The Government Spending Multiplier at the Zero Lower Bound: Evidence from the United States

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Abstract

We estimate state-dependent government spending multipliers for the United States. We use an Interacted Vector Autoregression (IVAR) model to capture the time-varying monetary policy characteristics including the recent zero interest rate lower bound (ZLB) state. We identify government spending shocks by sign restrictions and use a government spending growth forecast series to account for the effects of anticipated fiscal policy. In our baseline specification we find that government spending multipliers range from 3.4 to 3.7 at the ZLB. Away from the ZLB, multipliers range from 1.5 to 2.7. Next, we address the limited information problem typically inherent in VARs by the help of a Factor-Augmented IVAR (FAIVAR). We find that multipliers are lower in this case, ranging from 2.0 to 2.1 at the ZLB and between 1.5 and 1.8 away from it. Thus, in both specifications we find that multipliers are higher, when the interest rate is lower. Our results are consistent with recent theories that predict larger multipliers at the ZLB.

Keywords
Interacted VAR, Fiscal Policy, Government Spending, Zero Interest Rate Lower Bound

JEL Classification
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Introduction

How large is the government spending multiplier in normal times and how large is it when monetary policy is constrained by the zero interest rate lower bound (ZLB)? The Great Recession has revived the debate regarding this question among policy circles and in academia as it is of high practical relevance. If fiscal stimulus by means of an increase in government spending raises real GDP by more than one-for-one, i.e., each dollar of the government spending increase raises real GDP by more than one dollar, then such a stimulus is highly desirable from a policymaking perspective.

The recent debate has given particular attention to the fact that since the outbreak of the 2008 financial crisis the Fed’s monetary policy was accommodative, or, even constrained by the ZLB. It is worthwhile that the accommodative stance also included unconventional monetary policy. Figure 1 illustrates monetary and fiscal policy from 1960Q1 to 2015Q4. The key observation regarding the most recent recession is that the Federal Funds Rate was abruptly cut to near zero and has remained there until 2015Q4. Moreover, there has been a dramatic deficit-financed increase in government expenditures during this period. It is frequently argued that in such an extraordinary situation, an increase in government spending is even more effective than in normal times.

A growing theoretical literature examines this claim. There is an increasing number of New Keynesian DSGE models that generates predictions consistent with this claim. See, for instance, Christiano et al. (2011), Eggertsson (2010), Woodford (2011), Davig and Leeper (2011), or, Coenen et al. (2012). These models predict a government spending multiplier in the range of 3 to 5. Likewise, there is an emerging literature developing reasonable theories that suggest that the government spending multiplier at the ZLB is one or below, and lower than in times without the ZLB binding. See, for instance, Braun et al. (2013), Mertens and Ravn (2014), Aruoba et al. (2017).

Given the wide range of theoretical predictions for the size of the government spending multiplier at the ZLB, empirical evidence is a crucial need for policymakers and academia.

However, the empirical literature providing state-dependent evidence on the size of the aggregate government spending multiplier at the ZLB is still in its infancy. To date, Ramey and Zubairy (2017) is the single paper in this literature according to our knowledge. Ramey and Zubairy (2017) use the local projection method developed by Jordà (2005) and find that the government spending multiplier at the ZLB can be as large as 1.5 in some specifications. Moreover, there is a related, but distinct empirical literature quantifying state-dependent fiscal multipliers in recessions based on regime-switching VAR models. However, as Figure 1 illustrates, recessions and episodes where the ZLB is binding do not necessarily coincide. Thus, there is a need for more evidence on the government spending multiplier at the ZLB. The objective of this paper is to provide further state-dependent evidence on the size of the government spending multiplier at the ZLB from the United States. We extend the literature by proposing an alternative framework to quantify the state-dependent government spending multiplier. To this end we use an Interacted Vector Autoregression model (IVAR) building on the panel IVAR in Towbin and Weber (2013) and Sá et al. (2014). The interaction term allows us to derive impulse response functions (IRFs) to a government spending shock at different percentiles of the interest rate distribution. This methodology enables us to investigate among the entire range of historical interest rates for the sample considered: within the same setup, we are capable of computing multipliers for median and low levels of the interest rate distribution, with no need to restrict the sample.

