delivers important information. We first present the SPF forecast errors of real GDP growth over 5 quarters, $\hat{\gamma}_{y_t}(5)$, in Figure 10.[24] Some interesting observations are as follows.

Note that private forecasters tend to over-estimate the real GDP growth rate ($\hat{\gamma}_{y_t}(5) > 0$) during the pre-Volcker era ($1968Q4 - 1979Q2$), while they predominantly under-estimate it ($\hat{\gamma}_{y_t}(5) < 0$) during the post-Volcker era until the early 2000s. During the 2000s period, the SPF forecasts stay overall optimistic ($\hat{\gamma}_{y_t}(5) > 0$) until the beginning of the Great Recession, followed by much weaker but still optimistic forecasts.

We conjecture that these systemic forecast errors are closely related with the structural break in the monetary policy stance suggested by Clarida et al. (2000), who pointed out that the Federal Reserve's interest rate setting behavior has changed when Paul Volcker took office in the third quarter of 1979. To put it differently, they suggested that the monetary policy

[23]See the data description section for a detailed explanation on how these data are constructed.

[24]The vertical line is the break date, which is estimated by the structural break test presented in the next section.

stance had stayed overall accommodative during the pre-Volcker era, while the stance of monetary policy had turned hawkish during the post-Volcker era.

This implies that private forecasters formulated more optimistic GDP growth forecasts when monetary policy was coordinating well with fiscal policy during the pre-Volcker era. On the other hand, it seems that they have formulated more pessimistic GDP growth forecasts during the post-Volcker era when monetary policy stayed hawkish until the beginning of the 2000s. In the early 2000s, Greenspan has initiated an array of aggressive rate cuts to fight the recession triggered by the burst of the so-called dot com bubble in 2001. Such optimism in the early 2000s has become subdued rapidly when the Great Recession began in $2007-2008$, but has revived with an array of unconventional dovish monetary policy actions such as quantitative easing and operation twist.

Figure 10 around here

We investigate this possibility by examining the time-varying relationship between $\gamma_{g_t}^{SPF}(1)$ and $\gamma_{y_t}^{SPF}(5)$ via the following least squares (LS) regression over time using a fixed-size rolling window scheme.

$$\gamma_{y_t}^{SPF}(5) = \alpha + \beta\gamma_{g_t}^{SPF}(1) + \varepsilon_t \tag{11}$$

The motivation of this regression analysis is the following. When market participants receive new information on government spending growth, $\gamma_{g_t}^{SPF}(1)$, the realized (actual) patterns of revisions of their real GDP growth forecasts in the future, $\gamma_{y_t}^{SPF}(5)$, would reveal their view about the effectiveness of the government spending shock. That is, β is likely to be greater when forecasters are optimistic on the effect of fiscal stimulus. As forecasters become less optimistic or even pessimistic, β will decrease to zero or even become negative.

Figure 11 presents the LS estimates $\hat{\beta}_{LS}$ for β in (11) over time with a 44-quarter fixed size rolling window, that is, the initial point estimate corresponds to β from the pre-Volcker era. We also report the 90% confidence bands that are obtained from the normal approximation. This initial $\hat{\beta}_{LS}$ is 0.843, which is significant at the 5% level. However, the $\hat{\beta}_{LS}$ estimate rapidly declines as the sample window starts shifting towards the post-Volcker era. Note that confidence bands expand greatly since then, reflecting dramatic changes in $\hat{\beta}_{LS}$ after Mr. Volcker started extremely hawkish anti-inflation policies. This implies that market participants may formulate expectations of a lot weaker and statistically insignificant output effects of fiscal policy when the stance of monetary policy becomes hawkish.

The $\hat{\beta}_{LS}$ becomes stabilized eventually until it begins rising from the early 2000s, reflecting accommodative monetary policy actions implemented by Mr. Greenspan to fight the recession that began in 2001, followed by the burst of the so-called dot com (IT) bubble. Note that the $\hat{\beta}_{LS}$ point estimates remain overall high even during the Great Recession as the monetary policy becomes extremely accommodative with three rounds of quantitative easing (QE). However, the confidence bands become wider possibly reflecting high degree uncertainty and the fact that the role of monetary policy has become limited during the zero-lower-bound (ZLB) era. The $\hat{\beta}_{LS}$ starts falling around the mid 2010s when the Fed began the normalization plan of monetary policy. Putting all together, Figures 8 through 10 provide strong evidence of time-varying sentiment responses to the government spending shock through the lens of the SPF.

Figure 11 around here

4.2 Statistical Evidence of Structural Breaks

This subsection presents statistical evidence in favor of our conjectures presented in the previous section, which imply the presence of structural breaks in β. For this, we employ Bai and Perron (2003) and Hansen (1997, 2001)'s structural break tests for (11).[25]

When one tests the existence of a *single* structural break, we may consider the following alternative hypothesis, $H_A : \beta_1 \neq \beta_2$, where $\beta = \beta_1,\ t \in [1, \tau]$ and $\beta = \beta_2,\ t \in (\tau, T]$, which implies a break at time $t = \tau$. One may employ the following statistics proposed by Andrews (1993) using the full sample (T).

$$\text{Sup}F_T = \sup_{k_1 \leq k \leq k_2} F_T(k), \tag{12}$$

where $F_T(k)$ denotes the Lagrange Multiplier statistics, or alternatively the Wald or Likelihood Ratio test statistics, for the null hypothesis of no structural break, $H_0 : \beta_1 = \beta_2$, given a fine candidate grid point over $k \in [k_1,\ k_2]$.

Table 1 reports key results from the test procedures of Bai and Perron (2003), allowing a maximum of 3 structural breaks that are recommended for the case of roughly 10 years of the minimal length of a segment. $\text{Sup}F_T(k)$ is the test statistic for the null hypothesis

[25] The Chow test is not a feasible option because the structural break date is unknown.

of no structural break versus the alternative hypothesis of k breaks. The test rejects the null hypothesis favoring the alternative hypothesis of $k = 1$ and $k = 3$. Note that the two double maximum tests, $UD\max$ and $WD\max$, also reject the null of no break, favoring the existence of (unknown) number of structural breaks. $\text{Sup}F_T(k+1|k)$ is the sequential test of k breaks (null hypothesis) against the alternative hypothesis of $k+1$ breaks, which favors $k = 1$ and $k = 3$ against $k = 2$. It should be noted, however, that both the sequential and the BIC procedures select $k = 1$ using a 5% size. Given this, we conclude there exists a single structural break for (11) around the beginning of the post-Volcker era, $1978Q2$ with its 95% confidence band $[1974Q2, 1982Q3]$.

We further examine this possibility by employing the test procedure proposed by Hansen (1997, 2001), which is a sequential test procedure of a single break test utilizing the following test statistics as well as (12) that were proposed by Andrews (1993) and Andrews and Ploberger (1994).[26]

$$\text{Exp}F_T = \ln\left(\frac{1}{k_2 - k_1 + 1}\sum_{t=k_1}^{k_2}\exp\left(\frac{1}{2}F_T(k)\right)\right) \tag{13}$$

$$\text{Ave}F_T = \frac{1}{k_2 - k_1 + 1}\sum_{t=k_1}^{k_2}F_T(k)$$

Table 2 reports strong evidence of the presence of a structural break in (11). That is, the three statistics all reject the null hypothesis of no structural break with virtually zero p values. The $\text{Sup}F_T$ test procedure selects $1978Q2$ as the identified break date from the full sample. Following the sequential test approach by Hansen (2001), we seek additional break dates in the two sub-sample periods that are identified by the first structural break date, $1969Q4 - 1978Q2$ and $1978Q2 - 2017Q3$. However, we obtained no further evidence of a structural break in both the sub-sample periods even at the 10% significance level, concluding there was a single break in $1978Q2$.[27]

Table 1 and 2 around here

As can be seen in Tables 1 and 2, all tests support the presence of a single structural break from the full sample ($1969Q4 - 2017Q3$). Further, these procedures identified the break date

[26] We used conventional trimming parameter values, $k_1 = 0.15T$ and $k_2 = 0.85T$. p values are obtained using the method by Hansen (1997).

[27] Test results for the earlier period are available upon request.

at 1978Q2, which roughly corresponds to the beginning of the post-Volcker era. These test results highlight the implications of the systemic patterns of the forecast errors in Figure 9. During the pre-Volcker era that corresponds to regime D in our baseline theoretical model, market participants tend to over-estimate real GDP growth, whereas they formulate their forecasts more pessimistically during the post-Volcker era. These results are also confirmed by the expected output effects (β) of the government spending shock shown in Figure 10.

Putting all together, empirical findings presented in this section imply that the effectiveness of fiscal stimulus greatly hinges upon the coordination of monetary and fiscal policies through an important role of the sentiment channel. Hawkish monetary policy that conflicts with fiscal stimulus generates consumer pessimism, which ultimately weakens output effects by reducing private spending.

5 Concluding Remarks

Can increases in government spending help stimulate private activity? What variables play a dominant role in propagating government spending shocks to private spending? Empirical evidence is still at best mixed, and the profession has failed to reach a consensus.

Motivated by the work of Leeper et al. (2017), we present a New Keynesian macroeconomic model that yield strong output effects of fiscal stimulus only when monetary policy coordinates well with fiscal policy. When the central bank responds to inflation aggressively, private spending tends to fall in response to government spending shocks because the central bank raises the interest rate faster than inflation, resulting in an increase in the real interest rate.

Employing the post-war U.S. macroeconomic data, we confirm these predictions about the output effects of fiscal policy employing both the frequentist and the Bayesian models. During the pre-Volcker era, the private GDP rises as consumption increases rapidly in response to government spending shocks. Such strong stimulus effects rapidly disappear when the sample period moves toward the post-Volcker era. Although the empirical findings are overall consistent with theoretical predictions as to the output effects of fiscal policy, we observe a negligible role of the real interest rate in the propagation mechanism of fiscal stimulus to private spending, which is at odds with New Keynesian models.

The present paper proposes a sentiment channel as an alternative propagation mechanism. We demonstrate that sentiment leads innovations in consumption rather than passively reflecting changes in consumption. Employing the Survey of Professional Forecasters data, we show forecasters tend to make systemic forecast errors. More specifically, they tend to over-estimate (optimism) real GDP growth when monetary and fiscal policies coordinate well with each other. When policies conflict with each other, however, they often formulated more pessimistic forecasts. That is, they were prone to under-estimate economic growth in the near future.

We further investigate how forecasters revise their economic prospects when they receive new information on fiscal actions. Our regression analyses with a rolling window scheme demonstrate that positive innovations in government spending tend to trigger more optimistic GDP growth forecasts when monetary policy-makers maintain a dovish stance. When the central bank responds to inflation aggressively, however, forecasters are likely to formulate more pessimistic economic prospects. That is, fiscal stimulus under such circumstances generates pessimism that decreases private spending, ultimately weakening the output effects of fiscal policy. We corroborate our analyses by further providing statistical test results that confirm our claims.

