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How Has the Effects of Government Spending Evolved in the Post-War United States?

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Abstract

This paper examines the evolution of the output and price effects of government spending in the post-war United States. Exploiting the flexibility of a time-varying parameter vector autoregressive (TVP-VAR) framework and sign restriction identification, both cyclical and structural variations in multipliers are simultaneously observed: multipliers are large in recessions and have declined since the 1980s. The results also suggest that the inflationary effects of government spending became stronger along with an accumulation of public debt, but that regime change is unlikely to have occurred. The Ricardian channel is still considered the main cause of the decline in multipliers.

JEL classification: E30, E60, H60.

Keywords: Bayesian VARs; Time-varying parameters; Fiscal multipliers; Fiscal policy; Monetary policy; Fiscal theory of the price level.

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1 Introduction

The past decade has seen increased attention paid to the role of fiscal policy after the Global Financial Crisis of 2007-2008. As monetary policy became less effective, the World Bank, the IMF, and the OECD recommend using fiscal policy as a stabilization tool for countries with ample fiscal space (e.g., World Bank (2015); IMF (2016); OECD (2016)).[1] However, fiscal space is narrowing in many advanced countries, including the United States, where the debt-to-output ratio has been increasing largely as a result of the aging population. The rapid accumulation of U.S. public debt and the U.S. Federal Reserve's (Fed) adoption of unconventional monetary policy after the Crisis also revived the debate over the fiscal theory of the price level (FTPL).[2] This study aims to provide new evidence on the changes in the effects of government spending on output and prices in the post-war United States within the framework of a time-varying parameter vector autoregressive (TVP-VAR) model.

The relationship between public debt and the effects of fiscal policy on output has been extensively studied since Giavazzi and Pagano (1990) suggested the presence of non-Keynesian effects on the basis of Danish and Irish experiences in the 1980s. In searching for the underlying cause of non-Keynesian effects, Perotti (1999) use panel data from OECD countries to show that the impact of government spending shocks can differ depending on the initial level of debt. Furthermore, recent studies investigate sources of heterogeneity in multipliers across countries (e.g., Favero, Giavazzi and Perego (2011); Corsetti, Meier and Müller (2012*b*); Ilzetzki, Mendoza and Végh (2013); Nickel

[1] "Fiscal space" is defined as the difference between a current debt-to-output ratio and its upper limit. The upper limit is suggested to be measured in terms of either market access or fiscal sustainability, called "debt limit" or "fiscal limit" respectively (e.g., OECD (2016)). The former thinks of the limit as a point beyond which a government is unable to roll over its debt due to doubts raised about solvency based on its past fiscal adjustment record (e.g., Ghosh et al. (2013)). The latter considers it a probability that a tax rate may reach the peak of the Laffer curve meaning that a government will be unable to repay its debt in the future (e.g., Cochrane (2011); Davig, Leeper and Walker (2011); Bi (2012)). These two concepts of the limit are closely related as the probability of default affects market access through sovereign risk premia. In either case, fiscal space tends to narrow as the actual debt-to-output ratio increases.

[2] See, for example, Canzoneri, Cumby and Diba (2011) for a comprehensive overview of the FTPL literature.

and Tudyka (2014)) and provide evidence that the size of the government spending multiplier is large in economies with a low debt-to-output ratio. Several studies suggest households' Ricardian behavior as the cause of the debt-dependent output effects of fiscal policy (e.g., Sutherland (1997); Perotti (1999); Bi, Shen and Yang (2016)).

Although time variation in the multiplier in the United States is an area of active research, existing studies focus on its state-dependent nature across business cycles. A growing body of evidence suggests that government spending multipliers are larger in recessions than in expansions (e.g., Auerbach and Gorodnichenko (2012); Bachmann and Sims (2012); Candelon and Lieb (2013); Caggiano et al. (2015)).[3] The basic idea behind the state dependency of multipliers is that crowding out is less likely to occur in the presence of economic slack. Studies that consider data from a panel of countries also report similar results (e.g., Auerbach and Gorodnichenko (2013); Riera-Crichton, Vegh and Vuletin (2015)). Whereas Auerbach and Gorodnichenko (2013) consider debt-dependent effects in studying the state-dependent nature of multipliers on the basis of cross-country panel data, the role of public debt in the effects of government spending in U.S. time series data has not yet been examined. In addition, empirical studies of multipliers over business cycles typically conduct analysis using only real variables, despite the fact that monetary policy response is the crucial determinant of the crowding out effects of government spending.[4]

The effects of government spending shocks may also differ across policy regimes. Several studies suggest that multipliers decreased because of changes in the conduct of monetary policy after Paul Volcker's appointment as the Fed Chairman in 1979 (e.g., Perotti (2004); Bilbiie, Meier and Müller (2008)). However, disagreement exists over whether the effect of monetary policy on the economy changed drastically.[5] On the other hand, a recent strand of the FTPL literature provides a new look

[3] In contrast, Ramey and Zubairy (2017) do not find evidence that multipliers differ across the state of the U.S. economy using Ramey (2011)'s extended military news series and Jordà (2005)'s local projection method.

[4] Recent theoretical studies highlight the importance of monetary policy in determining the size of the multiplier during a liquidity trap (e.g., Woodford (2011); Christiano, Eichenbaum and Rebelo (2011); Erceg and Lindé (2014)).

[5] Several studies find little evidence in favor of the view that monetary policy played a main role in the Great

at the importance of monetary and fiscal policy regimes in determining the effects of government spending (e.g., Davig and Leeper (2011*a*); Leeper, Traum and Walker (2017)). Whereas studies in this strand report rather mixed results regarding the timing of monetary and fiscal policy regime changes, these studies generally support the presence of Ricardian fiscal policy regime during the post-Volcker period (e.g., Davig and Leeper (2011*a*); Traum and Yang (2011); Bhattarai, Lee and Park (2016); Bianchi and Ilut (2017)).[6] In contrast, studies based on linear models tend to find a Ricardian fiscal regime throughout the post-war period (e.g., Bohn (1998); Canzoneri, Cumby and Diba (2001); Canzoneri, Cumby and Diba (2011)).[7] Although the U.S. economy has supposedly not yet reached the fiscal limit (e.g., Davig, Leeper and Walker (2010);Cochrane (2011)), De Graeve and Queijo von Heideken (2015) suggest that the concerns over fiscal inflation have already started increasing their inflationary pressures since 2006.

Against this background, this study examines the evolution of the output and price effects of government spending in the post-war United States paying attention to the aforementioned various strands of the literature. For this purpose, we employ a TVP-VAR model with stochastic volatility along the lines of Primiceri (2005), who considers time-dependent contemporaneous relations among variables. The TVP-VAR model allows the parameters to vary continuously over time in a stochastic manner and, hence, is suitable for capturing permanent changes in the transmission mechanism.[8] Therefore, the model may well describe possible changes in household behavior and in the conduct of monetary and fiscal policy. Although rapid changes in the economic state are difficult to capture within the model *per se*, we consider them with the assistance of sign restrictions in the spirit of

Moderation starting in the mid-1980s (e.g., Cogley and Sargent (2005); Primiceri (2005); Sims and Zha (2006); Gambetti, Pappa and Canova (2008)).

[6] This strand of literature examines whether linearized monetary and fiscal policy rules are "active" or "passive" in the sense of Leeper (1991) within a framework of regime-switching DSGE models. When fiscal policy regime is found to be passive, the government is supposed to follow Ricardian fiscal regime but not necessarily vice versa. As to the difference in the concept of fiscal policy regimes, see Canzoneri, Cumby and Diba (2011), for example.

[7] Woodford (2001) suggests that the Fed's bond-price support in the 1940s is a good illustration of a non-Ricardian fiscal regime.

[8] Primiceri (2005) provides a succinct discussion of the advantages and disadvantages of TVP-VAR models over regime switching models.

Canova and Pappa (2011). The state-dependent effects of government spending can be captured by imposing additional sign restrictions in accordance with the economic state of each period. We choose an actual monetary policy phase as a proxy measure of economic slack in considering the economic state. To implement the sign restriction approach on a period-by-period basis, we exploit the algorithm proposed by Rubio-Ramírez, Waggoner and Zha (2010), which allows us to identify several shocks in a highly parameterized TVP-VAR model with great efficiency.

A large strand of the literature documents the time-varying effects of monetary policy within the TVP-VAR framework (e.g., Cogley and Sargent (2005); Primiceri (2005); Benati (2008); Canova and Gambetti (2009)), but only a few studies employ the methodology to investigate possible changes in the effects of fiscal policy. The notable exceptions are Pereira and Lopes (2014), Kirchner, Cimadomo and Hauptmeier (2010), and Rafiq (2012), who report a decline in the effectiveness of fiscal policy in the United States, the Euro area, and Japan, respectively.[9] Our study differs from them in that we examine time variation in both the output effects and the price effects of government spending and consider the economic state of each period using the assistance of sign restriction identification.

In the following, using findings from previous studies, we consider monetary policy and public debt as promising candidates for the driving forces behind the changes in the output and price effects of government spending. Therefore, we work with a medium-scale TVP-VAR model that incorporates monetary variables and public debt. We first identify shocks through basic agnostic sign restrictions, while relying on a traditional recursive identification scheme for the purpose of the robustness check. Regardless of the identification scheme used, we find changes in the transmission of fiscal policy and observe little time variation in changes in monetary policy transmission. The

[9] Kirchner, Cimadomo and Hauptmeier (2010) is the only study that performs exercises to investigate the driving forces behind the changes. By conducting regression analysis using the estimated government spending multipliers calculated from their estimated TVP-VAR model and possible explanatory factors, they conclude that increasing public debt is the main cause of the observed decline in multipliers in the Euro area.

unconditional estimates of the impulse responses indicate that government spending multipliers have declined substantially since the 1980s. In contrast, the price effects of government spending show an upward trend during the same period. Then, we calculate the impulse responses to government spending shocks by imposing additional identification restrictions that reflect the actual monetary policy phase of each period. The method allows us to simultaneously find both cyclical and structural variations in multipliers within a single TVP-VAR framework. We obtain larger multipliers in recessions than in expansions and smaller ones in the post-Volcker period than in the preceding period. With respect to price responses, we observe relatively strong inflationary effects during the boom periods of the 2000s and the 2010s. Because we find strong correlation between the output and price effects of government spending and debt-to-output ratio, we examine possible changes in the prevalence of either Ricardian or non-Ricardian fiscal regimes by applying the bivariate VAR methodology in Canzoneri, Cumby and Diba (2001) and Canzoneri, Cumby and Diba (2011) to our TVP-VAR framework. The analysis confirms that the degree of Ricardian behavior of the U.S. government was strengthened during most of the post-Volcker period. We further find unidirectional Granger causality from the government's stronger corrective action to the decline in the multipliers. Since regime change is unlikely to have occurred, the observed debt-dependent price effects can be attributed to the heightened inflation expectations given the narrowing of the fiscal space. Our results indicate that the public debt accumulation that began in the 1980s is the underlying cause behind the decline in multipliers and the increase in inflationary effects of government spending during the period.