1 For instance, the Fed announced three rounds of quantitative easing: in November 2008, in November 2010, and in September 2012.
2 Christiano et al. (2011, p.81) argue: ‘The simple models discussed above suggest that the multiplier can be large in the zero-bound state. The obvious next step would be to use reduced-form methods, such as identified VARs, to estimate the government-spending multiplier when the zero bound binds.’
3 Crafts and Mills (2013) and Ramey (2011b) provide evidence for ZLB episodes suggesting multipliers below unity.
By using the IVAR framework, we can address several potentially problematic issues of alternative frameworks that are used in the literature on state-dependent multipliers. For instance, compared to regime-switching approaches in general, such as Threshold VAR (TVAR) methods, and, the Ramey and Zubairy (2017) approach in particular, the IVAR model does not require to define a particular threshold. Regime-switching approaches use such a threshold to distinguish observations of normal times from ZLB episodes. However, such a threshold is subject to discretion. In addition, the IVAR uses all the information available for the full sample, while a threshold model uses the information of each state under consideration separately. Moreover, the IVAR does not rely on a particular assumption on an approximation of monetary policy over the sample period, i.e., a Taylor (1993)-rule. For instance, parts of the theoretical literature regard the contemporaneous-data Taylor-rule applied in Ramey and Zubairy (2017) as a problematic approximation of monetary policy, because it is not operational. For instance, real-time data is hardly available even for central banks, see McCallum (1999) for a discussion. Finally, an interest rate value implied by an ex post application of a Taylor (1993)-rule below some threshold does not necessarily mean that the economy is at the ZLB. Ramey and Zubairy (2017, pp.23-24) are aware of this point and then eliminate certain episodes on a discretionary basis.

An alternative may be to consider a Smooth Transition VAR (STVAR) framework as used by Caggiano et al. (2015) to estimate the government spending multiplier in recessions. However, there are also concerns regarding an STVAR approach that do not apply to the IVAR model. First, the STVAR, similar to a threshold model, allows only for a finite number of states in practice. Second, as emphasized in Caggiano et al. (2017, p.11), the change in monetary policy in times of crises is frequently abrupt and not smooth. The STVAR framework is not designed to capture such abrupt changes. In sum, compared to threshold-based approaches or the STVAR framework, the IVAR offers clear advantages. The interaction term can capture abrupt policy changes and allows for a large number of states. The number of states can equal the number of available observations. Another key strength of our empirical strategy is that we identify the government spending shock by using sign restrictions and the series of government spending growth forecasts errors used in Auerbach and Gorodnichenko (2012, 2013). The sign restriction approach allows us to use a minimum of economically meaningful and rather uncontroversial identification restrictions. The forecasts errors enable us to address the concerns related to fiscal foresight in Leeper et al. (2013). The series captures the surprise component in a broad measure of government spending and, as we show, is a relevant and strong instrument for the our post WWII sample. An alternative would be to consider the defense news series used in Ramey and Zubairy (2017). However, this is a rather narrow measure that captures just a particular component of government spending. Furthermore, as Ramey (2011b) reports, defense news appears to be a rather weak instrument, when a post WWII sample does not cover the period of the Korean War.

For our sample from 1966Q4 to 2015Q4, we consider two different specifications. Our baseline specification involves the forecast error series, government spending, GDP and the average tax rate. At the ZLB, government spending multipliers are between 3.42 and 3.66. When monetary policy is not constrained by the ZLB, government spending multipliers are between 1.54 and 2.56. Our second specification addresses the generic limited information problem inherent in VARs as a robustness check. On the one side, introducing more and more variables to the VAR adds more information. However, adding additional variables to the VAR implies a loss of degree of freedom. We handle this trade-off by considering a Factor-Augmented IVAR (FAIVAR). Compared to the baseline specification, we obtain lower multipliers in the FAIVAR. Nevertheless, the bottom line result is the same: multipliers are higher when interest rates are lower. At the ZLB multipliers range from 1.98 to 2.10 while multipliers range from 1.48 to 1.79 away from the ZLB. Thus, our results

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4 The sign restrictions approach is developed in Canova and De Nicolò (2002), Uhlig (2005), Mountford and Uhlig (2009) apply it to fiscal policy.
are qualitatively and quantitatively consistent with the claim that increases in government spending are even more effective at the ZLB. The paper proceeds as follows: Section 1 outlines the IVAR model, our baseline specification and data, our inference and identification approach and how we calculate the multipliers; Section 2 discusses the main results; Sections 3 addresses misspecification concerns; Section 4 concludes.