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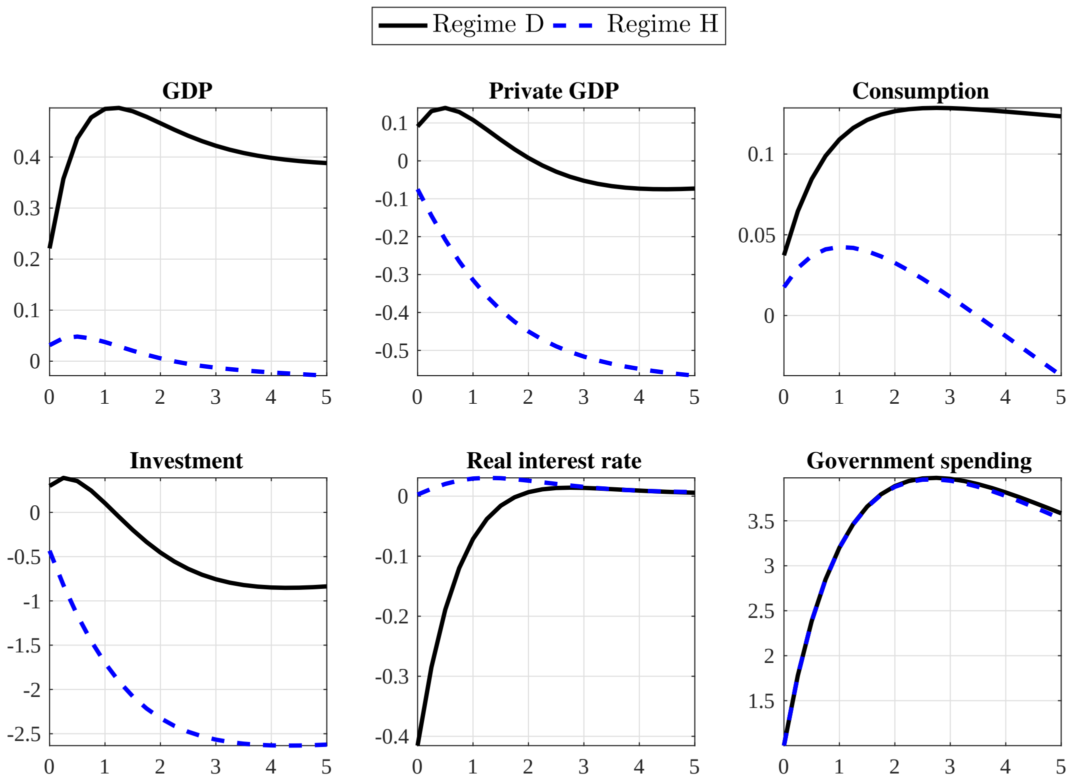
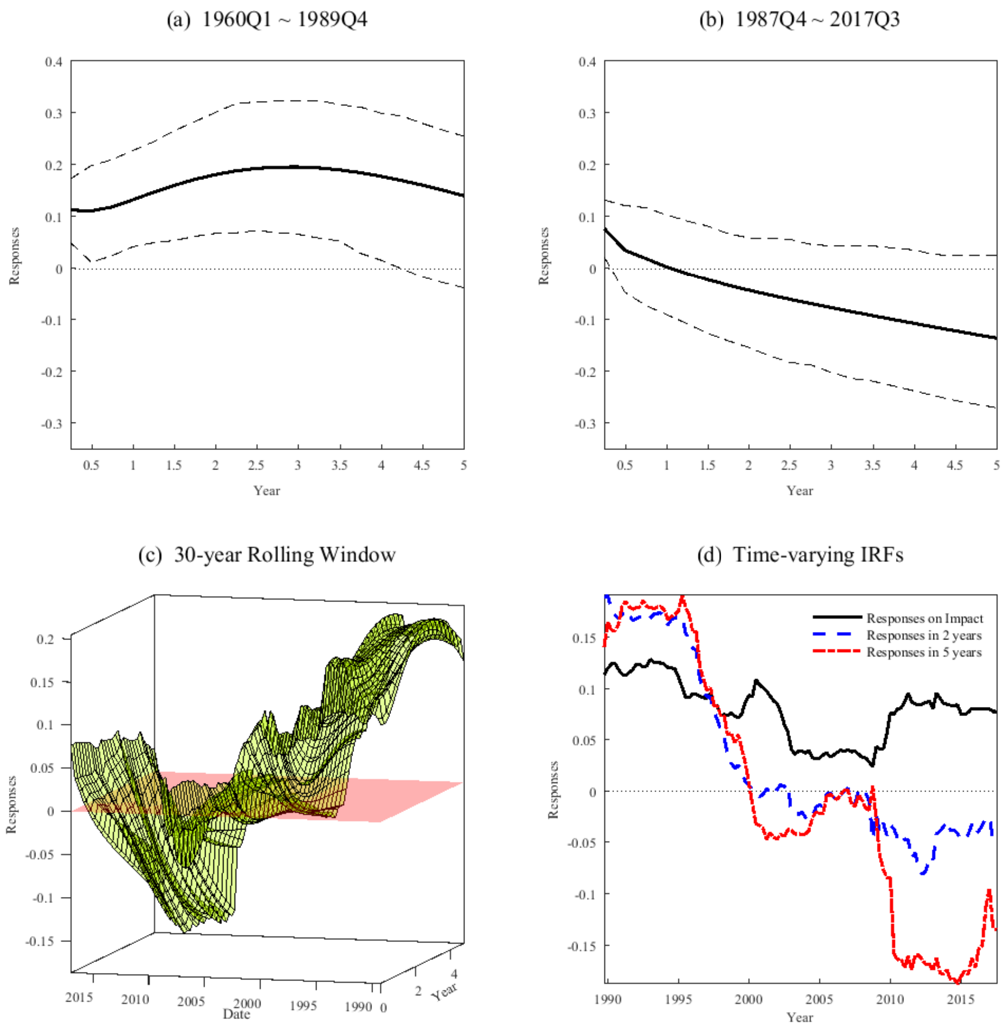
Figure 1. Simulated Impulse Responses to the Government Spending Shock



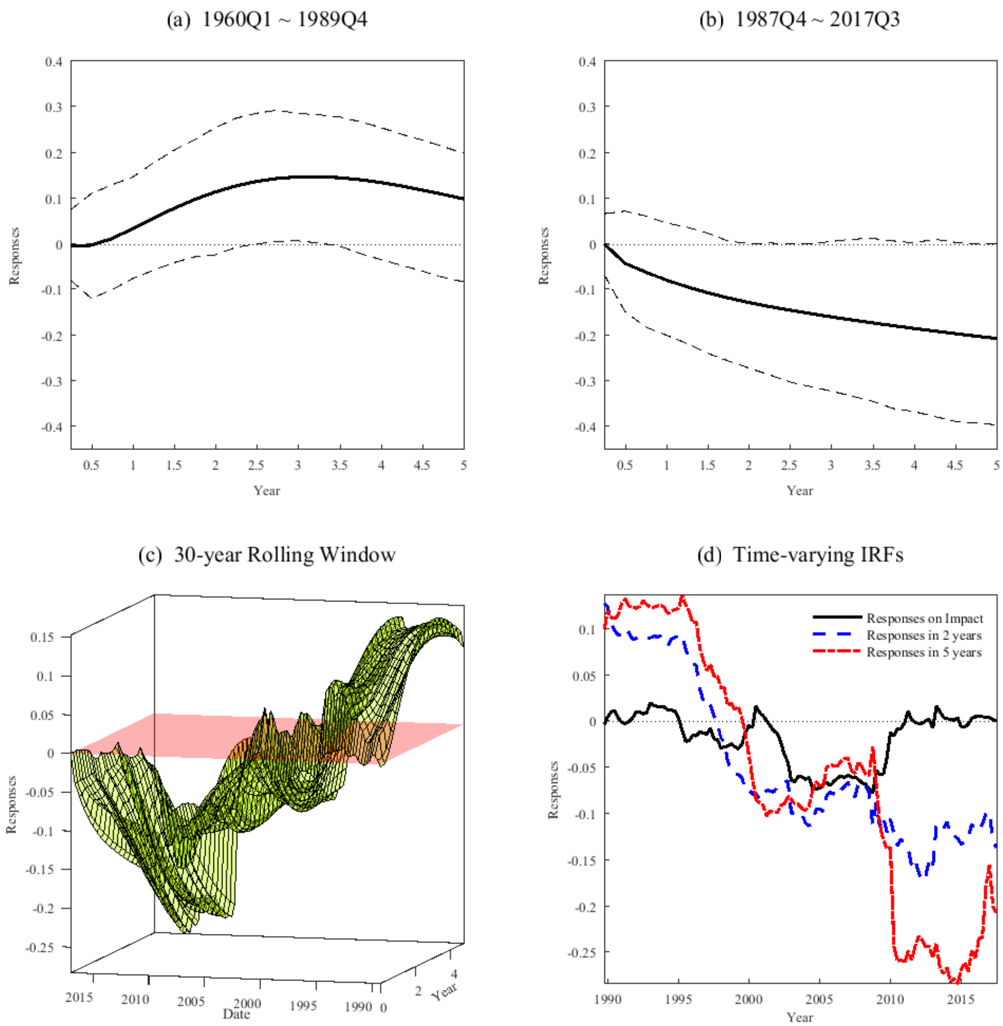
Note: We report simulated responses over 5 years to a 1% government spending shock in each regime. The monetary authority is assumed to maintain an accommodative stance that coordinates well with expansionary fiscal policy under the Regime D. On the other hand, the central bank maintains a hawkish policy stance that conflicts the dovish stance of the government under the Regime H.

Figure 2. GDP Responses to the Government Spending Shock



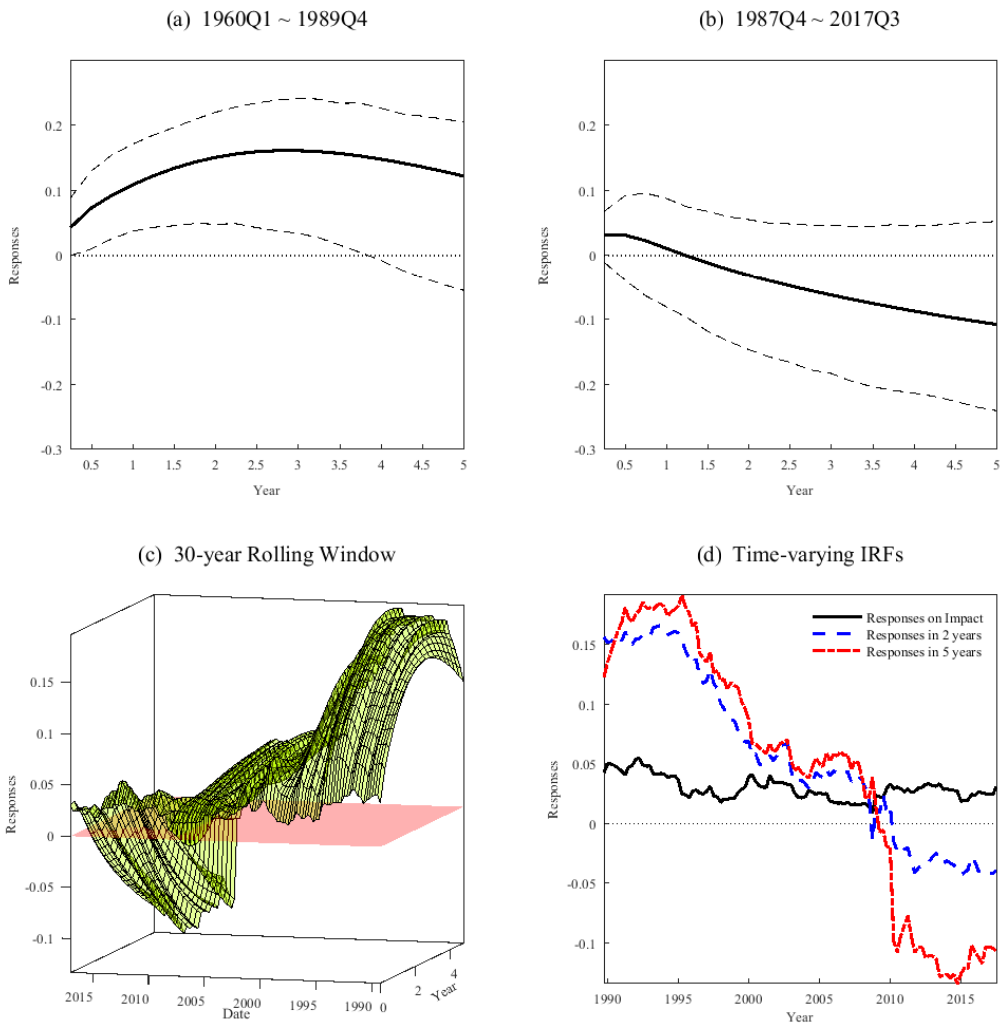
Note: We report the impulse-response function (IRF) estimates from $\mathbf{x}_t = [g_t, gdp_t, int_t, mon_t]'$ to a 1% government spending shock. Panel (a) and Panel (b) report the IRF estimates (solid) of the total GDP along with its 90% confidence bands (dashed) that were obtained from 500 bootstrap simulations with empirical distributions. Panel (c) reports an array of IRFs to the government spending shock with a 30-year fixed-size rolling window scheme. Panel (d) provides the IRFs in the short- to the long-run by dissecting the surface graph (panel (c)) at $y = 0,\ 2,\ 5$ (years) of the year-axis.