The remainder of this paper is organized as follows. Section 2 discusses the empirical methodology. Section 3 reports the changes in fiscal and monetary policy transmission. Section 4 investigates the mechanism underlying the time-varying effects of government spending with a particular focus on the role of monetary policy and public debt. Section 5 concludes.

2 Empirical Methodology

2.1 A VAR model with Time-Varying Parameters and Stochastic Volatility

We consider the following VAR (p) model with time-varying parameters and stochastic volatility:

$$y_t = B_{1,t}y_{t-1} + \cdots + B_{p,t}y_{t-p} + u_t, \tag{1}$$

for $t = p+1, ..., T$, where y_t is a $k \times 1$ vector of observed variables and $B_{i,t}$, $i = 1, ...p$, are $k \times k$ matrices of time-varying coefficients. The u_t is a $k \times 1$ vector of heteroskedastic shocks that are assumed to be normally distributed with a zero mean and a time-varying covariance matrix, Ω_t. Following established practice, we decompose u_t as $u_t = A_t^{-1}\ \Sigma_t \varepsilon_t$, where

$$A_t = \begin{bmatrix} 1 & 0 & \cdots & 0 \\ a_{21,t} & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ a_{k1,t} & \cdots & a_{kk-1,t} & 1 \end{bmatrix}, \tag{2}$$

$$\Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \cdots & 0 \\ 0 & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma_{k,t} \end{bmatrix}, \tag{3}$$

and $\varepsilon_t \sim N(0, I_k)$. It follows that $A_t \Omega_t A_t' = \Sigma_t \Sigma_t'$. Let β_t be a stacked $k^2 p \times 1$ vector of the elements in the rows of the $B_{1,t}, ...B_{p,t}$, and a_t be the vector of non-zero and non-one elements of the A_t. We assume that these vectors follow a random walk process:

$$\beta_{t+1} = \beta_t + u_{\beta,t}, \tag{4}$$

$$a_{t+1} = a_t + u_{a,t}, \tag{5}$$

$$h_{t+1} = h_t + u_{h,t}, \tag{6}$$

$$\begin{bmatrix} \varepsilon_t \\ u_{\beta,t} \\ u_{a,t} \\ u_{h,t} \end{bmatrix} \sim N\left(0, \begin{bmatrix} I & O & O & O \\ O & \Sigma_\beta & O & O \\ O & O & \Sigma_a & O \\ O & O & O & \Sigma_h \end{bmatrix}\right), \tag{7}$$

where $h_t = [h_{1,t}, \ldots, h_{k,t}]'$ with $h_{j,t} = \ln \sigma_{j,t}^2$ for $j = 1, ..., k$, and I is a k-dimensional identity matrix. The prior distributions for the initial values are given by $\beta_{p+1} \sim N(\mu_{\beta_0}, \Sigma_{\beta_0})$, $a_{p+1} \sim N(\mu_{a_0}, \Sigma_{a_0})$, and $h_{p+1} \sim N(\mu_{h_0}, \Sigma_{h_0})$. Observe that the model allows both the parameters that govern contemporaneous relations among variables and the log of the variance for the shocks to evolve over time as a random walk.

The stochastic volatility assumption makes the likelihood function of the model difficult to construct and Bayesian inference via Markov Chain Monte Carlo (MCMC) methods are required. To estimate a model that contains a relatively large number of variables, we rely on the efficient algorithm proposed by Nakajima, Kasuya and Watanabe (2011). Following Nakajima (2011*a*), we further assume for simplicity that Σ_β, Σ_a, Σ_h, Σ_{β_0}, Σ_{a_0}, and Σ_{h_0} are all diagonal matrices.[10] Regarding the sampling of β_t and a_t, we use the simulation smoother of de Jong and Shephard (1995) because the model can be written as a linear Gaussian state space form conditional on

[10] Although the assumption is not essential, it greatly simplifies the inference procedures for a_t and h_t, thereby contributing to increase the efficiency of the algorithm (e.g., Primiceri (2005); Nakajima, Kasuya and Watanabe (2011)). Moreover, we do not expect a significant difference in results from allowing for correlations among elements of a_t, β_t, and h_t, as in Primiceri (2005), Nakajima (2011*a*), and Nakajima (2011*b*), respectively.

the rest of the parameters.[11] In contrast, in sampling h_t, we employ the multi-move sampler of Shephard and Pitt (1997) and Watanabe and Omori (2004) for non-linear and non-Gaussian state space models. The multi-move sampler is more efficient than the single-move sampler of Jacquier, Polson and Rossi (1994).[12] Furthermore, it enables us to draw a sample from the exact conditional posterior density of the stochastic volatility, unlike the mixture sampler of Kim, Shephard and Chib (1998). Appendix B provides a more detailed outline of the MCMC algorithm used in this study.

2.2 Data and Identification Strategies

We use U.S. quarterly data for the period from 1952:Q1 to 2013:Q4.[13] The sample covers the post-Global Financial Crisis period, during which the government rapidly accumulated its debt and the Fed continued to expand its quantitative easing. The observed variables include government spending, gross domestic product (GDP), private consumption, debt-to-output ratio, GDP deflator, and nominal interest rate. See Appendix A for a detailed description of the data sources. Because the level of public debt and the conduct of monetary policy affect the output and price effects of government spending, we include the debt-to-output ratio and monetary variables, such as price level and interest rate, in the TVP-VAR model. This setup also allows us to consider the role of the monetary policy phase, which is neglected by most previous VAR studies on the effects of fiscal policy. Note that we include the debt-to-output ratio without imposing any restrictions, as in

[11] We employ the simulation smoother of de Jong and Shephard (1995) instead of the multi-state sampler of Carter and Kohn (1994), which is widely used in previous TVP-VAR studies. The multi-state sampler generates the entire state vector at once and therefore converges more quickly than the single-state sampler that yields a strong correlation among the samples. However, the method is prone to the problem of degeneracies because the entire state vector is constructed recursively. The simulation smoother of de Jong and Shephard (1995) avoids the problem by drawing disturbances rather than states.

[12] The shortcoming of using the single-move sampler is that it leads to slow convergence when state variables are highly autocorrelated. The multi-move sampler reduces the inefficiency by generating randomly selected blocks of disturbances rather than each state variable at a time.

[13] The sample period starts at the time when quarterly data series on public debt is available. The sample excludes the period of monetary policy normalization, which began with the tapering of quantitative easing in January 2014. Although our sample period covers the zero interest-rate policy period, we do not explicitly consider the zero lower bound in light of the findings in Nakajima (2011a). Using Japanese data from 1977 to 2010, Nakajima (2011a) provides evidence that a TVP-VAR model with stochastic volatility can produce almost the same result as one that incorporates the zero lower bound.

Corsetti, Meier and Müller (2012a), because we do not rule out the possibility that the price level adjusts to satisfy the government's intertemporal budget constraint.[14] The first three variables are expressed in real per capita terms. We use the logarithm for all variables except the debt-to-output ratio and the nominal interest rate. All variables are detrended with a linear and quadratic trend, and are seasonally adjusted, except for the interest rate. The lag length is set to $p = 4$, following Blanchard and Perotti (2002).

To implement the sign restriction approach within the TVP-VAR framework, we exploit the algorithm proposed by Rubio-Ramírez, Waggoner and Zha (2010) (RWZ algorithm, hereafter), as in Benati (2008). The RWZ algorithm proceeds as follows. We draw an independent standard normal $k \times k$ matrix Z_s for period s. The QR decomposition of Z_s gives an orthogonal matrix Q_s that satisfies $Q_s Q_s' = I$ and an upper triangular matrix R_s. Using $A_s^{-1}\Sigma_s Q_s$, we generate impulse responses for each MCMC replication. If the impulse response satisfies the restrictions, we keep the draw; otherwise, we discard it. The combination of Q_s' and ε_s, $\varepsilon_s^* = Q_s'\varepsilon_s$ is now regarded as a new set of structural shocks with the same covariance matrix as the original shock ε_s. Because Q_s is orthogonal, the new shocks are orthogonal to each other by design. The RWZ algorithm allows us to impose orthogonality conditions only by identifying other uncorrelated shocks. The algorithm is particularly appealing for identifying several shocks within our highly parameterized medium-scale TVP-VAR model because it is computationally efficient as Fry and Pagan (2011) address.

The TVP-VAR framework allows parameters to vary continuously over time in a stochastic manner; hence, it is not suitable for capturing rapid changes in the economic state. Nevertheless,

[14] Chung and Leeper (2007) and Favero and Giavazzi (2012) impose equations that represent the government's intertemporal budget constraint on their VAR models to capture feedback from debt to fiscal instruments. Note that their specification includes tax revenue whereas ours does not, as in Corsetti, Meier and Müller (2012a). We include in our system the price level rather than the inflation rate because we are interested in the possibility of price level adjustments, although the estimation results do not differ much if we use the inflation rate instead of the price level. Using the GDP deflator in estimating VARs can be found in Uhlig (2005), Sims and Zha (2006), and Mountford and Uhlig (2009).

we can consider the effects of such changes using sign restrictions and by exploiting the RWZ algorithm that we implement in each period. As Canova and Pappa (2011) suggest, the sign restriction approach enables us to study the effectiveness of fiscal policy under a certain economic state by imposing additional sign restrictions. Together with the assumption that the parameters governing contemporaneous relations among variables are time variant, we can impose different sets of sign restrictions on a period-by-period basis, considering the economic state of each period. Thus, it is possible to replicate the impact of government spending shocks that reflect the effects of rapid changes in the economic state.