1. Methodology

1.1 Empirical Model

We use an Interacted VAR Model based on Towbin and Weber (2013) and Sá et al. (2014).\(^5\) The recursive-form is given by

\[
B_t Y_t = \kappa + \sum_{k=1}^L \Gamma_k Y_{t-k} + \kappa^1 X_t + \sum_{k=1}^L \Gamma_k^1 X_t Y_{t-k} + \varepsilon_t
\]

(1)

where \(t = 1, \ldots, T\) denotes time and \(k = 1, \ldots, L\) denotes the lag length. \(Y_t\) is a \(q \times 1\) vector which contains explanatory variables, \(\kappa\) is the constant term, \(\Gamma_k\) is a \(q \times q\) matrix of autoregressive coefficients, \(\varepsilon_t \sim N(0, \Sigma)\) is the vector of residuals. Moreover, \(X_t\) denotes the interaction term, which can influence both the dynamic relationship between endogenous variables and their level, through \(\Gamma_k^1\) and \(\kappa^1\) respectively. The matrix \(B_t\) is a \(q \times q\) lower triangular matrix with ones on the main diagonal. Each component \(B_t(w, q)\) represents the contemporaneous effect of the \(q\)th-ordered variable on the \(w\)th-ordered variable. It is constructed as follows:

\[
B_t = \begin{cases} 
B_t(w, q) = 0 & \text{for } q > w \\
B_t(w, q) = 1 & \text{for } q = w \\
B_t(w, q) = B(w, q) + B^1(w, q)X_t & \text{for } q < w 
\end{cases}
\]

(2)

where \(B(w, q)\) and \(B^1(w, q)\) are regression coefficients capturing the marginal effects of a change in the interaction term. The recursive structure imposes that all the variables in the system react contemporaneously to the first ordered variable, but the latter does not react on impact to any other variables. The recursive form of the matrix \(B_t\) also implies that the covariance matrix of the residuals, \(\Sigma\), is diagonal.

1.2 Baseline Specification

Our data set consists of U.S. quarterly data and goes from 1966Q4 to 2015Q4.\(^6\) In our baseline specification our vector \((1)\) of endogenous variables is:

\[
Y_t = [FE_t, G_t, GDP_t, T_t]^\prime.
\]

(3)

This vector \(Y_t\) includes variables that are commonly used in the literature (e.g., Blanchard and Perotti, 2002). \(G_t\) represents real government spending and we use government consumption expenditures and gross investment as a proxy. \(T_t\) denotes the average tax revenue. We use federal government current receipts as a proxy for this variable. Moreover, \(GDP_t\) stands for real gross domestic product. Finally, \(FE_t\) denotes a series of forecast errors of the annualized growth rate of real government spending following Auerbach and Gorodnichenko (2012).\(^7\) By this series we address fiscal

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\(^5\) The exposition follows Sá et al. (2014) although we do not consider a panel of countries.

\(^6\) The choice of this time period is motivated by the availability of the Greenbook and SPF government spending forecasts.

\(^7\) Appendix A contains further information on the computation of this variable.
foresight. In Appendix B we provide evidence that $FE_t$ has high explanatory power regarding the variation in growth of $G_t$ and is therefore a relevant instrument to control for fiscal foresight that cannot be considered weak.\footnote{For further explanations about the fiscal foresight critique see Leeper et al. (2013).}

Variables $G_t$ and $GDP_t$, are expressed in real terms and considered in levels. $T_t$ is in nominal terms and divided by nominal GDP. With the exception of the average tax rate, the other variables and $FE_t$ have been normalized with an estimate of real potential GDP. Ramey and Zubairy (2017) show that the usual approach of using log of variables requires an \textit{ex post} conversion to dollar equivalents of the estimated elasticities that can produce serious bias. The problem is even more acute in nonlinear models and in particular in our model, where several multipliers can be potentially computed, since the \textit{ex post} conversion requires a factor which is based on the sample average of the ratio of GDP to government spending. With the kind of normalization just described, government spending multipliers can be computed directly\footnote{Further details about the computation of the government spending multipliers are described below. For more details on the bias caused by the \textit{ex post} conversion of the elasticities see Ramey and Zubairy(2017).}. Further details about all variables that we use, transformation and so on, are provided in Table 2.