Figure 3. Private GDP Responses to the Government Spending Shock



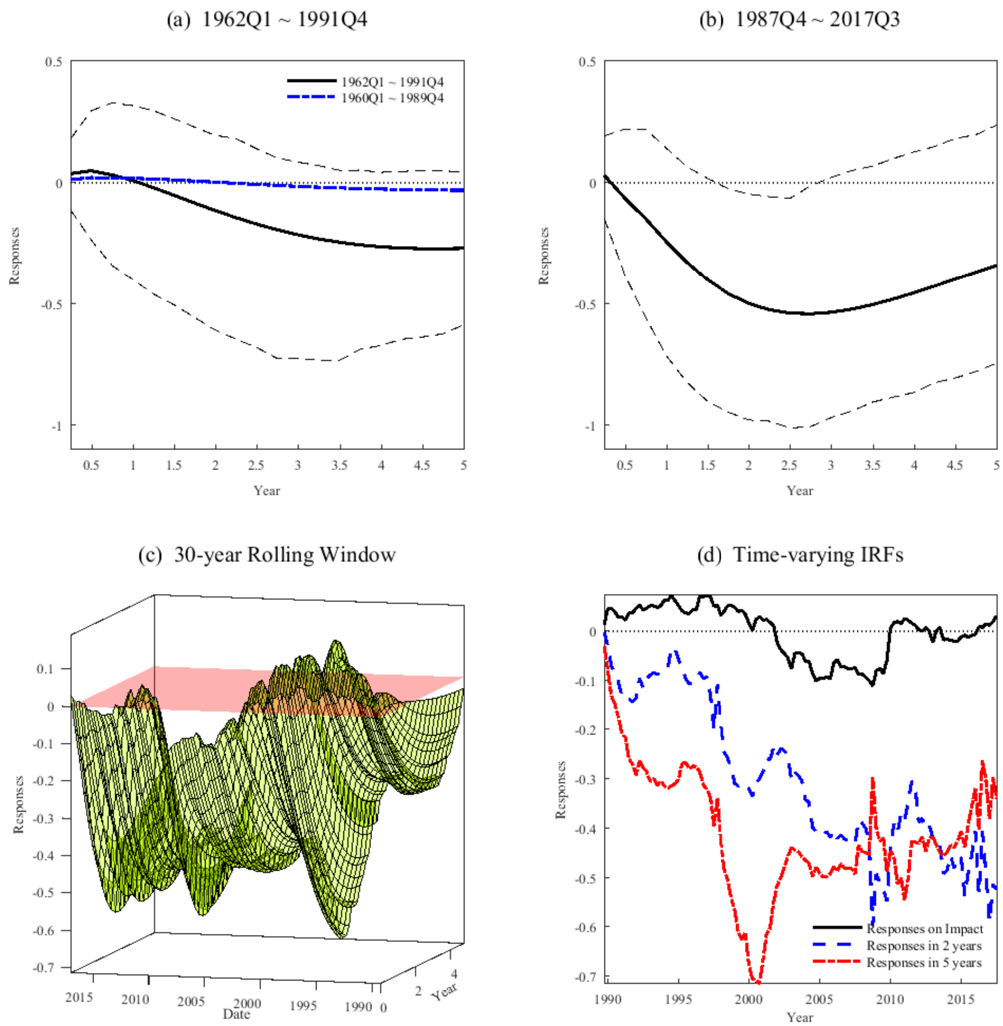
Note: We report the impulse-response function (IRF) estimates from $\mathbf{x}_t = [g_t, pgdp_t, int_t, mon_t]'$ to a 1% government spending shock. Panel (a) and Panel (b) report the IRF estimates (solid) of private GDP along with its 90% confidence bands (dashed) that were obtained from 500 bootstrap simulations with empirical distributions. Panel (c) reports an array of IRFs to the government spending shock with a 30-year fixed-size rolling window scheme. Panel (d) provides the IRFs in the short- to the long-run by dissecting the surface graph (panel (c)) at $y = 0,\ 2,\ 5$ (years) of the year-axis.

Figure 4. Consumption Responses to the Government Spending Shock



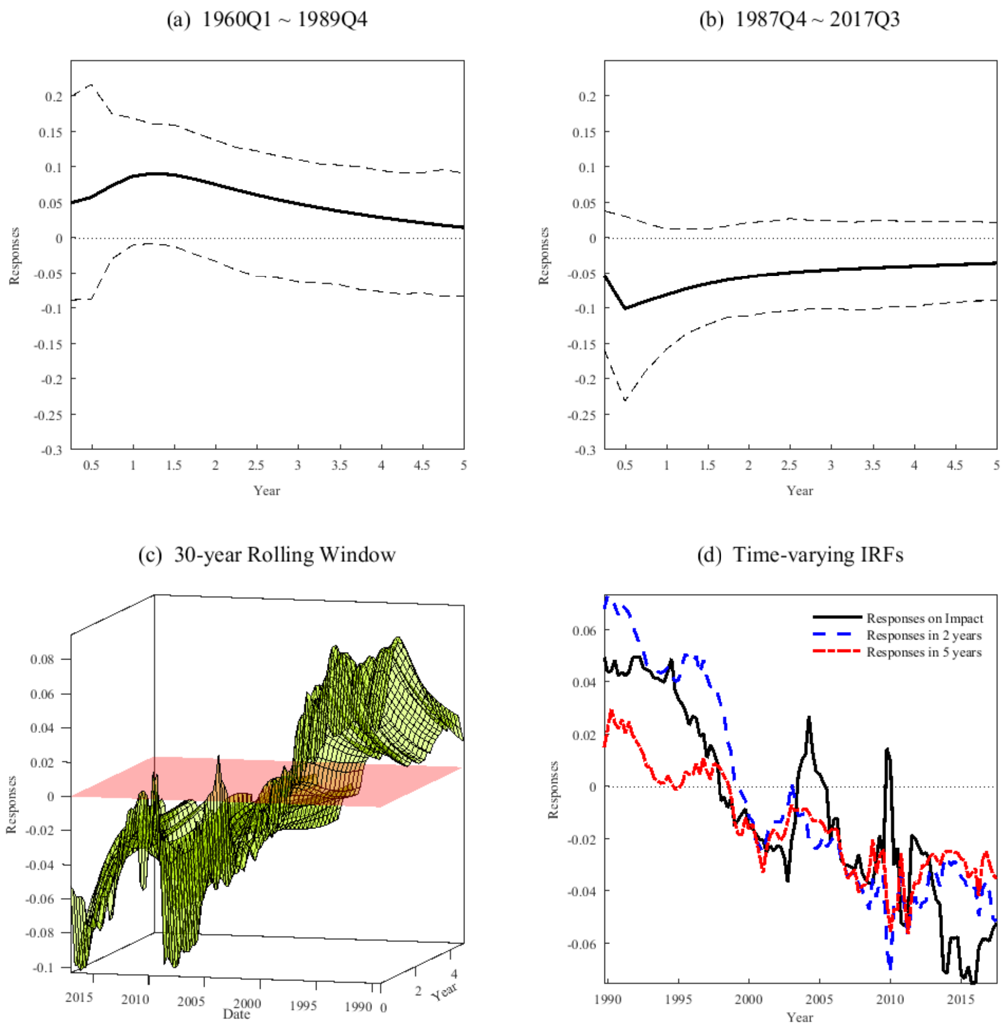
Note: We report the impulse-response function (IRF) estimates from $\mathbf{x}_t = [g_t, conm_t, int_t, mon_t]'$ to a 1% government spending shock. Panel (a) and Panel (b) report the IRF estimates (solid) of consumption along with its 90% confidence bands (dashed) that were obtained from 500 bootstrap simulations with empirical distributions. Panel (c) reports an array of IRFs to the government spending shock with a 30-year fixed-size rolling window scheme. Panel (d) provides the IRFs in the short- to the long-run by dissecting the surface graph (panel (c)) at $y = 0,\ 2,\ 5$ (years) of the year-axis.

Figure 5. Investment Responses to the Government Spending Shock



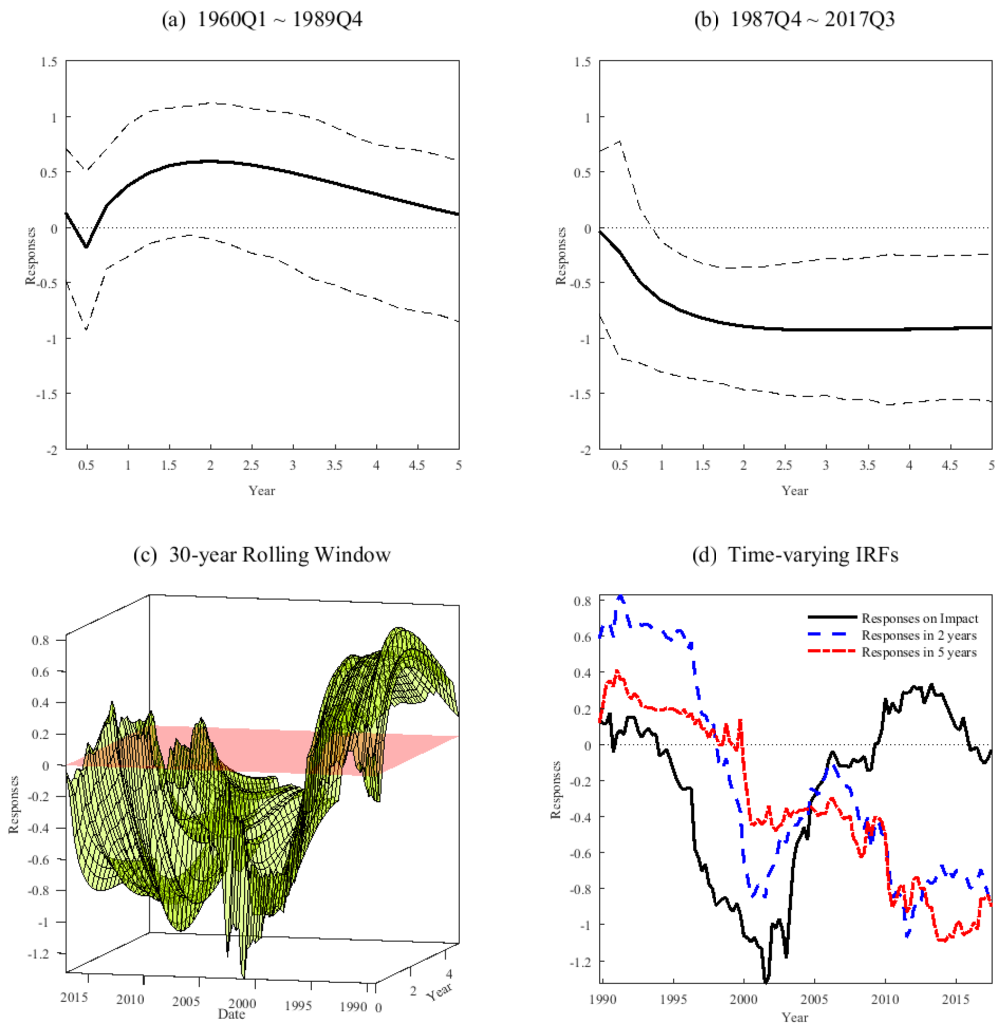
Note: We report the impulse-response function (IRF) estimates from $\mathbf{x}_t = [g_t, invt_t, int_t, mon_t]'$ to a 1% government spending shock. Panel (a) and Panel (b) report the IRF estimates (solid) of investment along with its 90% confidence bands (dashed) that were obtained from 500 bootstrap simulations with empirical distributions. Panel (c) reports an array of IRFs to the government spending shock with a 30-year fixed-size rolling window scheme. Panel (d) provides the IRFs in the short- to the long-run by dissecting the surface graph (panel (c)) at $y = 0,\ 2,\ 5$ (years) of the year-axis.