3 Evolution of the Transmission Mechanism

3.1 Basic Results

Figure 1 presents the point estimates (posterior means) for the stochastic volatility of the reduced-form innovations, ε_t. The time variation in the volatility estimates of reduced-form shocks for interest rate and prices are largely consistent with those reported in previous studies (e.g., Cogley and Sargent (2005); Primiceri (2005); Koop, Leon-Gonzalez and Strachan (2009); Mumtaz and Zanetti (2013)). The volatility of reduced-form interest rate shock increased substantially around Paul Volcker's appointment as Fed Chairman in 1979 and showed a large decline during the early 1980s. The volatility of price shock reached its highest peak during the Great Inflation of the mid-1970s. The smoother variation in the volatility of the price level compared with that of the inflation rate reported in previous studies can be attributed to the difference in the variables. The volatility of the output shock declined sharply in the early 1980s, which shows a similar pattern to that of unemployment reported in Cogley and Sargent (2005). Justiniano and Primiceri (2008) also show a reduction in the volatility of government spending. It is worth noting that the volatility of interest rate declined significantly during the zero interest-rate policy period.[15] The results are in line with the findings in Nakajima (2011*a*) that suggest that the zero lower bound of nominal interest rates has negligible effects on impulse responses in a TVP-VAR model with stochastic volatility.[16] Since the estimation results here are largely consistent with those reported in previous studies, we can conclude that the time-varying volatilities are well captured in our model. The inclusion of stochastic volatility in the TVP-VAR model appears to be essential to appropriately detecting structural changes in the transmission of government spending shocks.

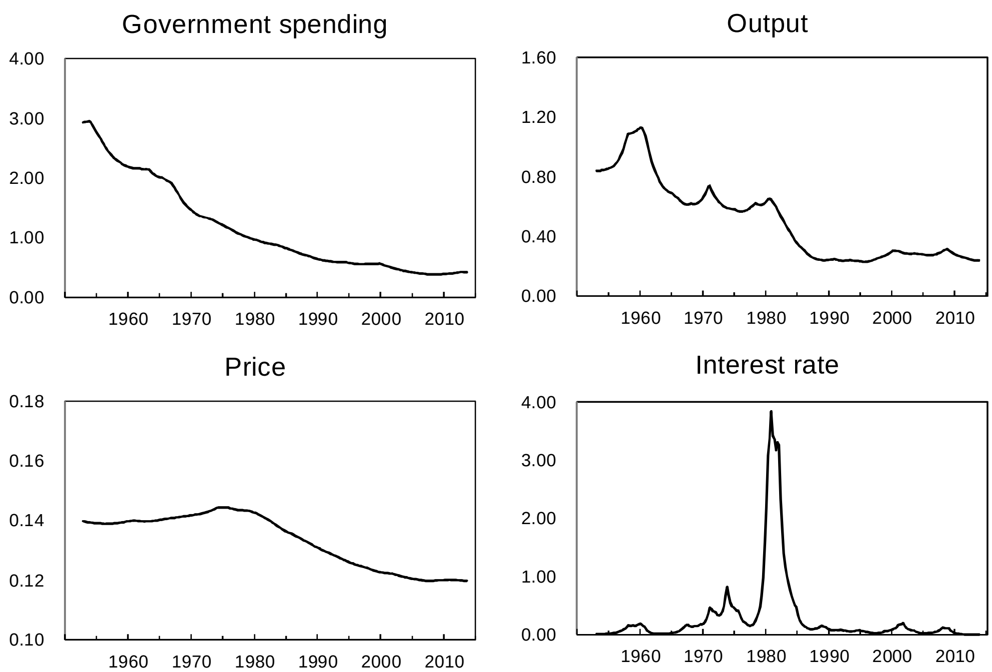

FIG. 1. Posterior mean volatilities of selected reduced-form innovations.

We present the sign restrictions that we employed in calculating impulse responses in Table 1.[17] We impose a minimum set of contemporaneous restrictions to make our identification as

[15] The decline in volatility is significant, particularly after the Fed's second and third rounds of quantitative easing (QE2 and QE3, respectively), which were announced in November 2010 and September 2012, respectively.

[16] Nakajima (2011*a*) finds an effectively low level of stochastic volatility for the monetary policy shocks during the zero interest-rate period in Japan based on an original TVP-VAR model that has no constraints on nominal interest rates. Including stochastic volatility is suggested as a source of similarity in the impulse responses between the original TVP-VAR model and the extended TVP-VAR model that incorporates the zero lower bound.

[17] To compare the results with those of other studies, we restrict our focus in this study to a traditional unanticipated government spending shock.

agnostic as possible.[18] In particular, we do not impose restrictions on output responses to fiscal and monetary policy shocks, as in Uhlig (2005) and Mountford and Uhlig (2009). As Canova and Pappa (2011) argue, existing theories do not provide definitive answers to the short-run dynamics after a government spending shock. Furthermore, it is computationally burdensome to estimate impulse responses from a TVP-VAR model that imposes sign restrictions for several periods.

That an increase in government spending has a positive impact on the debt-to-output ratio is the key identifying restriction that distinguishes government spending shocks from other shocks.[19] We also require government spending shocks to be orthogonal to monetary policy and business cycle shocks, following Mountford and Uhlig (2009). We borrow their definition for the restrictions to identify "monetary policy" and "business cycle" shocks. Notice that the effects of government spending studied here do not consider the state of the economy and hence can be regarded as unconditional effects. Although the automatic response of monetary policy to a government spending shock in subsequent periods is considered in the model's transmission mechanism, the actual monetary policy phase of the period is not taken into account.

TABLE 1

Sign restrictions

	Government spending	Monetary policy	Business cycle
Government spending	$+$		
Output			$+$
Consumption			
Price		$-$	
Interest rate		$+$	
Debt-to-output ratio	$+$		$-$

Notes: The table shows the signs imposed on the impulse responses of the variables to an expansionary government spending shock, a contractionary monetary policy shock, and a positive business cycle shock. A blank indicates that the variable's response is unrestricted. A positive [negative] sign indicates that the variable's response is restricted to being positive [negative] on impact.

[18] The choice of the period during which to restrict the responses does not change the basic results. Similar time variation in government spending multipliers appear when we impose sign restrictions for a year after a shock.

[19] The restriction shares similarities with those in previous studies (e.g., Pappa (2009); Enders, Müller and Scholl (2011); Bouakez, Chihi and Normandin (2014))

The agnostic sign restriction identifications, however, have a drawback that a wide range of impulse responses tends to be chosen (e.g., Uhlig (2005); Fry and Pagan (2011)). Thus, we compare the results with those obtained using the traditional recursive identification as a robustness check. We follow Corsetti and Müller (2006) for the Cholesky decomposition to order the real variables before the monetary variables. As in Blanchard and Perotti (2002), government spending is ordered first. Regarding monetary policy, the interest rate is ordered last, following Uhlig (2005) and Primiceri (2005).

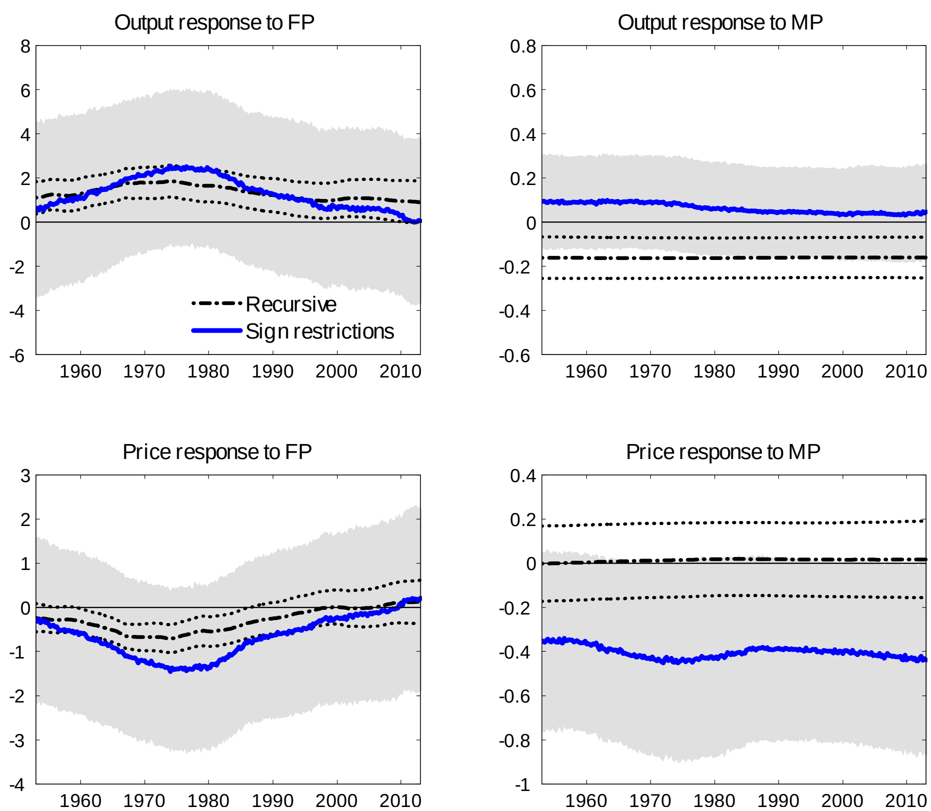

FIG. 2. The output and price responses to fiscal and monetary policy shocks after 12 quarters. *Notes:* The upper [lower] left panel displays responses to a one-dollar [one-percent] increase in government spending. The right panels display responses to a one percentage point increase in the interest rate. The solid [dash-dotted] lines represent posterior mean responses with the shaded areas [areas between dotted lines] representing the 16th-84th percentile ranges for sign restriction [recursive] identification.

Figure 2 presents the output and price responses to fiscal and monetary policy shocks at a three-year horizon. The impulse response at time t is computed for each MCMC replication on the basis of the estimated time-varying parameters at time t.[20] We divide the output responses to government spending shocks by the sample average ratio of output and government spending as in Auerbach and Gorodnichenko (2012), enabling the responses to be interpreted as output multipliers.[21] Although the magnitude of the changes differs depending on the identification scheme employed, the time variation in the output and price responses show similar patterns: the output effects of government spending shocks were on a downward trend since the 1980s, whereas the price effects were on an upward trend during the same period.

In line with the findings in Primiceri (2005), output and price responses to contractionary monetary policy shocks show, in contrast, little time variation. As Uhlig (2005) addresses, monetary policy shocks identified through sign restrictions have no clear effect on output and are followed by a decline in prices, whereas contractionary effects on output and the "price puzzle" first demonstrated by Sims (1992) are observed for recursive identification.