For the interaction term we use the U.S. Shadow Federal Funds Rate developed by Wu and Xia (2016), i.e., $X_t = s_r t-1$. The interaction term allows us to examine how the time-varying interest rate environment affects the transmission mechanism of the government spending shock among the variables in $Y_t$. In particular, we investigate effects of a government spending shock when $s_r t-1$ is at the 1st, 5th, 13th, 25th, 50th and 75th percentile of its distribution. We consider the range from the 1st to the 13th percentile of the Shadow Rate distribution as the \textit{low} interest state, as the 13th percentile coincides with a value of the interest rate equal to 0.25. The latter value is conventionally accepted by the literature as the lower bound for monetary policy in using the Federal Funds rate as instrument. Results for the 25th percentile and above are associated with the \textit{high} interest state. It is important to emphasize that we use this categorization of percentiles in order to structure the discussion of results later on. However, this is not a threshold that affects our results.

We use the U.S. Shadow Federal Funds Rate as this rate is a more precise indicator of monetary policy after the Federal Funds Rate reached the ZLB: away from the ZLB this series is equal to the effective federal funds rate, but at ZLB Wu and Xia (2016) use a Gaussian Affine Term Structure Model (GATSM) to generate an effective rate. Figure 1 illustrates this point. After the abrupt cut in the Federal Funds Rate during the most recent recession, the Federal Funds Rate has been near zero and shows little variation. However, unconventional monetary policy measures have been implemented and the variation in the Shadow Federal Funds Rate in the same period captures this policy. We use first lag of the shadow rate to address potential endogeneity concerns. Specifying $X_t = s_r t-1$ implies that the monetary policy instrument is not endogenous to $Y_t$. If we were to specify $X_t = s_r t$ reverse causality could be a problem.\footnote{Notice that specifying $X_t = s_r t$ does not have a significant effect on our results and conclusions below.}

Finally, notice that we choose a lag length of $L = 1$ in order to preserve the parsimony of the model.\footnote{The lag length has been chosen on the base of the Hannan-Quinn(HQ) and Schwarz-Bayes(SBC) information criteria.}

\subsection{1.3 Inference and Identification}

As in Uhlig (2005) and Sá et al. (2014) we use Bayesian estimation by setting an uninformative normal-Wishart prior, and start with the estimation of the recursive model described in equation (1). Since we know that the covariance matrix $\Sigma$ is diagonal by construction we can proceed by estimating the model equation by equation. For each equation we draw the recursive-form
parameters jointly from the posterior. We evaluate them at a pre-specified value of the interaction term and compute reduced form parameters by inverting the matrix $B_t$:

$$Y_t = B_t^{-1}K + B_t^{-1} \sum_{k=1}^{L} \Gamma_k Y_{t-k} + B_t^{-1}X_t + B_t^{-1} \sum_{k=1}^{L} \Gamma_k Y_{t-k} + B_t^{-1} \epsilon_t$$  \hspace{1cm} (4)

$$Y_t = C + \sum_{k=1}^{L} A_k Y_{t-k} + C'1 X_t + \sum_{k=1}^{L} A_k Y_{t-k} + \epsilon_t$$  \hspace{1cm} (5)

where the vector of residuals $\epsilon_t \sim N(0, \Sigma_t')$ and the Cholesky decomposition of the reduced form covariance matrix is given by $V_t = B_t^{-1} \Sigma_t^{\frac{1}{2}}$.

Government spending shocks are identified by imposing sign-restrictions. Once we have obtained the Cholesky decomposition of the reduced form covariance $V_t$, the general idea is to obtain combinations of $V_t$ by using an orthogonal matrix $Q$ such that $V_t' = QV_t$, where orthogonality of the shocks is preserved. Rubio-Ramírez et al. (2010) propose to draw a $W$ matrix from a $N(0,1)$ and use the QR decomposition (householder transformation), obtaining $W = QR$, where $Q$ is the orthogonal matrix required to impose the sign restrictions that allows to preserve orthogonality of the shocks derived from the Cholesky decomposition, since $QV_t Q' = I$. In this way, candidate draws for the impulse vector are obtained and the impulse responses are calculated, discarding any $V_t'$ where the sign restrictions are violated in all its columns. Repeating such operations until a desired number of draws meet the required sign restrictions allow to calculate the median responses over the accepted draws.