Figure 6. Real FFR Responses to the Government Spending Shock



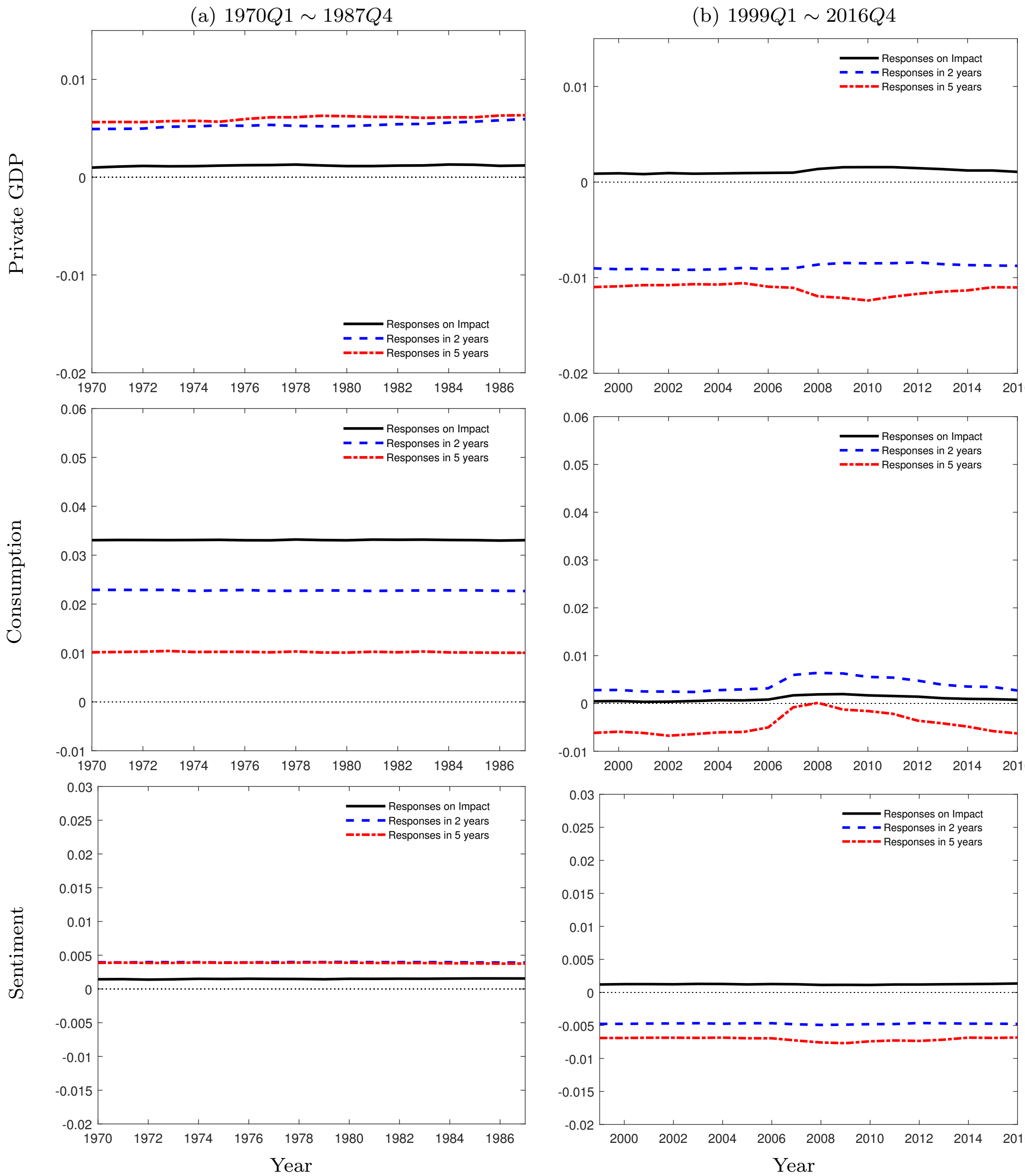
Note: We report the impulse-response function (IRF) estimates from $\mathbf{x}_t = [g_t, gdp_t, rint_t, mon_t]'$ to a 1% government spending shock. Panel (a) and Panel (b) report the IRF estimates (solid) of real interest rate along with its 90% confidence bands (dashed) that were obtained from 500 bootstrap simulations with empirical distributions. Panel (c) reports an array of IRFs to the government spending shock with a 30-year fixed-size rolling window scheme. Panel (d) provides the IRFs in the short- to the long-run by dissecting the surface graph (panel (c)) at $y = 0,\ 2,\ 5$ (years) of the year-axis.

Figure 7. Sentiment Responses to the Government Spending Shock



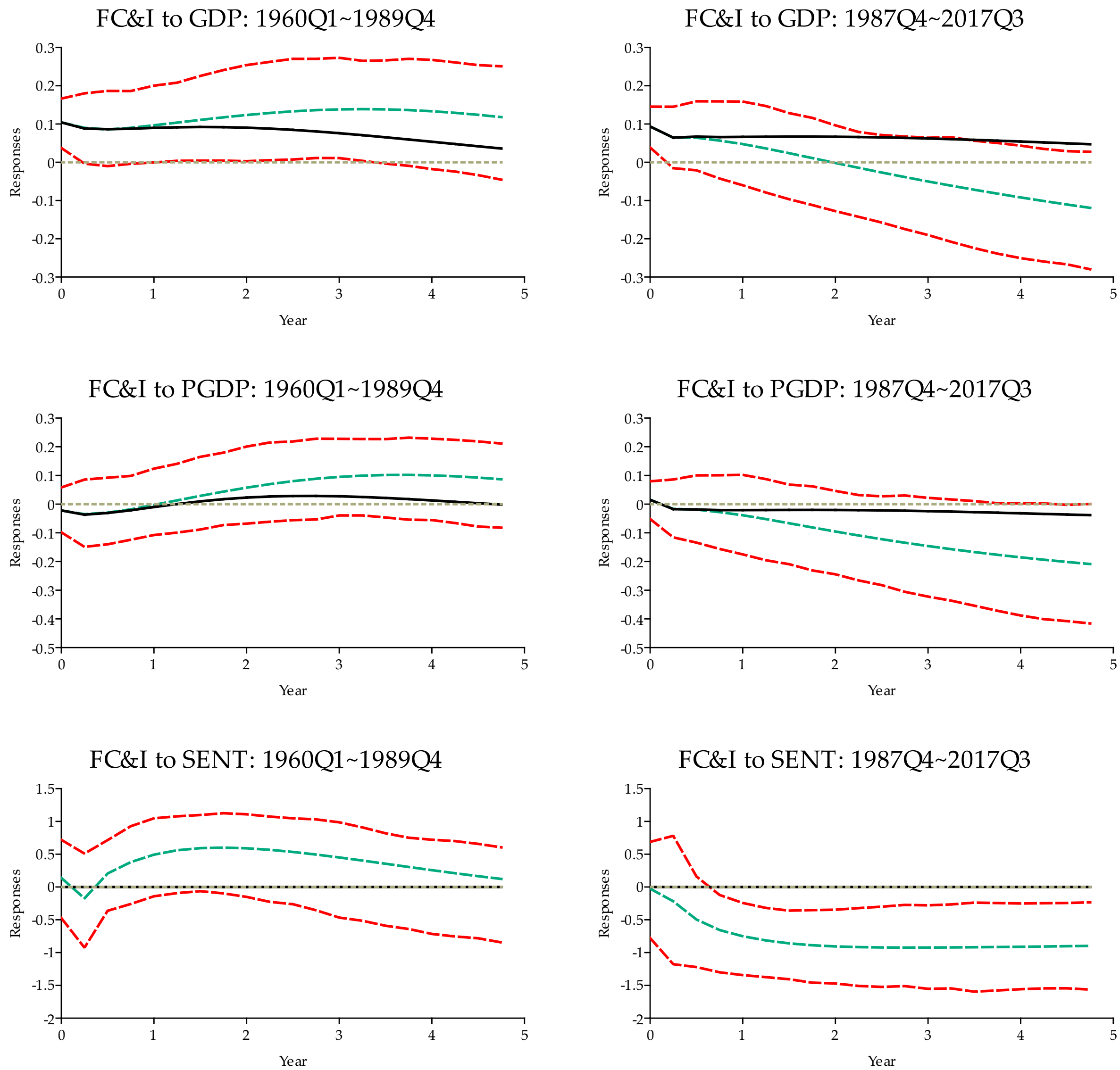
Note: We report the impulse-response function (IRF) estimates from $\mathbf{x}_t = [g_t, gdp_t, rint_t, mon_t, sent_t]'$ to a 1% government spending shock. Note that the location of $sent_t$ is irrelevant given that g_t is placed first. Panel (a) and Panel (b) report the IRF estimates (solid) of sentiment along with its 90% confidence bands (dashed) that were obtained from 500 bootstrap simulations with empirical distributions. Panel (c) reports an array of IRFs to the government spending shock with a 30-year fixed-size rolling window scheme. Panel (d) provides the IRFs in the short- to the long-run by dissecting the surface graph (panel (c)) at $y = 0,\ 2,\ 5$ (years) of the year-axis.