[20] Koop, Pesaran and Potter (1996) propose a method to calculate impulse responses considering the history of observations that affect impulse responses in non-linear models. However, because we expect a slight difference from using the method as in Koop, Leon-Gonzalez and Strachan (2009), which can be computationally demanding, we follow the simple computational procedure used in Primiceri (2005), Koop, Leon-Gonzalez and Strachan (2009), and Nakajima, Kasuya and Watanabe (2011).

[21] Ramey and Zubairy (2017) point out a potential problem arising from the use of the sample average ratio to calculate multipliers by considering the large variation found in their long samples of historical data. Nevertheless, we use the average ratio not only because it is relatively stable in our post-war sample, but also because we intend to highlight the time variation in the multipliers caused by changes in the transmission and economic state without interference from changes in the ratio.

3.2 Time-Variation in the Unconditional Impulse Responses

We now observe the changes in the shapes of the impulse responses to government spending shocks. Because the point estimate of the maximum impact on output (peak multipliers) takes the largest and the smallest values in 1974:Q1 and 2012:Q2, respectively, for the sign restriction identification, Figure 3 shows the impulse responses of output and consumption to government spending shocks (i.e., output and consumption multipliers) in these periods. The output responses for the recursive identification are also presented for comparative purposes. Although the sizes of multipliers are different across the identification schemes, their shapes and time variations are similar. After almost one year of decline, the output multipliers increase and reach their highest peak around a three-year horizon. It is worth noting that the output responses for recursive identification show quite similar shapes to that estimated by Blanchard and Perotti (2002). Their output response, which is estimated using U.S. data for 1947:Q1 to 1997:Q4, lies largely between our estimates in 1974:Q1 and 2012:Q2 for recursive identification. Figure 3 also illustrates the similarity between the time variation patterns in output and consumption multipliers. Their similar patterns suggest that the time variation in the effects of government spending on output is mostly led by that on consumption.

Figure 4 presents the price responses to government spending shocks. As in the case of multipliers, the mean responses show similar shapes and time variation across identification schemes. The negative price responses to government spending shocks in 1974:Q1 seem puzzling, but these counterintuitive results are prevalent across existing studies, as Perotti (2004) and Mountford and Uhlig (2009) point out (e.g., Edelberg, Eichenbaum and Fisher (1999); Fatás and Mihov (2001); Canova and Pappa (2011)). Nevertheless, the results suggest that government spending may have had a positive impact on the price level in 2012:Q2.

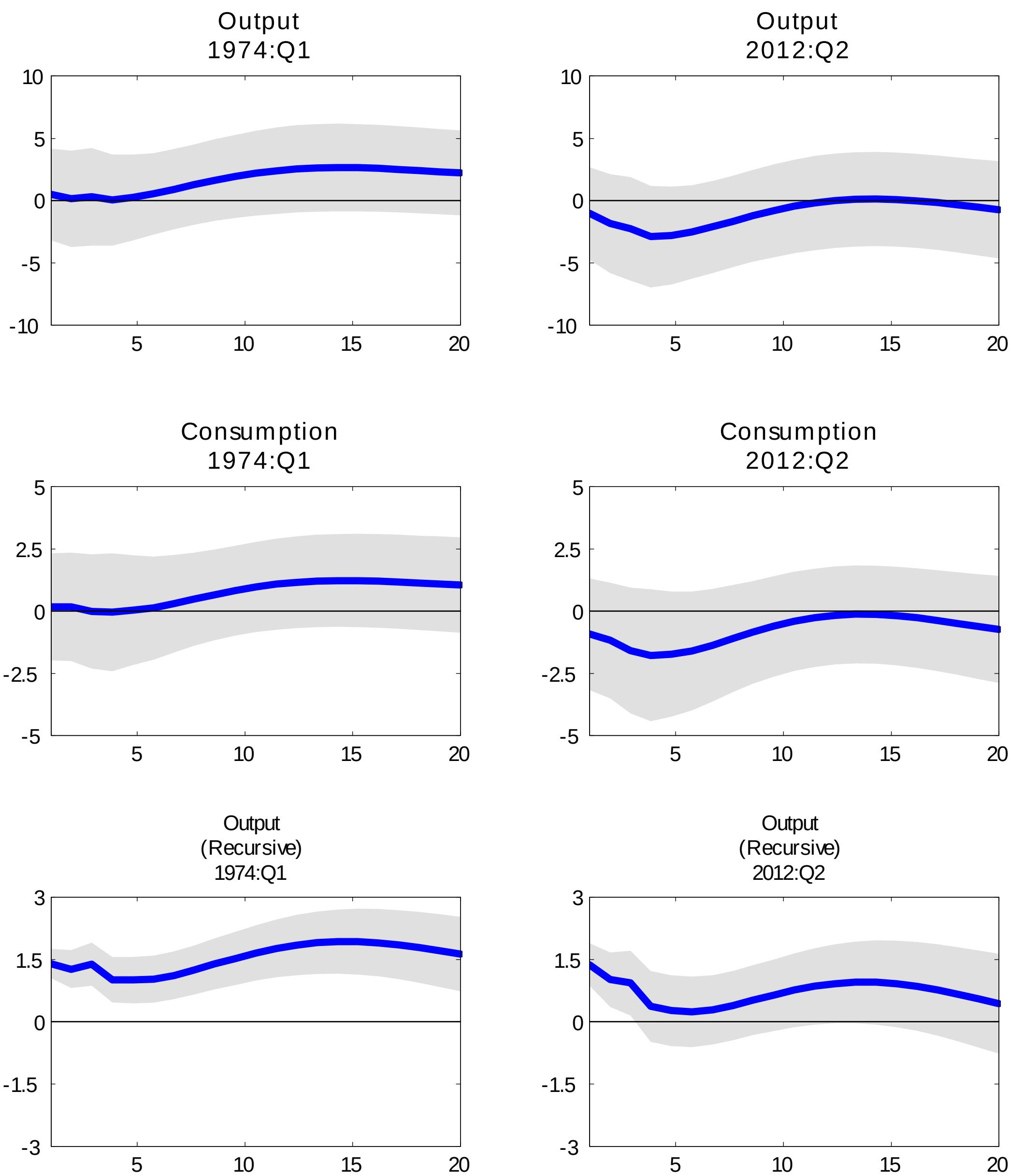

FIG. 3. The unconditional output and consumption multipliers. *Notes:* The solid lines represent posterior mean responses to a one-dollar increase in government spending with the shaded areas representing the 16th-84th percentile ranges. The top and middle [bottom] panels show responses obtained using sign restriction [recursive] identification.

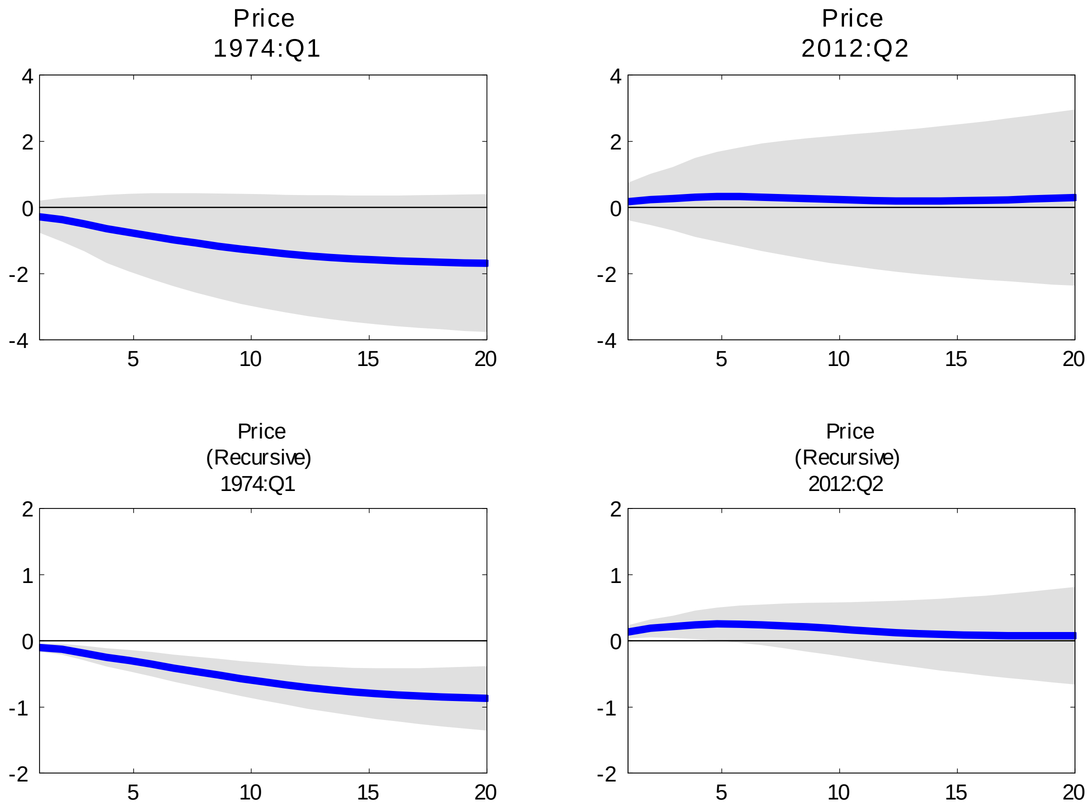

FIG. 4. The unconditional impulse responses of prices. *Notes:* The solid lines represent posterior mean responses to a one-percent increase in government spending with the shaded areas representing the 16th-84th percentile ranges. The top [bottom] panels show responses obtained using sign restriction [recursive] identification.

Overall, the shapes of impulse responses to government spending shocks are consistent with those reported in existing studies. While the difference in the magnitude of impulse response across identification schemes are significant, their time variation is found to be quite similar. To examine the state-dependent effects of government spending using sign restrictions in the following sections, we therefore focus our attention to the time variation of the effects rather than their magnitude.

4 Role of Monetary Policy and Debt Accumulation

4.1 Impulse Responses in Different Monetary Policy Phases

Recent studies highlight the concept that the size of the multiplier depends on the state of the economy because government spending shocks are less likely to crowd out private demand in the presence of economic slack (e.g., Auerbach and Gorodnichenko (2012)). Although rapid changes in economic state are difficult to capture within our TVP-VAR framework, government spending shocks during certain economic states can be identified by imposing additional sign restrictions in the spirit of Canova and Pappa (2011). Because we assume time variant contemporaneous relations among variables, we can employ a different set of sign restrictions and implement the RWZ algorithm for each period. By imposing sign restrictions that reflect the economic state of the period, we can capture the state-dependent effects of government spending on a period-by-period basis.