The set of sign restrictions imposed to obtain identification of a government spending shock are as follow: GDP and government spending responses are constrained to be positive for at least four quarters, while the forecast error for only one quarter (see also Table 1). No restrictions are imposed on the average tax variable.

Our procedure accounts for identification uncertainty: for each stable parameter draw of the posterior we find a set of 100 orthonormal matrices that satisfies the sign restrictions. We then compute the corresponding IRFs saving only the median of the 100 identified models. We then repeat this step for each stable draw of the posterior described above for 20,000 parameter draws considering the median of the medians as our estimate of interest.

### 1.4 Multipliers

We estimate the model in normalized levels, similar to Ramey and Zubairy (2017). Thus, there is no need to normalize IRFs in any way, or, to carry out the ex-post conversion that is typically applied in the existing literature (see, e.g., Ramey, 2011b). Our IRFs represent the change in the variable of interest to a surprise change in government spending.

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12 As in Sá et al. (2014); Cogley and Sargent (2005) we avoid the possibility to have explosive IRFs by discarding the explosive draws from the unrestricted posterior.

13 This approach was developed by Canova and De Nicolò (2002), Faust and Rogers (2003), Uhlig (2005). As in Sá et al. (2014) we use the algorithm developed by Rubio-Ramírez et al. (2010).

14 Identification based on sign restrictions is in principle less sensitive to the estimation of the covariance matrix than identification based on short-run restrictions. However, we start with estimation of the structural model, where the order might influence results. This is why we have estimated the model with alternative orderings and do not find significant changes.

15 Uncertainty about identification is due to the fact that we have limited information about the true structural shock. For further details see Sá et al. (2014) and Cogley and Sargent (2005).

16 We compute 20,000 stable draws, discarding the first 10,000 as burn-in draws.
We compute three types of multipliers denoted by $\mathcal{M}_i \in \{1, 2, 3\}$. $\mathcal{M}_1$ is based on Ramey (2011b), who makes a discrete approximation of the integral of the median IRFs over time horizon $h = 0, 1, ..., H$ given by

$$\mathcal{M}_1 = \frac{\sum_{h=0}^{H} d\text{GDP}(h)}{\sum_{h=0}^{H} d\text{G}(h)}$$  \hspace{1cm} (6)$$

Multipliers 2 and 3 are computed using numerical integration, through the use of the Trapezoidal and Simpson’s rule, respectively. The goal of these two computations is to give more accurate approximations of the integrals in

$$\mathcal{M}_{2,3} = \frac{\int_0^{H} (d\text{GDP}(h))dh}{\int_0^{H} (d\text{G}(h))dh}$$  \hspace{1cm} (7)$$

Following Auerbach and Gorodnichenko (2012), we choose $H = 20$.

2. Main Results

In this section we present the macroeconomic effects of a one unit government spending forecast shock obtained for our baseline specification. Figures 2 and 3 show the IRFs of endogenous variables for the low and high interest rate state respectively.

First, observe that IRFs for government spending and GDP in both states are persistently different from zero, except for very high interest rates. Moreover, the median IRF of the average tax rate is mostly insignificant in the low interest rate state. Thus, we have identified a predominantly deficit-financed spending increase in both states.

The behavior of government spending is similar among states. Government spending peaks on impact and is persistently different from zero throughout the time horizon.

Thus, what are the effects on GDP? In sum, the IRFs for GDP qualitatively resemble the behavior of government spending in their respective state. GDP peaks on impact and has a persistently positive response in the subsequent quarters. However, in the high interest rate state, the median IRF becomes insignificant at an earlier point in time. Taking the behavior of GDP and government spending together, the IRFs suggest that when the interest rate is at the ZLB, a comparable exogenous increase in government spending is more effective in stimulating GDP.

The implied multipliers are consistent with our observations, see Table 4. Multipliers, depending on the definition, are in the range of 3.42 and 3.66 in the low interest rate state and around 1.54 and 2.56 in the high interest rate state. Thus, the multipliers also suggest that government spending increases are more effective in the low interest rate state. Moreover, notice that the multipliers for both states are large compared to the VAR literature in general (see, e.g., Ramey, 2011a) and compared to the findings of Ramey and Zubairy (2017) who report multipliers of at most 1.5 at the ZLB and multipliers below unity away from the ZLB.