Figure 8. Bayesian Impulse Response Estimates in the Short- to the Long-Run



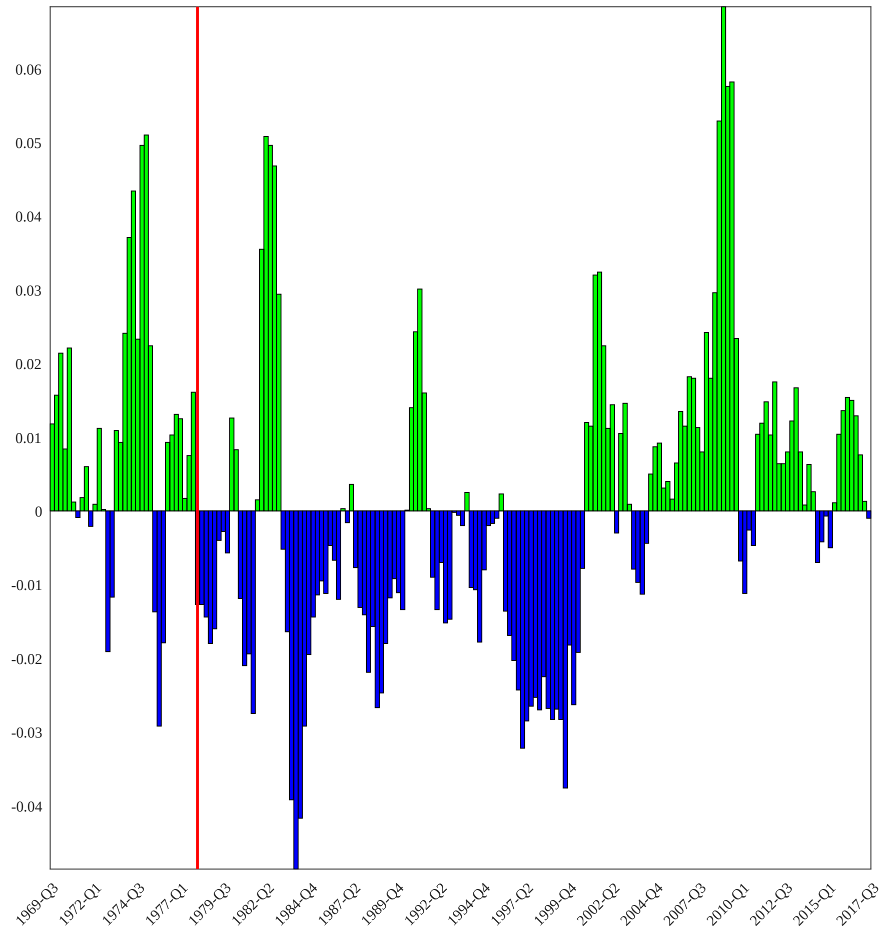
Note: We report the impulse-response function (IRF) estimates with a Bayesian approach for private GDP ($pgdp_t$), consumption ($conm_t$), and sentiment ($sent_t$) to a 1% government spending shock. First 10-year observations were used to form the prior for each of the two subsamples. All IRFs are available from authors upon requests.

Figure 9. Counterfactual Simulation Exercises with Alternative Identification Scheme



Note: We report the impulse-response function (IRF) estimates from 5-variate VAR models with $sent_t$ ordered last to a 1% government spending shock. Solid lines are hypothetical response functions with additional sentiment shocks that are designed to hold sentiment unchanged for all forecast horizons. Dashed lines are the impulse-response function point estimate from unconstrained models accompanied by its 90% confidence bands.

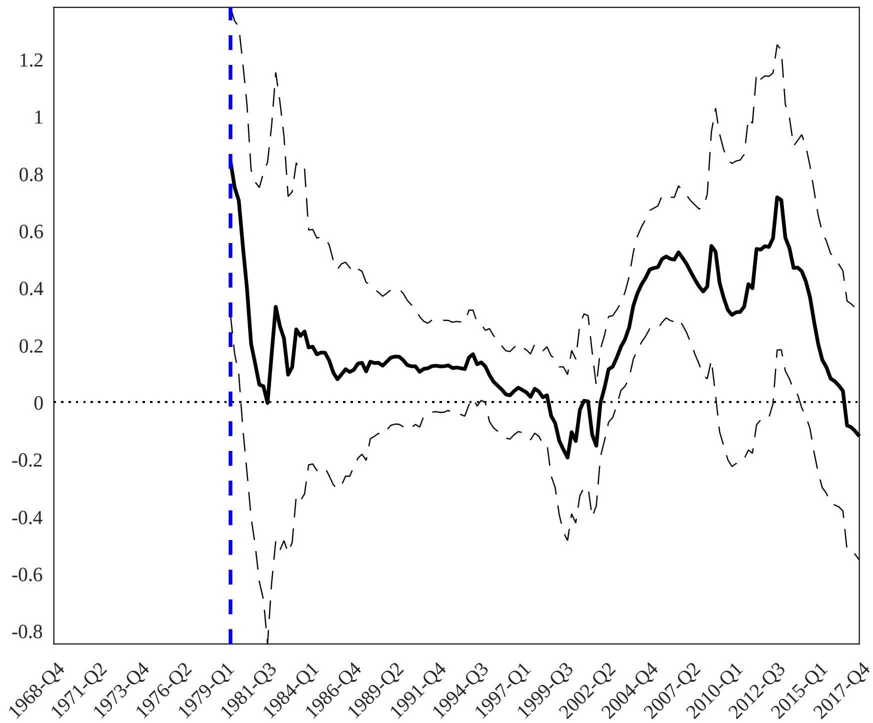
Figure 10. SPF Forecast Errors for the Real GDP Growth Rate



Note: We report the 5-quarter ahead SPF forecast errors for the real GDP growth rate. The vertical line represents the break date estimate from SupF_T test.

Figure 11. LS Estimates for β with a Fixed Size Rolling Window Scheme

$$\gamma_{y_t}^{SPF}(5) = \alpha + \beta\gamma_{g_t}^{SPF}(1) + \varepsilon_t$$



Note: We report the LS estimates $\hat{\beta}_{LS}$ for β over time with a 44-quarter fixed size rolling window so that the initial point estimate corresponds to the pre-Volcker era ($1968Q4 \sim 1979Q3$). We obtained the 90% confidence bands (dashed lines) via the normal approximation.

Table 1. Bai and Perron's Structural Break Tests for β in $\gamma_{y_t}^{SPF}(5) = \alpha + \beta\gamma_{g_t}^{SPF}(1) + \varepsilon_t$

Specifications[1]				
$z_t = \{1, \gamma_{g_t}^{SPF}\}$	$q = 1$	$p = 0$	$h = 39$	$M = 3$
Tests				
Sup$F_T(1)$	Sup$F_T(2)$	Sup$F_T(3)$	$UDmax$	$WDmax$
$9.4918^\dagger$	4.8099	$13.8476^\dagger$	$13.8476^\dagger$	$13.8476^\dagger$
Sup$F_T(2\|1)$	Sup$F_T(3\|2)$			
2.7656	$28.2002^\dagger$			
Number of breaks Selected				
Sequential	1			
BIC[2]	1			
Estimates for break dates selected by the sequential procedure				
$\hat{T}_1$				
1978Q2				
$(1974Q2 - 1982Q3)$				

[1] See Bai and Perron (2003): q is the number of regressors whose coefficient could break; p is the number of regressors whose coefficient does not break; h is the minimal length of a segment; M is the maximum number of breaks allowed.
[2] BIC represents the Bayesian Information Criterion.

Table 2. Hansen's Structural Break Tests for β in $\gamma_{y_t}^{SPF}(5) = \alpha + \beta\gamma_{g_t}^{SPF}(1) + \varepsilon_t$

		Test stat.		
Sample Period	Break Date	SupF_T	ExpF_T	AveF_T
$1968Q4 \sim 2017Q4$	**1978Q2**	19.19	6.18	7.17
		(0.00)	(0.00)	(0.00)
$1978Q2 \sim 2017Q4$	N/A	5.64	1.56	2.60
		(0.46)	(0.29)	(0.24)

Appendix

A The Theoretical Model

A.1 Firms and Price Setting

The final good (y_t) is a composite good of a continuum of intermediate goods (y_{it}), characterized by a Dixit and Stiglitz (1977) aggregator, $y_t = \left[\int_0^1 y_{it}^{(\theta_p-1)/\theta_p} di\right]^{\theta_p/(\theta_p-1)}$, where $\theta_p > 1$ governs the degree of substitution between the inputs. Taken input prices (P_{it}) and the output price $P_t = \left(\int_0^1 P_{it}^{1-\theta_p} di\right)^{1/1-\theta_p}$ as given, the profit maximization yields the demand for intermediate good i, $y_{it} = (P_{it}/P_t)^{-\theta_p} y_t$. The intermediate good i is produced by a monopolistically competitive firm who has the following production function:

$$y_{it} = (k_{it}^s)^{\alpha} n_{it}^{1-\alpha}, \tag{1}$$

where $\alpha \in (0,1)$. n_{it} and k_{it}^s denote the level of labor hours and capital services used by firm i, respectively.[1]

Each monopolistically competitive firm solves a two-stage problem. In the first stage, taken input prices (w_t and r_t^k), as given, each firm rents labor (n_{it}) and capital (k_{it}^s) to minimize its operating cost, $w_t n_{it} + r_t^k k_{it}^s$, subject to its production function (1). Cost minimization yields the identical real marginal cost:

$$mc_t = \sigma w_t^{1-\alpha} \left(r_t^k\right)^{\alpha}, \tag{2}$$

where $\sigma = \left(\frac{1}{1-\alpha}\right)^{1-\alpha} \left(\frac{1}{\alpha}\right)^{\alpha}$. In the second stage, each intermediate goods firm chooses its price (P_{it}) to maximize the discounted present value of future profits subject to the demand for y_{it}.

Following the price-setting scheme proposed by Calvo (1983), intermediate firm i can reset its price (P_{it}^*) with a fixed probability $(1-\omega_p)$. With probability ω_p, it partially indexes its price to past inflation according to the following rule:

$$P_{it} = \pi_{t-1}^{\iota_p} \bar{\pi}^{1-\iota_p} P_{it-1}, \tag{3}$$

where $\pi_t \equiv \frac{P_t}{P_{t-1}}$ is the gross inflation rate between $t-1$ and t, while $\bar{\pi}$ is the steady state inflation. Note that indexation is controlled by the parameter $\iota_p \in [0,1]$ that allows any combinations of the two types of indexation usually employed in the literature, steady state inflation (e.g., Yun (1996)) and the past inflation rate (e.g., Christiano et al. (2005)). Throughout this paper, variables with a bar denote steady state values.

[1] k_{it}^s is the effective amount of capital, which is introduced in the next section.

The profit maximization problem for firm i that reoptimizes its price at time t is:

$$\max_{P_{it}^*} \; E_t \sum_{s=0}^{\infty} (\omega_p \beta)^s \, \frac{\lambda_{t+s}}{\lambda_t} \left[\left(\frac{\Xi_{t,s}^p P_{it}^*}{P_{t+s}} - mc_{t+s} \right) y_{it+s} \right] \tag{4}$$

$$s.t. \quad y_{it+s} = \left(\frac{\Xi_{t,s}^p P_{it}^*}{P_{t+s}} \right)^{-\theta_p} y_{t+s}$$

$$\Xi_{t,s}^p = \left\{ \begin{array}{cc} 1 \quad \text{for} \quad s = 0 \\ \prod_{k=1}^{s} \pi_{t+k-1}^{\iota_p} \bar{\pi}^{1-\iota_p} \quad \text{for} \quad s \geq 1 \end{array} \right\},$$

where the profit at time $t+s$ is discounted by the pricing kernel $\beta^s \left(\lambda_{t+s}/\lambda_t\right)$ and λ_t is the marginal utility (or shadow price) of wealth of households at time t that appears in the following subsection. The optimality condition from (4) implies:

$$E_t \sum_{s=0}^{\infty} (\omega_p \beta)^s \, \frac{\lambda_{t+s}}{\lambda_t} \left[\frac{\Xi_{t,s} P_{it}^*}{P_{t+s}} - \mathcal{M}^p mc_{t+s} \right] y_{it+s} = 0 \tag{5}$$

where $\mathcal{M}^p \equiv \frac{\theta_p}{\theta_p - 1}$. The aggregate price index evolves as follows:

$$1 = (1 - \omega_p) \left(\pi_t^* \right)^{1-\theta_p} + \omega_p \left(\frac{\pi_{t-1}^{\iota_p} \pi^{1-\iota_p}}{\pi_t} \right)^{1-\theta_p} \tag{6}$$

where $\pi_t^* = \frac{P_t^*}{P_t}$.