To consider the economic state of each period, we choose the actual monetary policy phase as a proxy measure of economic slack for two main reasons. First, interest rates directly affect the degree of crowding out. As Canova and Pappa (2011) argues, monetary policy is the most relevant factor affecting the size of the multiplier. Unless monetary policy accommodates fiscal policy in a coordinated manner, government spending leads to substantial crowding out of private demand. Second, although output and prices are natural candidates for measures of economic slack, we want to make our analysis agnostic on these variables because their responses to government shocks are of interest. We define expansionary and contractionary monetary policy phases as the periods during which the detrended interest rate series declines and increases by more than 0.25 percentage points for two consecutive quarters, respectively.[22] We calculate the multipliers during expansionary or contractionary monetary policy phases by adding negative or positive restrictions on interest rates

[22] Auerbach and Gorodnichenko (2012) and Bachmann and Sims (2012) use the seven-quarter moving average of the output growth rate as an index that changes the probability of the economic state. Ramey and Zubairy (2017) use a 6.5 percent unemployment rate as the threshold value to define high and low unemployment states.

to the set of sign restrictions presented in Table 1 in accordance with these criteria. For periods distinguished as neither expansionary nor contractionary monetary policy phases, we use the same sign restrictions as those used in the previous section.

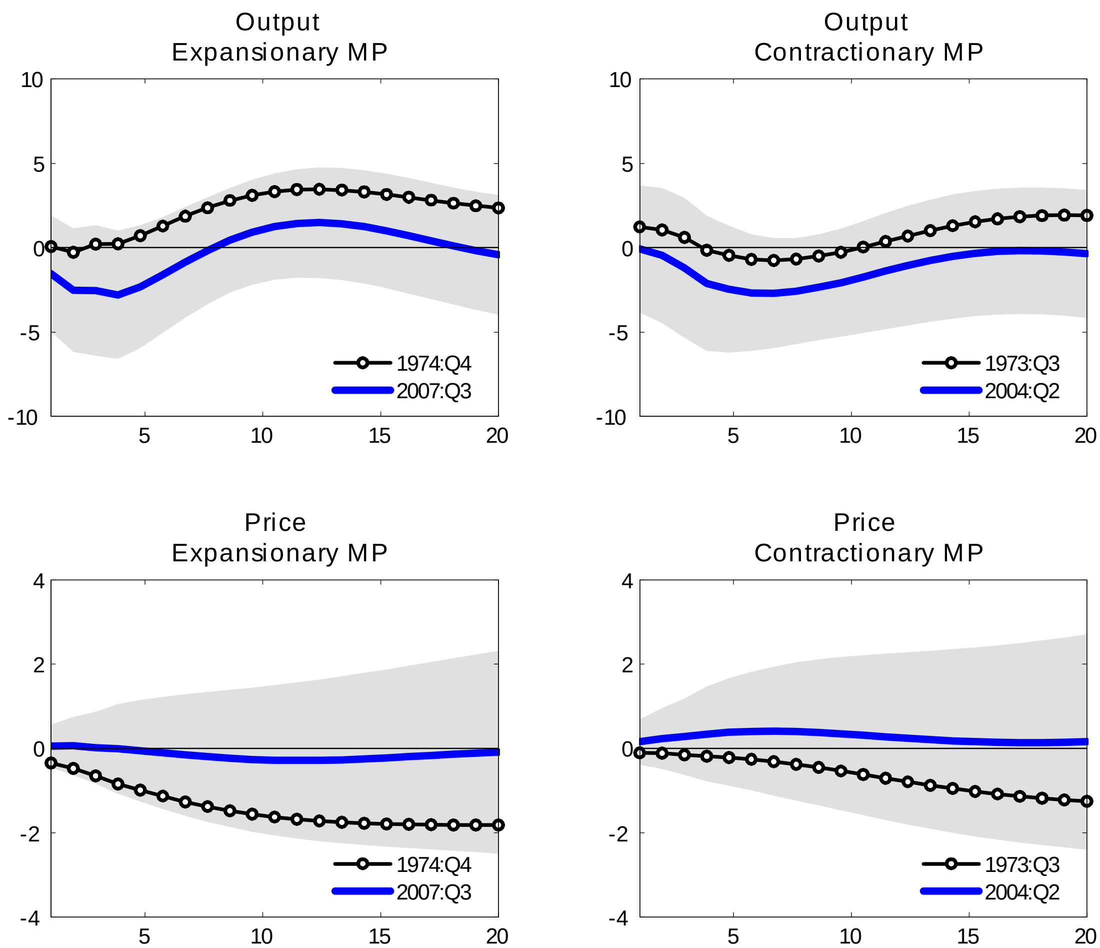

FIG. 5. The output and price responses to an increase in government spending during different monetary policy phases. *Notes:* The upper [lower] panels display responses to a one-dollar [one-percent] increase in government spending. The solid lines in the left [right] panels represent posterior mean responses in 2007:Q3 [2004:Q2] with the shaded areas representing the 16th-84th percentile ranges. The solid lines with circles in the left [right] panels represent posterior mean responses in 1974:Q4 [1973:Q3].

Figure 5 presents output and price responses to government spending shocks under different monetary policy phases. We present the responses of the periods when peak multipliers take the largest and smallest values. As expected, government spending shocks during expansionary monetary policy phases have larger output effects than those during contractionary monetary policy phases. When monetary policy is expansionary, output shows a hump-shaped pattern of increase after a government spending shock. In contrast, output declines immediately after a government spending shock if the Fed conducts a contractionary monetary policy. The counterintuitive negative response of prices can be observed more clearly when monetary policy is expansionary, whereas the effects became unclear in 2000s regardless of the phases of monetary policy. Under both monetary policy phases, output and price effects of government spending show similar time variations to those we have seen in previous section.

TABLE 2
Multipliers

	Average		Range	
	Peak	Cumulative	Peak	Cumulative
Recursive	1.50	1.19	1.19 – 1.93	0.82 – 1.54
Sign restrictions	1.40	0.35	0.13 – 2.66	-1.71 – 1.55
Sign restrictions (MP considered)	1.44	0.24	-0.08 – 3.47	-1.71 – 2.00
Recession periods	2.13	0.84	0.47 – 3.47	-1.06 – 2.00
Expansion periods	1.28	0.12	-0.08 – 3.38	-1.71 – 1.98
Pre-Volcker period	1.91	0.83	0.60 – 3.47	-0.70 – 2.00
Post-Volcker period	1.08	-0.22	-0.08 – 3.23	-1.71 – 1.79

Notes: The table shows the averages and ranges of the peak and cumulative multipliers calculated over either the entire sample or the subsample periods (recession, expansion, pre- and post-Volcker) for different identification schemes. The peak and cumulative multipliers are evaluated at a five-year horizon for each period using the posterior mean impulse responses to a one-dollar increase in government spending. The cumulative multipliers are calculated as the cumulative change in output over the cumulative change in government spending.

4.2 Cyclical and Structural Variations in the Effects of Government Spending

To investigate further the time variation in the effects of government spending shocks, we compute the cumulative responses of output, prices, and interest rates to government spending shocks. Figure 6 shows the time profiles of their point estimates. For the cumulative output multiplier, we find both cyclical and structural variations: multipliers are large in recessions and have declined since the 1980s, in line with findings from existing studies. Table 2 reports the averages of the peak and cumulative multipliers and their ranges over either the entire sample or a subsample period. The average peak multiplier during a recession period is 2.13, which is similar in size to that reported in Auerbach and Gorodnichenko (2012), whereas the average cumulative multiplier is less than one across both recession and expansion periods, in line with the findings in Ramey and Zubairy (2017). The substantial difference observed between the cumulative multipliers in the pre- and post-Volcker periods corroborates the findings in Perotti (2004) and Bilbiie, Meier and Müller (2008).

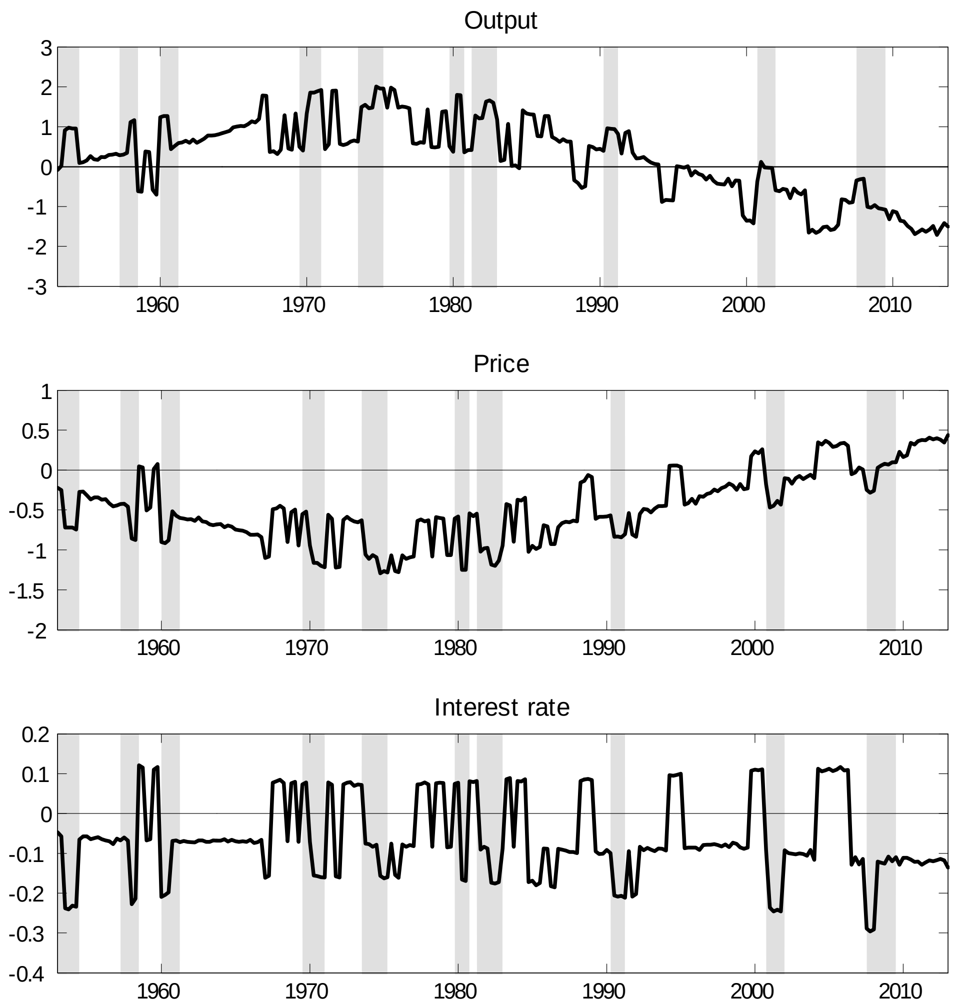

FIG. 6. The evolution of cumulative responses of output, prices, and interest rates to government spending shocks evaluated at a five-year horizon. *Notes:* The cumulative responses are calculated as the cumulative changes in output [prices and interest rates] over the cumulative change in government spending after its one-dollar [one-percent] increase using the posterior mean responses estimated by considering the monetary policy phase of the period. The shaded areas represent recessions as defined by the NBER.