A.2 Households and Wage Setting

There is a continuum of households on the unit interval $[0,\ 1]$ indexed by j. In addition to hours worked (n_{jt}), each household j derives utility from composite consumption (c_{jt}^*) which consists of private goods (c_{jt}) and public goods (g_t), that is, $c_{jt}^* \equiv c_{jt} + \alpha_g g_t$. Parameter α_g governs the degree of substitutability/complementarity of the consumption goods. When $\alpha_g < 0$, private and public consumption are complements (Leeper et al. (2017)), whereas $\alpha_g > 0$ implies that these are substitutes with each other (Christiano and Eichenbaum (1992); Ambler and Paquet (1996); Finn (1998)). Household j maximizes the following lifetime utility,

$$E_t \sum_{t=0}^{\infty} \beta^t \left[\ln \left(c_{jt}^* - h c_{t-1}^* \right) - \chi \frac{n_{jt}^{1+\eta}}{1+\eta} \right], \tag{7}$$

where $\beta \in (0,1)$ is the discount factor and $h \in (0,1)$ denotes the external habit formation parameter. To put it differently, we define the habit stock by a fraction of lagged aggregate consumption (hc_{t-1}^*). χ is the disutility parameter from work and $1/\eta$ determines the Frisch elasticity of labor supply.

The household's real flow budget constraint is given by:

$$c_{jt} + i_{jt} + \frac{B_{jt}}{P_t} \leq \frac{R_{t-1} B_{jt-1}}{P_t} + (1 - \tau^n) \, w_{jt} n_{jt} + \left[\left(1 - \tau^k \right) r_t^k u_{jt} - a \left(u_{jt} \right) \right] k_{jt-1} + d_{jt} + tr \tag{8}$$

where the left-hand side represents the uses of income, private consumption (c_{jt}), investment (i_{jt}), and purchases of nominal government debt (B_{jt}) deflated by P_t. The right-hand side denotes the sources of income consisting of real interest payments of government debt, after-tax real wage (w_{jt}) and capital rental (r_t^k) income, dividends distributed by the intermediate goods firms (d_{jt}), and constant lump-sum transfer payments (tr) from the government. τ^n and τ^k are constant tax rates levied on labor income and capital, respectively.[2]

The effective amount of capital services is represented by $k_{jt}^s \equiv u_{jt}k_{jt-1}$, whereas $a(u_{jt})k_{jt-1}$ describes the physical cost associated with variations in the degree of capacity utilization, which is parameterized by a quadratic function, $a(u_{jt}) = \zeta_1 (u_{jt} - 1) + \frac{\zeta_2}{2}(u_{jt} - 1)^2$. [3] Note that $u = 1$ and $a(1) = 0$ in the steady state. We also define $\frac{a''(1)}{a'(1)} \equiv \frac{\zeta_2}{1-\zeta_2}$ following Smets and Wouters (2007).[4]

The law of motion for capital is:

$$k_{jt} = (1-\delta) k_{jt-1} + \left[1 - S\left(\frac{i_{jt}}{i_{jt-1}}\right)\right] i_{jt}, \tag{9}$$

where δ is the depreciation rate and $S(\cdot)$ denotes an adjustment cost function, proposed by Christiano et al. (2005), such that $S(1) = S'(1) = 0$, and $\kappa \equiv S''(1) > 0$.

There is a representative, competitive labor agency that hires a continuum of differentiated labor from each household with the following aggregator:

$$n_t = \left[\int_0^1 n_{jt}^{\frac{\theta_w - 1}{\theta_w}} dj\right]^{\frac{\theta_w}{\theta_w - 1}}, \tag{10}$$

where $0 \leq \theta_w < \infty$ is the elasticity of substitution among different types of labor. This competitive labor agency maximizes its profit subject to this production function, taking all differentiated labor wages (w_{jt}) and the aggregate wage (w_t) as given, yielding:

$$n_{jt} = \left(\frac{w_{jt}}{w_t}\right)^{-\theta_w} n_t, \tag{11}$$

where w_t is the aggregate real wage that satisfies $w_t = \left(\int_0^1 w_{jt}^{1-\theta_w} dj\right)^{\frac{1}{1-\theta_w}}$.

Following Erceg et al. (2000), wage stickiness is introduced in a way that is analogous to price stickiness described above. In each period, a fraction $1 - \omega_w$ of households can adjust their wages to w_{jt}^* and others can only index their wages by past inflation as $w_{jt} = \pi_{t-1}^{\iota_w} \bar{\pi}^{1-\iota_w} w_{jt-1}$, where indexation is controlled by the parameter $\iota_w \in [0, 1]$. Therefore, the

[2] We assume constant tax rates to focus mainly on the transmission channel of government spending given the tax policy.

[3] Note that we use the *end of period stock* timing convention. For example, k_{t-1} is the capital stock that was determined by investment at time $t-1$, but is used at time t in the production function for y_t.

[4] We need this condition to linearize the model presented here.

wage-setting problem of households who reset their wages at time t can be written as:

$$\max_{w^*_{jt}} \; E_t \sum_{s=0}^{\infty} (\omega_w \beta)^s U(c_{jt+s}, n_{jt+s}) \tag{12}$$

$$s.t. \quad n_{jt+s} = \left(\frac{\Xi^w_{t,s} w^*_{jt}}{w_{t+s}} \right)^{-\theta_w} n_{t+s}$$

$$\Xi^w_{t,s} = \left\{ \begin{array}{cc} 1 \quad \text{for} \quad s=0 \\ \prod_{k=1}^{s} \pi^{\iota_w}_{t+k-1} \bar{\pi}^{1-\iota_w} \quad \text{for} \quad s \geq 1 \end{array} \right\}$$

The first order condition associated with this wage-setting problem can be written as:

$$\sum_{s=0}^{\infty} (\omega_w \beta)^s E_t \left[\frac{n_{jt+s}}{\tilde{c}_{t+s}} \left(\frac{\Xi^w_{t,s} w^*_{jt}}{P_{t+s}} - \mathcal{M}^w MRS_{jt+s} \right) \right] = 0 \tag{13}$$

where $\mathcal{M}^w \equiv \frac{\theta_w}{\theta_w - 1}$, $\tilde{c}_{t+s} \equiv c^*_{t+s} - hc^*_{t+s-1}$, and $MRS_{jt+s} \equiv \varrho \tilde{c}_{t+s} n^{\eta}_{jt+s}$ is the relevant marginal rate of substitution between consumption and labor hours in period $t+s$. Therefore, the aggregate wage index is described as follows:

$$1 = (1 - \omega_w) \left(\pi^*_{w,t} \right)^{1-\theta_w} + \omega_w \left(\frac{\pi^{\iota_w}_{t-1} \pi^{1-\iota_w}}{\pi_t} \frac{w_{t-1}}{w_t} \right)^{1-\theta_w} \tag{14}$$

where $\pi^*_{w,t} = \frac{w^*_t}{w_t}$.

A.3 Monetary and Fiscal Authorities

The monetary policy follows a Taylor rule. It adjusts the gross nominal interest rate (R_t) in response to deviations of inflation (π_t) and output (y_t) from their respective steady state levels:

$$R_t = R^{\psi_r}_{t-1} \left[\bar{R} \left(\frac{\pi_t}{\bar{\pi}} \right)^{\phi_\pi} \left(\frac{y_t}{\bar{y}} \right)^{\phi_y} \right]^{1-\psi_r} \tag{15}$$

where $0 \leq \psi_r < 1$ is the interest rate smoothing parameter, $\bar{R}$ is the equilibrium real interest rate, ϕ_π and ϕ_y are the policy response parameters to the inflation gap and the output gap, respectively.

The government collects tax revenues from capital and labor in addition to its sales of one-period debt to finance its expenditures that include interest payments, government expenditures (g_t) and transfer payments (tr). The government's flow budget constraint is:

$$\frac{B_t}{P_t} + \tau^n w_t n_t + \tau^k r^k_t u_t k_{t-1} = \frac{R_{t-1} B_{t-1}}{P_t} + g_t + tr \tag{16}$$

Government expenditures (g_t) obey the following stochastic process:

$$g_t = g^{\psi_g}_{t-1} \left[\bar{g} \left(b_{t-1}/\bar{b} \right)^{-\gamma_g} \right]^{1-\psi_g} \nu_{g,t} \tag{17}$$

where the parameter $\psi_g \in (-1,1)$ governs the degree of the persistence of g_t. Following Leeper et al. (2017), we allow government spending to respond to deviations of the (lagged) real debt $b_{t-1} = \frac{B_{t-1}}{P_{t-1}}$ from its stead state value $\bar{b}$. That is, the parameter $\gamma_g > 0$ triggers a correction of government spending when real debt deviates from its steady state value. $\nu_{g,t}$ is a government spending shock, which is assumed to follow a stationary ($\rho_g < 1$) AR(1) process:

$$\ln \nu_{g,t} = \rho_g \ln \nu_{g,t-1} + \sigma_g \varepsilon_{g,t}, \varepsilon_{g,t} \sim \mathbb{N}(0,1) \tag{18}$$

A.4 Market Clearing and Aggregation

We consider a symmetric equilibrium in which all intermediate good firms make identical choices so that the subscript i can be omitted. All goods and asset markets clear in the equilibrium. Specifically, the goods market clear condition requires the following aggregate resource constraint:

$$y_t = c_t + i_t + g_t + a(u_t)\,k_{t-1}, \tag{19}$$

where capital evolves according to the law of motion for capital (9). Equilibrium conditions and their log-linearized equivalents around the deterministic steady state are given in the Appendix. The log-linearized model is solved using the Sims (2002) gensys algorithm.

B Calibrations

The model is calibrated at a quarterly frequency. Regime specific monetary policy parameters are based on the estimates reported by Clarida et al. (2000) to investigate the effects of structural breaks in the Fed's behavioral equation. Other model parameters are along the lines of research works in the literature or were calibrated using U.S. data over the period $1960Q1 - 2017Q3$. Benchmark calibration parameter values are summarized in Table 1.

The discount factor (β) is set to 0.9958, which equals $(1/T)\sum_{t=1}^{T} \pi_t/(1+(FFR_t/100))^{1/4}$ where T is the sample size from the data, π_t denotes the quarterly gross inflation rate, and FFR_t is the effective federal funds rate. The inverse Frisch labor supply elasticity ($1/\eta$) is fixed at 2, which is a common value in the current literature. We set $\delta = 0.025$ for the quarterly depreciation rate for capital that implies an annual depreciation rate of 10%. The disutility parameter (χ) is an implied parameter that is calibrated with other parameters so that hours worked in the steady state is close to 1/3 in a model with divisible labor.[5] The habit formation coefficient (h) and the complementarity parameter (α_g) of consumption between private goods and public goods are set to 0.99 and -0.2, respectively, which are similar to the ones in Leeper et al. (2017).

The Cobb-Douglas factor share of capital (α) is set to 0.33. The price elasticity of demand for individual good (θ_p) and the elasticity of substitution among different types of labor (θ_w) are all calibrated to be 8. The capital utilization rate (ζ_2) and the adjustment cost

[5] This roughly matches the observation that individuals spend 1/3 of their time engaged in market activities and 2/3 of their time in non-market activities. See Hansen (1985).

for investment (κ) are set to 0.15 and 5, respectively, being consistent with the estimation results in Leeper et al. (2017). The parameters for price stickiness (ω_p) and wage stickiness (ω_w) are both assumed to be 0.8, implying a slightly over one-year average duration of price/labor contracts.

Monetary and fiscal parameters are calibrated based on the mean values from U.S. data over the same sample period in the present paper. The steady state gross quarterly inflation rate ($\bar{\pi}$) is assumed to be 1.0082. The total government spending-to-GDP ratio (s_g) is set to 0.0945. The government debt-to-GDP ratio (s_b) is 1.3707. The persistence parameter (ρ_g) of government spending is assumed to be 0.98. The average labor tax rate (τ^n) is set to 0.2171 and the capital tax rate (τ^k) is 0.2497.

In regime D, the (long run) coefficients on inflation (ϕ_π) and on the output (ϕ_y) are set to 0.83 and 0.27, respectively, while the interest rate smoothing parameter (ψ_r) is assumed to be 0.68. These values are based on the work of Clarida et al. (2000) for the pre-Volcker era that ends right before Paul Volcker took office as the new Federal Reserve chairman in 1979Q3. Government spending is assumed not to respond to the government debt, that is, γ_g is set to 0 implying that the fiscal authority also implements their stimulus policies aggressively.

For the specification of regime H, we employ the parameter values from Clarida et al. (2000) for the post-Volcker era. That is, we set ϕ_π, ϕ_y, and ψ_r to 2.15, 0.93, and 0.79, respectively. The fiscal authority in regime H maintains a less dovish stance than its stance in regime D, implementing mildly expansionary fiscal policy with $\gamma_g = 0.07$. We assume $\rho_g = 0.8$ and $\sigma_g = 0.01$ for the stochastic process of the government spending shock in (18) in both regimes.

Table B1. Parameter Calibrations

Preference and HHs	
β, discount factor	0.9958
h, habit formation	0.99
η, inverse Frisch labor elas.	2
$\bar{n}$, steady-state labor	1/3
δ, depreciation rate	0.025
α_g, subs. of private/public cons.	-0.2
Frictions and Production	
α, capital share	0.33
θ_p, elas. of subs. b/w intermediate goods	8
θ_w, elas. of subs. b/w different types of labor	8
ω_p, Calvo price stickiness	0.8
ω_w, Calvo wage stickiness	0.8
ζ_2, capital utilization	0.15
κ, investment adj. cost	5
Monetary/Fiscal Calibrations	
$\bar{\pi}$, steady-state gross inflation rate	1.0082
ψ_g, lagged resp. for govt spending	0.98
s_g, steady-state govt spending-to-GDP ratio	0.0945
s_b, steady-state debt-to-GDP ratio	1.3707
$\bar{\tau}^n$, steady-state labor tax rate	0.2171
$\bar{\tau}^k$, steady-state capital tax rate	0.2497
Regime D	
Monetary Policy	
ϕ_π, interest rate resp. to inflation	0.83
ϕ_y, interest rate resp. to output	0.27
ψ_r, resp. to lagged interest rate	0.68
Fiscal Policy	
γ_g, govt spending resp. to debt	0
Regime H	
Monetary Policy	
ϕ_π, interest rate resp. to inflation	2.15
ϕ_y, interest rate resp. to output	0.93
ψ_r, resp. to lagged interest rate	0.79
Fiscal Policy	
γ_g, govt spending resp. to debt	0.07
Shocks	
ρ_g, govt spending persistence	0.8
σ_g, govt spending	0.01

* Parameters are calibrated at a quarterly frequency.

C Equilibrium Conditions

- The first order conditions of the household

$$
\begin{aligned}
\lambda_t &= \frac{1}{c_t^* - hc_{a,t-1}^*} \\
c_t^* &= c_t + \alpha_g g_t \\
{R_t}^{-1} &= \beta E_t \left(\frac{\lambda_{t+1}}{\lambda_t} \frac{1}{\pi_{t+1}} \right) \\
a'(\mu_t) &= \left(1-\tau^k\right) r_t^k \\
q_t &= \beta E_t \left\{ \frac{\lambda_{t+1}}{\lambda_t} \left[\left(1-\tau^k\right) \mu_{t+1} r_{t+1}^k - a(\mu_{t+1}) + q_{t+1}(1-\delta) \right] \right\} \\
1 &= q_t \left[1 - S\left(\frac{i_t}{i_{t-1}}\right) - S'\left(\frac{i_t}{i_{t-1}}\right) \frac{i_t}{i_{t-1}} \right] + \beta E_t q_{t+1} \frac{\lambda_{t+1}}{\lambda_t} S'\left(\frac{i_{t+1}}{i_t}\right) \left(\frac{i_{t+1}}{i_t}\right)^2 \\
\Omega_t &= \lambda_t w_t n_t \left(\pi_{w,t}^*\right)^{1-\theta_w} + \omega_w \beta E_t \left(\frac{\pi_t^{\iota_w} \bar{\pi}^{1-\iota_w}}{\pi_{t+1}} \right)^{1-\theta_w} \left(\frac{\pi_{w,t+1}^*}{\pi_{w,t}^*} \frac{w_{t+1}}{w_t} \right)^{\theta_w - 1} \Omega_{t+1} \\
\Omega_t &= \mathcal{M}^w \varrho \left(\pi_{w,t}^*\right)^{-\theta_w(1+\eta)} n_t^{1+\eta} + \omega_w \beta E_t \left(\frac{\pi_t^{\iota_w} \bar{\pi}^{1-\iota_w}}{\pi_{t+1}} \right)^{-\theta_w(1+\eta)} \left(\frac{\pi_{w,t+1}^*}{\pi_{w,t}^*} \frac{w_{t+1}}{w_t} \right)^{\theta_w(1+\eta)} \Omega_{t+1}
\end{aligned}
$$

- The wage index evolves as:

$$
1 = \omega_w \left(\frac{w_{t-1}}{w_t} \frac{\pi_{t-1}^{\iota_w} \bar{\pi}^{1-\iota_w}}{\pi_t} \right)^{1-\theta_w} + (1-\omega_w) \left(\pi_{w,t}^*\right)^{1-\theta_w}
$$

- The first order conditions of the firm

$$
\begin{aligned}
0 &= (1-\theta_p) F_{2,t} + \theta_p F_{1,t} \\
F_{1,t} &= \lambda_t mc_t y_t + \omega_p \beta E_t \left(\frac{\pi_t^{\iota_p} \bar{\pi}^{1-\iota_p}}{\pi_{t+1}} \right)^{-\theta_p} F_{1,t+1} \\
F_{2,t} &= \lambda_t \pi_t^* y_t + \omega_p \beta E_t \left(\frac{\pi_t^{\iota_p} \bar{\pi}^{1-\iota_p}}{\pi_{t+1}} \right)^{1-\theta_p} \left(\frac{\pi_t^*}{\pi_{t+1}^*} \right) F_{2,t+1} \\
\frac{w_t}{r_t^k} &= \frac{1-\alpha}{\alpha} \frac{\mu_t k_{t-1}}{n_t} \\
mc_t &= \alpha^{-\alpha} (1-\alpha)^{\alpha-1} w_t^{1-\alpha} \left(r_t^k\right)^{\alpha}
\end{aligned}
$$

- The price level evolves:

$$
1 = \omega_p \left(\frac{\pi_{t-1}^{\iota_p} \bar{\pi}^{1-\iota_p}}{\pi_t} \right)^{1-\theta_p} + (1-\omega_p) \left(\pi_t^*\right)^{1-\theta_p}
$$

- Monetary authority follow its Taylor rule

$$
R_t = (R_{t-1})^{\psi_r} \left[\bar{R} \left(\frac{\pi_t}{\bar{\pi}}\right)^{\phi_\pi} \left(\frac{y_t}{\bar{y}}\right)^{\phi_y} \right]^{1-\psi_r}
$$

- Government budget constraint:

$$\frac{B_t}{P_t} + \tau^n w_t n_t + \tau^k r_t^k u_t k_{t-1} = \frac{R_{t-1} B_{t-1}}{P_t} + g_t + tr$$

where

$$g_t = g_{t-1}^{\psi_g} \left[\bar{g} \left(b_{t-1} / \bar{b} \right)^{-\gamma_g} \right]^{1-\psi_g} \nu_{g,t}$$

- Markets clear:

$$\begin{aligned} y_t &= \frac{\left(u_t k_{t-1} \right)^{\alpha} \left(n_t \right)^{1-\alpha}}{\xi_{p,t}} \\ y_t &= c_t + i_t + g_t + a\left(u_t \right) k_{t-1} \end{aligned}$$

where

$$\begin{aligned} a\left(u_t \right) &= \zeta_1 \left(u_t - 1 \right) + \frac{\zeta_2}{2} \left(u_t - 1 \right)^2 \\ \xi_{p,t} &= \omega_p \left(\frac{\pi_{t-1}^{\iota_p} \bar{\pi}^{1-\iota_p}}{\pi_t} \right)^{-\theta_p} \xi_{p,t-1} + \left(1 - \omega_p \right) \left(\pi_t^* \right)^{-\theta_p} \end{aligned}$$

and

$$\begin{aligned} k_t &= \left(1 - \delta \right) k_{t-1} + \left[1 - S\left(\frac{i_t}{i_{t-1}} \right) \right] i_t \\ S\left(\frac{i_t}{i_{t-1}} \right) &= \frac{\kappa}{2} \left(\frac{i_t}{i_{t-1}} - 1 \right)^2 \end{aligned}$$

- The government spending shock evolves according to

$$\ln \nu_{g,t} = \rho_g \ln \nu_{g,t-1} + \sigma_g \varepsilon_{g,t}$$

D Steady State

Given the steady state labor hours, the steady state inflation rate and the steady state fiscal policy calibration, the remaining variables are defined by the system:

$$
\begin{aligned}
R &= \frac{\bar{\pi}}{\beta} \\
r^k &= \frac{\frac{1}{\beta} - (1-\delta)}{1-\tau^k} \\
a'(1) &= r^k \left(1-\tau^k\right) \\
w &= (1-\alpha)\left[mc\left(\frac{\alpha}{r^k}\right)^{\alpha}\right]^{\frac{1}{1-\alpha}} \\
k &= \frac{\alpha}{1-\alpha}\frac{w\bar{n}}{r^k} \\
i &= \delta k \\
y &= k^{\alpha}\bar{n}^{1-\alpha} \\
c &= y - i - g \\
tr &= \left(1-\frac{1}{\beta}\right) b + \tau^n w n + \tau^k r^k k - g \\
c^* &= c + \alpha_g g \\
\lambda &= \frac{1}{c^*(1-h)} \\
mc &= \frac{\theta_p - 1}{\theta_p} \\
\varrho &= \frac{w\lambda}{\mathcal{M}^w n^{\eta}} \\
\chi &= \varrho\left(1-\tau_n\right)
\end{aligned}
$$

E Log-Linearized System

Let $\hat{x}_t = \ln(x_t/\bar{x})$ denote the percentage deviation of a variable x_t from its steady-state $\bar{x}$.