The middle panel of Figure 6 displays the point estimates of cumulative responses of prices to government spending shocks. The price responses show both cyclical and structural variations, as in the case of output multipliers. Government spending shocks during expansionary monetary policy phases are accompanied by stronger deflationary effects, regardless of their stronger impact on output than those during contractionary phases. At the same time, price responses have trended upward since the 1980s. Although the counterintuitive negative response depicted in Figures 4 and 5 appear during most of the sample period, relatively strong inflationary effects are observed during the boom periods of the 2000s and 2010s.

In contrast, the cumulative interest rate response does not show any clear trend, as indicated in the bottom panel of Figure 6. Together with the little time variation in changes in transmission presented in Figure 2, monetary policy does not seem to have played a major role in the declining trend in multipliers since the 1980s.

Notice that both the downward trend in the output effects and the upward trend in the price effects of government spending appear during the period in which the debt-to-output ratio has consistently increased. Their strong correlations are illustrated in Figure 7. Recent studies of debt-dependent fiscal policy effects attribute the cause of the decline in the multiplier to a Ricardian channel, in which households reduce consumption in anticipation of a future tax increase in response to public debt accumulation (e.g., Bi, Shen and Yang (2016)). In contrast, the FTPL literature emphasizes fiscal effects on inflation. According to Cochrane (1999), the analytical content of FTPL can be summarized in the following version of the intertemporal budget constraint on the government:

$$\frac{\text{nominal debt}}{\text{price level}} = \text{present value of real surpluses.} \tag{8}$$

The FTPL literature suggests that there should be a price level increase to reduce the real value of nominal debt when an increase in government spending has a negative impact on the present value of real surpluses. We can infer from equation (8) that the degree of fiscal inflation depends on the amount of nominal debt outstanding: as the government accumulates more nominal debt, the price level increase attributable to the deterioration in real surpluses must be larger.

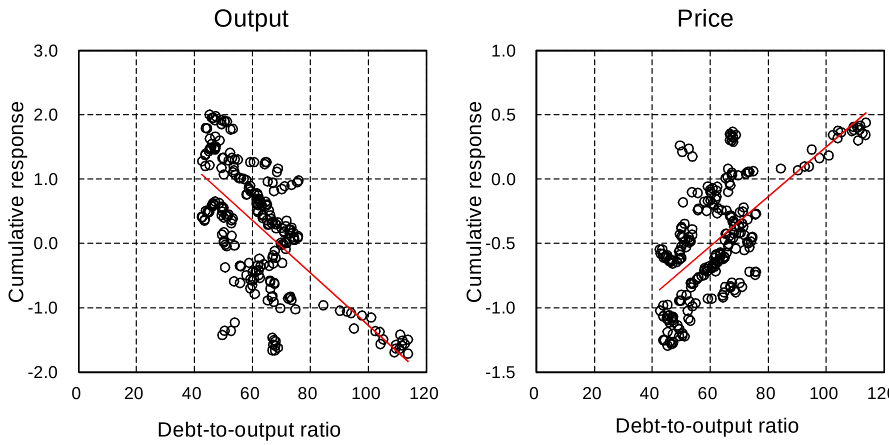

FIG. 7. The correlation between the output and price effects of government spending and debt-to-output ratio. *Notes:* The left [right] panel plots the debt-to-output ratio and cumulative responses of output [prices] to a one-dollar [one-percent] increase in government spending. The cumulative responses are calculated as the cumulative changes in output and prices over the cumulative change in government spending using the posterior mean responses estimated by considering the monetary policy phase of the period. R-squared: 0.436 (left); 0.463 (right).

The positive correlation between the price effects of government spending and debt accumulation may give the impression that debt-driven inflation suggested by the FTPL is already materialized and that a Ricardian fiscal regime is not in place. However, a debt-stabilizing fiscal policy need not be in effect within a finite period to meet the requirements for a Ricardian regime, as Canzoneri, Cumby and Diba (2001) and Canzoneri, Cumby and Diba (2011) argue. Therefore, we cannot reach a conclusion about the prevalence of either a Ricardian or non-Ricardian fiscal regime from observing only the debt-dependency of the price effects of government spending.

4.3 Testing for a Regime Change

We now turn our attention to the role of fiscal policy regimes in the time-varying effects of government spending. Although the empirical literature on fiscal regimes was developed by applying the analytical framework of Leeper (1991), we must assume linearized monetary and fiscal policy rules. Therefore, we take a different path by employing the methodology in Canzoneri, Cumby and Diba (2001) and Canzoneri, Cumby and Diba (2011), which allows us to examine the prevalence of either the Ricardian or the non-Ricardian fiscal regimes without assuming any particular type of policy rules. They estimate a linear bivariate VAR model in Surplus/GDP and Liabilities/GDP on post-war U.S. data and show that a Ricardian fiscal regime is more plausible. Their VAR-based methodology is attractive because we can easily extend it to our TVP-VAR framework.

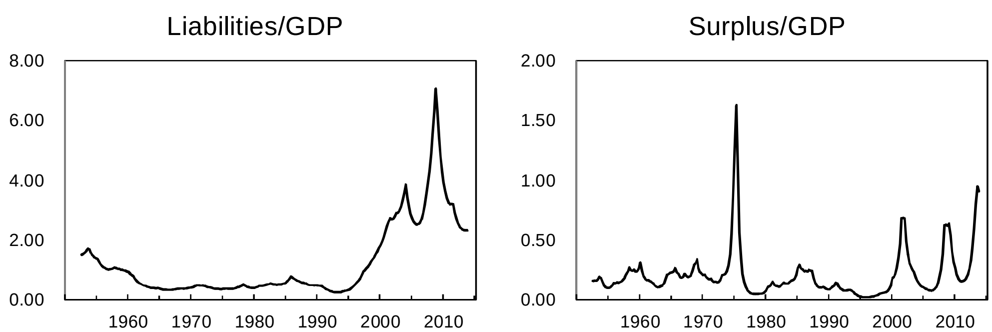

FIG. 8. Posterior mean volatilities.

We estimate a bivariate TVP-VAR model with two lags in Surplus/GDP and Liabilities/GDP. See Appendix A for a detailed description of the data sources. Figure 8 presents the point estimates for stochastic volatilities. The overall results for the volatility of Surplus/GDP shocks effectively capture the fiscal events and indicate a pattern similar to the estimate of tax shocks by Gonzalez-Astudillo (2013). The stochastic volatility of Surplus/GDP shocks increased the most around the time of the Tax Reduction Act of 1975, and increased during tax reforms and measures such as the Reagan Tax Reform of 1981 and 1986, the Bush Tax Cuts of 2001 and 2003, and the American Recovery and Reinvestment Act of 2009.

Figure 9 presents the impulse responses of Surplus/GDP and Liabilities/GDP to an increase in Surplus/GDP. The Surplus/GDP and Liabilities/GDP are ordered first in the left and the right panels, respectively. The former is consistent with a non-Ricardian regime and the latter makes more sense in a Ricardian regime. Because the point estimate of Liabilities/GDP declined the least and the most in 1972:Q1 and 2007:Q3, respectively, when Liabilities /GDP is ordered first, Figure 9 compares the impulse responses in these periods. Regardless of the ordering, Liabilities/GDP declined for several years in response to a Surplus/GDP shock across different sample dates. It is also shown that one unit of deterioration in surplus leads to a larger decline in Liabilities/GDP in 2007:Q3 than in 1972:Q1. Note that the decline in Liabilities/GDP can be observed throughout the estimation period and that the degree of decline shows a widening trend since the late 1970s.