- The first order conditions of the household

$$
\begin{aligned}
\hat{\lambda}_t &= -\frac{1}{1-h}\hat{c}_t^* + \frac{h}{1-h}\hat{c}_{t-1}^* \\
\hat{c}_t^* &= \frac{c}{c+\alpha_g g}\hat{c}_t + \frac{\alpha_g g}{c+\alpha_g g}\hat{g}_t \\
\hat{\lambda}_t &= E_t\hat{\lambda}_{t+1} + \hat{R}_t - E_t\hat{\pi}_{t+1} \\
\hat{r}_t^k &= \frac{\zeta_2}{1-\zeta_2}\hat{u}_t \\
\hat{q}_t &= E_t\hat{\lambda}_{t+1} - \hat{\lambda}_t + \beta(1-\delta)E_t\hat{q}_{t+1} + \beta\left(1-\tau^k\right)r^k E_t\hat{r}_{t+1}^k \\
0 &= E_t\hat{i}_{t+1} - (1+\beta)\hat{i}_t + \frac{1}{\kappa}\hat{q}_t + \hat{i}_{t-1} \\
\hat{\Omega}_t &= \omega_w\beta\hat{\Omega}_{t+1} - \omega_w\beta(1-\theta_w)\hat{w}_{t+1} - \omega_w\beta(1-\theta_w)\hat{\pi}_{t+1} - \omega_w\beta(1-\theta_w)\hat{\pi}_{w,t+1}^* \\
&\quad + (1-\omega_w\beta\theta_w)\hat{w}_t + \omega_w\beta(1-\theta_w)\iota_w\hat{\pi}_t + (1-\theta_w)\hat{\pi}_{w,t}^* - (1-\omega_w\beta)\left(\hat{\lambda}_t + \hat{n}_t\right) \\
\hat{\Omega}_t &= \omega_w\beta\hat{\Omega}_{t+1} + \omega_w\beta\theta_w(1+\eta)\hat{w}_{t+1} + \omega_w\beta\theta_w(1+\eta)\hat{\pi}_{t+1} + \omega_w\beta\theta_w(1+\eta)\hat{\pi}_{w,t+1}^* \\
&\quad -\omega_w\beta\theta_w(1+\eta)\hat{w}_t - \omega_w\beta\theta_w(1+\eta)\iota_w\hat{\pi}_t - \theta_w(1+\eta)\hat{\pi}_{w,t}^* + (1-\omega_w\beta)(1+\eta)\hat{n}_t
\end{aligned}
$$

- The wage index evolves as:

$$
0 = \hat{\pi}_{w,t}^* - \frac{\omega_w}{1-\omega_w}\hat{w}_t - \frac{\omega_w}{1-\omega_w}\hat{\pi}_t + \frac{\omega_w}{1-\omega_w}\hat{w}_{t-1} - \frac{\omega_w\iota_w}{1-\omega_w}\hat{\pi}_{t-1}
$$

- The first order conditions of the firm

$$
\begin{aligned}
0 &= \hat{F}_{1,t} - \hat{F}_{2,t} \\
\hat{F}_{1,t} &= \omega_p\beta E_t\hat{F}_{1,t+1} + \omega_p\beta\theta_p E_t\hat{\pi}_{t+1} - \omega_p\beta\theta_p\iota_p\hat{\pi}_t + (1-\omega_p\beta)\left(\hat{\lambda}_t + \hat{mc}_t + \hat{y}_t\right) \\
\hat{F}_{2,t} &= \omega_p\beta E_t\left(\hat{F}_{2,t+1} - \hat{\pi}_{t+1}^*\right) - \omega_p\beta(1-\theta_p)E_t\hat{\pi}_{t+1} + \hat{\pi}_t^* + \omega_p\beta(1-\theta_p)\iota_p\hat{\pi}_t \\
&\quad + (1-\omega_p\beta)\left(\hat{\lambda}_t + \hat{y}_t\right) \\
\hat{w}_t &= \hat{r}_t^k + \hat{u}_t - \hat{n}_t + \hat{k}_{t-1} \\
\hat{mc}_t &= (1-\alpha)\hat{w}_t + \alpha\hat{r}_t^k
\end{aligned}
$$

- The price level evolves:

$$
0 = \hat{\pi}_t^* - \frac{\omega_p}{(1-\omega_p)}\hat{\pi}_t + \frac{\omega_p\iota_p}{(1-\omega_p)}\hat{\pi}_{t-1}
$$

- Monetary authority follow its Taylor rule

$$
\hat{r}_t = (1-\psi_r)\phi_\pi\hat{\pi}_t + (1-\psi_r)\phi_y\hat{y}_t + \psi_r\hat{r}_{t-1}
$$

- Government budget constraint:

$$\tau^n wn\left(\hat{w}_t+\hat{n}_t\right)+\tau^k r^k k\left(\hat{r}_t^k+\hat{u}_t\right)+b\hat{b}_t+\frac{b}{\beta}\hat{\pi}_t-g\hat{g}_t \quad = \quad \frac{b}{\beta}\left(\hat{r}_{t-1}+\hat{b}_{t-1}\right)-\tau^k r^k k\hat{k}_{t-1}$$

where

$$\hat{g}_t \quad = \quad \psi_g\hat{g}_{t-1}-\gamma_g\left(1-\psi_g\right)\hat{b}_{t-1}+\hat{\nu}_{g,t}$$

- Markets clear:

$$\begin{aligned} y\hat{y}_t &= c\hat{c}_t+i\hat{i}_t+g\hat{g}_t+\left(1-\tau^k\right)r^k k\hat{u}_t \\ \hat{y}_t &= \left(1-\alpha\right)\hat{n}_t+\alpha\hat{k}_{t-1}+\alpha\hat{u}_t-\hat{\xi}_{p,t} \end{aligned}$$

where

$$\begin{aligned} \hat{k}_t &= \left(1-\delta\right)\hat{k}_{t-1}+\delta\hat{i}_t \\ \hat{\xi}_{p,t} &= \omega_p\theta_p\hat{\pi}_t-\left(1-\omega_p\right)\theta_p\hat{\pi}_t^*-\omega_p\theta_p\iota_p\hat{\pi}_{t-1}+\omega_p\hat{\xi}_{p,t-1} \end{aligned}$$

- The government spending shock evolves according to

$$\hat{\nu}_{g,t} \quad = \quad \rho_g\hat{\nu}_{g,t-1}+\sigma_g\varepsilon_{g,t}$$

F Bayesian Estimation

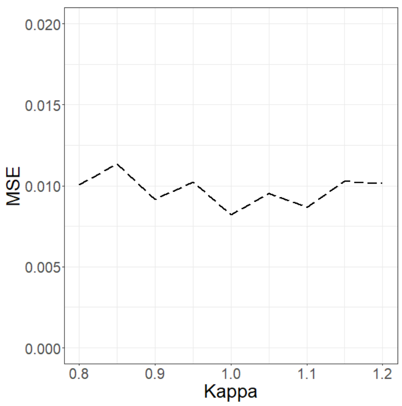
F.1 Prior Dominance

We find the full-sample estimates are highly sensitive to the prior values. The following perturbation experiment illustrates this point. Consider a fine grid as follows,

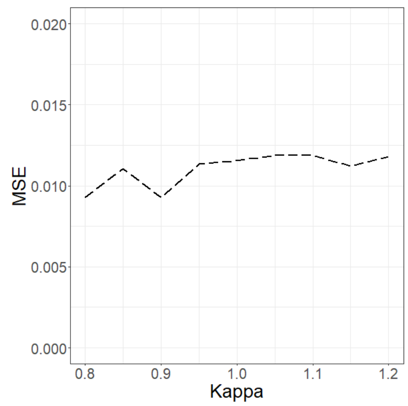
$$\kappa \in \{0.8, 0.85, 0.9, 0.95, 1, 1.05, 1.1, 1.15, 1.2\}$$

We scale the prior's mean by κ then estimate the model coefficients using the full sample.[6] Then we calculate the mean-squared difference between the coefficients' posterior mean with the prior's mean.

3/Prior.png



4/Prior.png



7/Prior.png

The left graph plots the average for figure 3. The middle graph plots the average for figure 4. And the right graph plots the average for figure 7. The flat graphs are suggesting the posterior estimates are closely following the changed prior values.

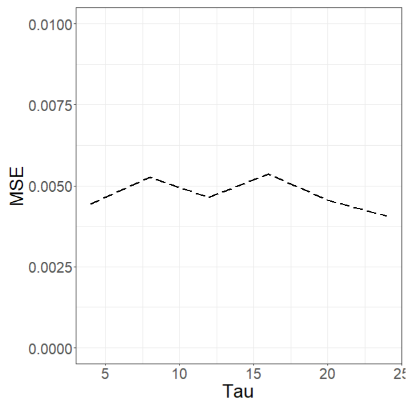
F.2 Diffusing Prior

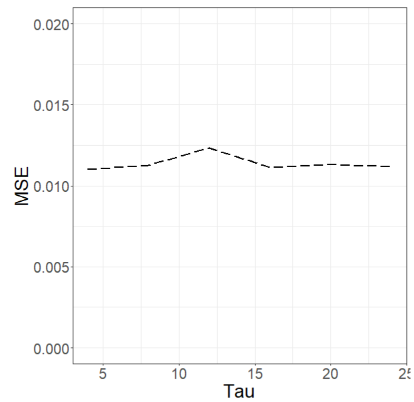
We also tried diffusing the prior distribution to weaken the prior's dominance. Consider the following grid,

$$\tau \in \{4, 8, 12, 16, 20, 24\}$$

τ is the multiplicative factor on the prior distribution's variance.[7] As τ increases, the prior distribution becomes more diffused. We calculate the mean-squared difference between the coefficients' posterior mean with the prior's mean over the different τ's.

[6]Obviously, $\kappa = 1$ is the benchmark case.
[7]$\tau = 4$ corresponds to the benchmark case.

3/Diffuse.png 

4/Diffuse.png 

7/Diffuse.pn

The left graph plots the average for figure 3. The middle graph plots the average for figure 4. And the right graph plots the average for figure 7. The flat graphs suggest the prior dominance persist even after diffusing the prior.

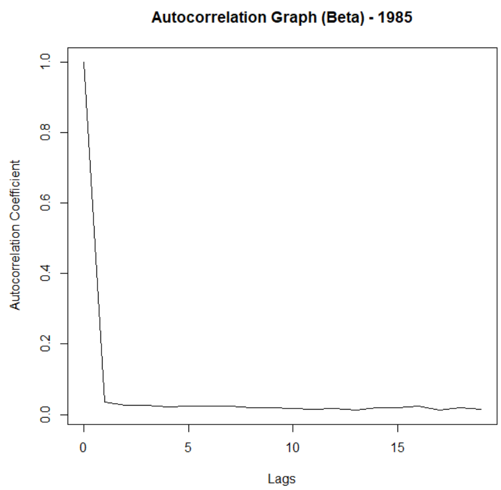
F.3 Comparing OLS Estimates: 1960-1969 vs 1989-1998

We compare the 1960-1969 OLS estimates against the 1989-1998 OLS estimates in fitting the data from 1999 to 2017. For each specification, we compute the average mean-squared errors from fitting to the 1999-2017 data basing first on the 1960-1969 OLS estimates and second on the 1989-1998 OLS estimates. The average is taken over time and over the four variables. Then we take a ratio of the first MSE over the second MSE. And a ratio being above one should suggest 1989-1998 OLS estimates provides a better fit.

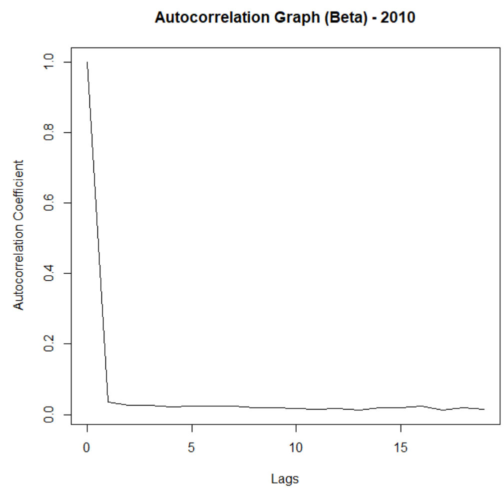
	Figure 3	Figure 4	Figure 7
Ratio	4.45795	1.322273	2.418615

The table shows the ratio is consistently above one for all specifications. So the 1989-1998 OLS estimates provides a better fit for observations from 1999 to 2017. And it is why we choose sample splitting to use 1989-1998 OLS estimates as priors for posterior estimates in 1999 to 2017.

Convergence - Benchmark



1.png



2.png

The left graph documents the average autocorrelation of autoregressive coefficients' posterior draws for the year 1985. And the right graph does the same for the year 2010. Both graphs show sharp decay in autocorrelation between sequential draws from the Gibbs Sampler. These posterior draws are kept after discarding the first 3000 draws treated as burn-ins. It seems using 3000 burn-ins is sufficient to simulate approximately independent posterior draws.

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