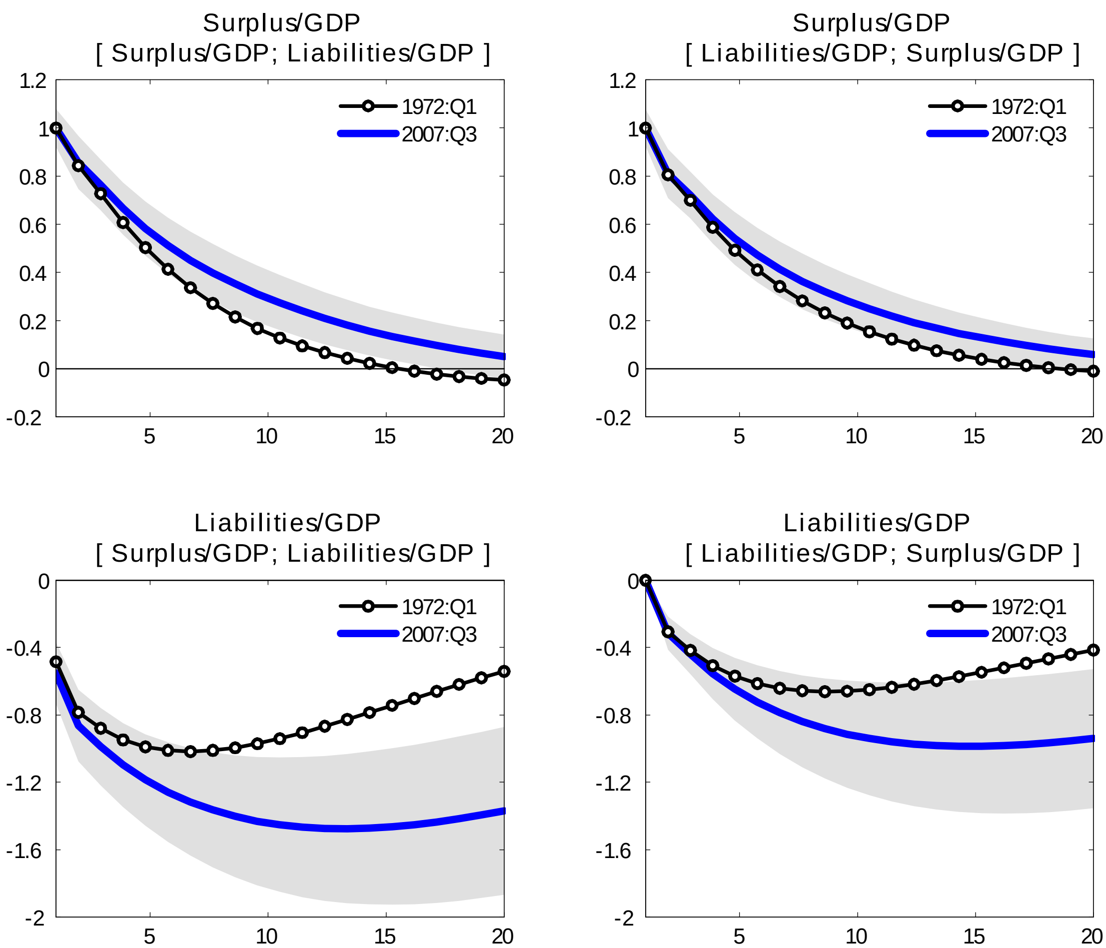

FIG. 9. Evolution of surplus and debt dynamics. *Notes:* The figure shows responses of Surplus/GDP and Liabilities/GDP to a one-percentage point increase in Surplus/GDP. The solid lines represent posterior mean responses for 2007:Q3 with the shaded areas representing the 16th-84th percentile ranges. The solid lines with circles represent posterior mean responses for 1972:Q1. Surplus/GDP is ordered first in the left column and is reversed in the right column.

The results are very similar to those obtained by Canzoneri, Cumby and Diba (2001) and Canzoneri, Cumby and Diba (2011), suggesting that the U.S. government followed a Ricardian fiscal regime throughout the post-war period. As Canzoneri, Cumby and Diba (2001) and Canzoneri, Cumby and Diba (2011) discuss, a non-Ricardian explanation is implausible because it requires a negative correlation between present and future surpluses, which we cannot observe. Furthermore, our application of their VAR-based methodology to a TVP-VAR model reveals that the degree of the government's Ricardian behavior was strengthened during most of the post-Volcker period. Because the debt-to-output ratio increased steadily during the period, the government's stronger Ricardian behavior can be attributed to its debt accumulation. The nonlinear relationship between the government's corrective action and the level of public debt is in line with the findings in Bohn (1998), who provides evidence that the marginal response of the U.S. surplus to changes in debt is an increasing function of the debt level.

4.4 Discussion

Our next question is whether the observed strengthening of the government's Ricardian behavior is the cause of the decline in multipliers. We therefore examine the Granger-causality between the cumulative responses of liabilities to surplus shocks and the cumulative multipliers. Note that monetary policy phase is not considered here in calculating the cumulative multipliers because we found little evidence that monetary policy plays a role in the decline in multipliers. To cope with possible non-stationarity of the variables, we employ the procedure of Toda and Yamamoto (1995) to test for their causal relationship. The first step of the procedure is to select the optimal lag length (k) of the VAR model in levels. The Akaike information criterion (AIC) and the Schwarz information criterion (SIC) both suggest $k = 3$. As a second step, we conduct the Augmented Dickey-Fuller (ADF) unit root test to determine the maximum order of integration ($d_{\max}$) that might occur in the model. Toda and Yamamoto (1995) show that we can test restrictions on the first k coefficient matrices of a $(k + d_{\max})$th-order VAR model in levels using the standard asymptotic theory, even if the variables are integrated or cointegrated. Letting $d_{\max} = 1$ based on the ADF test results reported in Table 3, we test the null hypothesis of no Granger causality using a standard Wald statistic for the first k coefficient matrices of the $(k+d_{\max})$th-order VAR model in levels. As shown in Table 3, the results suggest that the cumulative responses of liabilities appear to Granger-cause the cumulative multipliers, while the null hypothesis of no-Granger-causality in the opposite direction cannot be rejected. Considering the negative correlation between the debt-to-output ratio and the multipliers observed in Figure 7, we conjecture that the government's corrective action strengthened in the presence of higher indebtedness, thus serving as the major driving force for the observed decline in government spending multipliers. This explanation shares the view of the relationship between debt and multipliers in existing studies (e.g., Sutherland (1997); Perotti (1999); Bi, Shen and Yang (2016)).

TABLE 3
Unit-root and causality test results

Null hypothesis	Test statistics and p-values			
ADF	t-statistic at level		t-statistic at first difference	
LIAB has a unit root	-3.2081	(0.0853)*	-4.2524	(0.0044)**
MULT has a unit root	-2.6937	(0.2402)	-16.6765	(0.0000)**
Granger (Toda-Yamamoto procedure)	Wald chi-square test statistic			
LIAB does not Granger-cause MULT	10.7188	(0.0299)**		
MULT does not Granger-cause LIAB	6.8931	(0.1416)		

Notes: LIAB stands for the cumulative response of liabilities to surplus shocks calculated as the cumulative change in Liabilities/GDP over the cumulative change in Surplus/GDP using the posterior mean responses estimated in subsection 4.3. MULT stands for the cumulative multiplier calculated as the cumulative change in output over the cumulative change in government spending using the posterior mean responses estimated for the sign restriction identification scheme in subsection 3.2. Both LIAB and MULT are evaluated at a five-year horizon. Figures between parentheses are p-values. A double asterisk (**) denotes significant at the five percent level; a single asterisk (*) denotes significant at the ten percent level.

Notably, the strengthening in the government's Ricardian behavior occurred soon after the passage of the Congressional Budget and Impoundment Act of 1974, which established the Congressional Budget Office. Since then, Congress introduced a variety of budget rules in an attempt to impose fiscal discipline on the budgetary process. By examining the effects of budget rules, Auerbach (2008) concludes that these rules appear to have had some success with deficit control. The institutional change might have contributed to raising expectations of the future tax burden in the face of public debt accumulation, thereby leading to smaller multipliers.

Then, how can we explain the debt-dependency of the price effects of government spending observed in Figure 7, which is more plausible under a non-Ricardian fiscal regime? Leeper (2013) and Davig and Leeper (2011*b*) discuss the possible fiscal effects on inflation before a regime change when the probability of reaching the fiscal limit is increasing. However, they assume that inflation occurs as a result of the demand shift from government bonds to goods induced by heightened expectations of an increased supply of government bonds. This channel must be accompanied by a materialized output increase and therefore is difficult to reconcile with our results, in which multipliers show a declining trend along with an accumulation of public debt. In contrast, Cochrane (2011) considers the upward shift of the Phillips curve driven by heightened inflation expectations as a source of fiscal inflation. He argues that fiscal inflation is likely to be accompanied by stagflation and not a boom. This channel provides a more plausible explanation of our results. Although higher taxes are anticipated, they reduce the capacity for further tax increase at the same time, thereby increasing the probability of hitting the fiscal limit. A rise in inflation expectations can occur when people believe that a government spending increase will not be fully financed by future tax. If this is the case, a narrowing fiscal space may be the cause of the observed debt-dependent price effects of government spending. This interpretation is also consistent with the findings in De Graeve and Queijo von Heideken (2015), who suggest an increasing contribution of fiscal inflation anticipation to inflation along with a public debt accumulation.

5 Conclusion

This study provides new empirical evidence on the evolution of the output and price effects of government spending during the post-war period in the United States. We obtain large multipliers in recessions and small ones in the post-Volcker period, in line with the findings of existing studies. Whereas the empirical evidence for the negative correlation between debt and multipliers was established for cross-country data, this study provides it by analyzing U.S. time series data. The results also suggest that the inflationary effects of government spending became stronger along with an accumulation of public debt. Testing for changes in the fiscal policy regime, we further find that the degree of the government's Ricardian behavior was strengthened during most of the post-Volcker period. The stronger corrective action in the face of rising indebtedness is shown to Granger cause the decline in the multipliers.

Although we tackled a relatively wide range of topics, the findings are largely consistent with those documented in various strands of the literature. This study differs from previous ones in that we examine time variation in both the output effects and the price effects of government spending. We address their cyclical and structural variations simultaneously within a single TVP-VAR framework using the assistance of sign restriction identification. In particular, the study contributes to the FTPL literature by providing evidence of the debt-dependent price effects of government spending, whereas fiscal inflation is difficult to detect in the post-war U.S. data. Another contribution of the paper is the application of TVP-VAR technique to demonstrate the changes in the government's Ricardian behavior.

Nevertheless, there remains much work ahead. Although our atheoretical VAR-based approach is a flexible way to model the evolution of time series data, it has limitations in explaining the underlying mechanism. It would be worth exploring the development of a theoretical model that accounts for the time variation in the output and price effects reported in this paper. Investigating the effects of fiscal inflation anticipation on prices would be one of the most important topics. Extending the analysis to the period of monetary policy normalization, which requires a model of quantitative easing, is another interesting avenue to explore.

A Description of Data Sources

We obtain all quarterly data from the Federal Reserve Bank of St. Louis's FRED database. Seasonally adjusted series for real government spending, real gross domestic product, and real private consumption are Real Government Consumption Expenditures and Gross Investment (GCEC96), Real Gross Domestic Product (GDPC1), and Real Personal Consumption Expenditures (PCECC96), respectively. To convert the series into per-capita terms, we divide them by the seasonally adjusted Civilian Labor Force (CLF16OV). The ratios of output and consumption to government spending used to calculate the multipliers are constructed from seasonally adjusted series for Gross Domestic Product (GDP), Personal Consumption Expenditures (PCEC), and Government Consumption Expenditures and Gross Investment (GCE), respectively. We use the seasonally adjusted GDP Chain-type Price Index (GDPCTPI) as the price level. We use the 3-Month Treasury Bill Secondary Market Rate (TB3MS) as the nominal interest rate. The debt-to-output ratio is calculated by dividing the sum of federal, state, and local government liabilities by the seasonally adjusted Gross Domestic Product (GDP). We use the seasonally adjusted liabilities of the Federal government (FGDSLAQ027S) and those of the State and local governments (SLGLIAQ027S) in the calculation. The Surplus/GDP is calculated by dividing the seasonally adjusted primary surplus by the Gross Domestic Product (GDP). The primary surplus is defined as Net Government Saving (TGDEF) minus the difference between income receipts on assets (W059RC1Q027SBEA) and interest payments (A180RC1Q027SBEA). The Liabilities/GDP is calculated as in the debt-to-output ratio.

B Markov Chain Monte Carlo Methods

This appendix outlines the MCMC algorithm used to estimate the TVP-VAR models presented in this paper. Given the data, the algorithm allows us to sample parameters and hyperparameters from their posterior density. In what follows, x denotes the entire history of x_t to the end of the sample period. Letting $f(x \mid z)$ denote the conditional density of x given z, the MCMC algorithm repeats the following steps:

1. Initialize $\beta, a, h, \Sigma_\beta, \Sigma_a, \Sigma_h$.
2. Draw β from $f(\beta \mid a, h, \Sigma_\beta, y)$.
3. Draw Σ_β from $f(\Sigma_\beta \mid \beta, y)$.
4. Draw a from $f(a \mid \beta, h, \Sigma_a, y)$.
5. Draw Σ_a from $f(\Sigma_a \mid a, y)$.
6. Draw h from $f(h \mid \beta, a, \Sigma_h, y)$.
7. Draw Σ_h from $f(\Sigma_h \mid h, y)$.
8. Go to 2.

For the first step, we set the initial states of the parameters as $\beta_{p+1} \sim N(0, 10I)$, $a_{p+1} \sim N(0, 10I)$, and $h_{p+1} \sim N(0, 50I)$. We postulate an inverse-Gamma distribution for the m-th diagonal elements of the covariance matrices. The priors are specified as $(\Sigma_\beta)_m^2 \sim IG\left(10, 10^{-6}\right)$, $(\Sigma_a)_m^2 \sim IG\left(5, 10^{-3}\right)$, and $(\Sigma_h)_m^2 \sim IG\left(5, 10^{-3}\right)$. We execute 30,000 MCMC replications and discard the first 5,000 draws to estimate the TVP-VAR models. To reduce autocorrelation among the draws, we save only every fifth draw.

B.1 Drawing β

For notational convenience, we rewrite equation (1) as

$$y_t = X_t \beta_t + A_t^{-1} \Sigma_t \varepsilon_t, \tag{9}$$

where $X_t = I_k \otimes \left[y'_{t-1}, ..., y'_{t-k} \right]$ and $\otimes$ denotes the Kronecker product. Because the observation equation (B1) and state equation (4) constitute a linear Gaussian state-space representation for the dynamic behavior of y_t, we can apply the simulation smoother of de Jong and Shephard (1995) to draw samples of β from its posterior density conditioned on a, h, Σ_β, and y.

B.2 Drawing a

We can write equation (B1) as

$$A_t \left(y_t - X_t \beta_t \right) = \Sigma_t \varepsilon_t. \tag{10}$$

Let $\widehat{y}_t = y_t - X_t \beta_t$ and $\widehat{X}_t$ be the $k \times \frac{k(k-1)}{2}$ matrix defined by

$$\widehat{X}_t = \begin{bmatrix} 0 & \cdots & \cdots & 0 \\ -\widehat{y}_{1,t} & 0 & \cdots & 0 \\ 0 & -\widehat{y}_{[1,2],t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & -\widehat{y}_{[1,...,k-1],t} \end{bmatrix},$$

where $\widehat{y}_{[1,...,q],t}$ represents the row vector $[\widehat{y}_{1,t}, \widehat{y}_{2,t}, ..., \widehat{y}_{q,t}]$, where $q \leq k-1$. We can express equation (B2) as

$$\widehat{y}_t = \widehat{X}_t a_t + \Sigma_t \varepsilon_t. \tag{11}$$

Because the observation equation (B3) and state equation (5) can be treated as a linear Gaussian state-space model, assuming that Σ_a and Σ_{a_0} are diagonal matrices, we can apply the simulation smoother of de Jong and Shephard (1995) to draw samples of a from its posterior density conditioned on β, h, Σ_a, and y.

B.3 Drawing h

For the stochastic volatilities, we apply the multi-move sampler developed by Shephard and Pitt (1997) and Watanabe and Omori (2004) because the system of equations consists of (B1), and (6) is not linear in h. The diagonality assumptions of Σ_h and Σ_{h_0} allow us to make inferences on $\{h_{j,t}\}_{t=p+1}^{T}$ separately for $j = 1, ..., k$. Let $y^*_{j,t}$ be the j-th element of $A_t \widehat{y}_t$. Now, consider the following system of equations:

$$y^*_{j,t} = \exp\left(\frac{h_{j,t}}{2}\right)\varepsilon_{j,t}, \tag{12}$$

$$h_{j,t+1} = h_{j,t} + \eta_{j,t}, \tag{13}$$

$$\begin{pmatrix}\varepsilon_{j,t}\\ \eta_{j,t}\end{pmatrix} \sim N\left(0, \begin{pmatrix} 1 & 0 \\ 0 & \nu_j^2 \end{pmatrix}\right),$$

where $\varepsilon_{j,t}$ and $\eta_{j,t}$ are the j-th elements of ε_t and $u_{h,t}$, respectively, and ν is the j-th diagonal element of Σ_h. The prior distribution for the initial value is given by $\eta_{j,p} \sim N(0, \nu^2_{j,o})$, where $\nu^2_{j,o}$ is the j-th diagonal element of Σ_{h_0}.

Drawing samples of h conditional on β, a, Σ_h, and y is difficult because of an analytically intractable form of its posterior density. One way is to draw each sample of h_t conditional on $h_{\backslash t}$, β, a, Σ_h, and y; however, the method tends to produce a highly correlated sample sequence.[23] Therefore, we divide the state variables $\{h_{j,t}\}_{t=p+1}^{T}$ into $K + 1$ blocks and draw each block conditional on the elements of the other blocks and parameters. Let the end elements of the blocks be

[23] See Shephard and Pitt (1997) and Kim, Shephard and Chib (1998).

h_{j,k_n} for $n = 1, ..., K$. The end conditions of blocks k_n, called "stochastic knots," are determined randomly over iterations. To cope with possible degeneracies, we draw $\eta_{j,p+1}, ..., \eta_{j,T-1}$ instead of $h_{j,p+2}, ..., h_{j,T}$, which can be constructed using (B5) given the sampled $h_{j,p+1}$. Suppose we draw samples from a typical block $h_{j,r}, ..., h_{j,r+d}$, where $r \geq p+1$, $d \geq 1$, and $r + d \leq T$. By Bayes' theorem, the posterior conditional density of a block of disturbances can be expressed as

$$f(\eta_{j,r-1}, ..., \eta_{j,r+d-1} \mid h_{j,r-1}, h_{j,r+d+1}, y^*_{j,r}, ..., y^*_{j,r+d}, \nu_j, \nu_{j,o}) \tag{14}$$

$$\propto \prod_{t=r}^{r+d} \frac{1}{e^{h_{j,t}/2}} \exp\left(-\frac{y^{*2}_{j,t}}{2e^{h_{j,t}}}\right) \times \prod_{t=r-1}^{r+d-1} f\left(\eta_{j,t}\right) \times f\left(h_{j,r+d}\right),$$

where

$$f\left(\eta_{j,t}\right) = \begin{cases} \exp\left(-\frac{\eta^2_{j,p}}{2\nu^2_{j,o}}\right) & (t = p), \\ \exp\left(-\frac{\eta^2_{j,t}}{2\nu^2_j}\right) & (t \geq p+1), \end{cases}$$

$$f\left(h_{j,r+d}\right) = \begin{cases} \exp\left(-\frac{\left(h_{j,r+d+1} - h_{j,r+d}\right)^2}{2\nu^2_j}\right) & (r + d < T), \\ 1 & (r + d = T). \end{cases}$$

To draw $\eta_{j,r-1}, ..., \eta_{j,r+d-1}$ from the density (B6), we consider a proposal density expressed in logarithmic form by taking the logarithm of (B6) and applying the second-order Taylor approximation to

$$g\left(h_{j,t}\right) \equiv -\frac{h_{j,t}}{2} - \frac{y^{*2}_{j,t}}{2e^{h_{j,t}}}$$

around a certain point $h_{j,t} = \widehat{h}_{j,t}$, which we choose to be near the mode of the posterior density. We can sample from the proposal density by defining artificial variables

$$h^*_{j,t} = \begin{cases} \sigma^*_{j,t}\left(g'(\widehat{h}_{j,t}) - g''(\widehat{h}_{j,t})\widehat{h}_{j,t} + \frac{h_{j,t+1}}{\nu^2_j}\right) & (t = r + d < T), \\ \widehat{h}_{j,t} + \sigma^*_{j,t} g'(\widehat{h}_{j,t}) & (t = r, ..., r + d - 1 \text{ and } t = r + d = T), \end{cases}$$

where

$$\sigma^*_{j,t} = \begin{cases} \frac{\nu_j^2}{1-g''(\widehat{h}_{j,t})\nu_j^2} & (t = r + d < T), \\ -\frac{1}{g''(\widehat{h}_{j,t})} & (t = r, ..., r + d - 1 \text{ and } t = r + d = T), \end{cases}$$

and then considering the following equation

$$h^*_{j,t} = h_{j,t} + \zeta_{j,t}, \tag{15}$$

where $\zeta_{j,t} \sim N(0, \sigma^*_{j,t})$. Using (B5) and (B7), we can formulate a linear Gaussian state-space model and draw samples applying the simulation smoother of de Jong and Shephard (1995). The sampling is the same as that for the proposal density. Hence, we use the Accept-Reject Metropolis-Hastings algorithm in Tierney (1994) to produce draws from the correct density (B6) as the sampling process is iterated.

B.4 Drawing hyperparameters

Let $\theta = (\beta, a, h)$, m-th element of θ_t be $\theta_{m,t}$, and the prior for the m-th diagonal element of the covariance matrix of θ be given by $(\Sigma_\theta)^2_m \sim IG(s_{\theta_0}/2, S_{\theta_0}/2)$. Because we assume that Σ_θ is a diagonal matrix, the m-th diagonals of covariance matrix $(\Sigma_\theta)^2_m$ can be sampled independently. The posterior density conditioned on θ is then given by $(\Sigma_\theta)^2_m \mid \theta \sim IG(\widehat{s}_{\theta_m}/2, \widehat{S}_{\theta_m}/2)$, where

$$\widehat{s}_{\theta_m} = s_{\theta_0} + T - p - 1,$$

$$\widehat{S}_{\theta_m} = S_{\theta_0} + \sum_{t=p+1}^{T-1} (\theta_{m,t+1} - \theta_{m,t})^2 .$$

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