In sum, our findings cannot be reconciled with theories that suggest that the government spending multiplier at the ZLB is 1 or below, and lower than in the high interest rate state (see, e.g., Braun et al., 2013; Mertens and Ravn, 2014; Aruoba et al., 2017). In addition, our findings, especially for the high interest rate state, contradict with standard Real Business Cycle models (see, e.g., Baxter and King, 1993) that predict a negative wealth effect and lower multipliers due to crowding out of consumption.\textsuperscript{17}

In contrast, our results can be reconciled with New Keynesian DSGE models that predict government spending multipliers at the ZLB in the range of 3 to 5 (see, e.g. Christiano et al., 2011; Eggertsson, 2010; Woodford, 2011; Davig and Leeper, 2011; Coenen et al., 2012). For instance, in

\textsuperscript{17} An increase in government spending lowers the present value of after-tax income. As a consequence, agents lower consumption and increase labor supply. The latter decreases the real wage and higher employment can raise investment.
models such as Christiano et al. (2011) the negative wealth effect of a government spending stimulus is weakened by assumption. As a consequence, co-movement in consumption and real wages due to counter-cyclical markups is possible. An increase in government spending raises aggregate output, marginal cost, and expected inflation. At the ZLB, the key channel to explain the higher multipliers is related to the real interest rate. As expected inflation increases and the nominal interest rate is zero, the real interest rate must fall. In consequence, private consumption increases, raises aggregate output, marginal cost and expected inflation once more. Thus, the ZLB amplifies the effects of government spending on output. As the output increases require an increase in employment, these models also imply real wage increases.

3. Robustness

In this section we address misspecification concerns regarding our baseline specification and the results presented above. Notice that we maintain the identification approach described in Section 1.3 throughout the robustness analysis.

3.1 Factor-Augmented Interacted VAR Model

In particular, one may argue that our baseline specification is problematic for two reasons that can be addressed by developing a FAIVAR model. First, the choice of variables in \( Y_t \) is subject to discretion. Thus, one may argue that our results are due to the particular choice of variables in \( Y_t \). Second, given the considerations and results in Fragetta and Gasteiger (2014), one may argue that our Interacted VAR model is affected by a generic limited information problem. As a matter of fact, when economic agents make their decisions, they use all available information at the time. In contrast, an econometrician can only take into account a limited set of information, due to the problem related to degrees of freedom.

A FAIVAR model addresses both lines of critique. On the one hand it allows us to take into account the information from a large informational data set and to maintain a small set of variables in \( Y_t \) that is necessary for meaningful identification. Thus, discretion in the specification of \( Y_t \) is limited to a minimum. On the other hand, the FAIVAR model allows us to overcome the generic limited information problem.

We implement a two-step estimation procedure. Following Bernanke et al. (2005), we use the method of principal components to extract and summarize information from a large dataset. The Bai and Ng (2002, 2007) \( IC_{p2} \) criterion suggests to extract four static factors. Thus, we specify the vector of endogenous variables in the FAIVAR model as

\[
Y_t = [F_{t}, G_{t}, GDP_{t}, T_{t}, F_{t}]'.
\]  

where \( F_t \) is the 4 × 1 vector capturing the first four principal components of the informational dataset.

The IRFs in Figures 4 and 5 depict the low and high interest rate state respectively. Overall, IRFs of government spending and average taxes show a qualitatively similar pattern as in the baseline specification. However, IRFs for GDP reveal several differences compared to the baseline specification. First, IRFs for GDP are less persistent. This behavior is particular evident in the low interest rate state. In

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18 Thus, in such models multipliers can be large even without considering the ZLB (see, Galí et al., 2007).

19 We apply the principal components method by using the same informational dataset as used in Fragetta and Gasteiger (2014). Their informational dataset comprises 61 publicly available time series from the Federal Reserve Bank of St. Louis’ FRED R Economic Database. As in their case we transform variables to guarantee stationarity according to Dickey and Fuller (1979) and Kwiatkowski et al. (1992) tests.
the latter case, one can also find hump-shaped IRFs. Nevertheless, the lower persistence should be reflected in lower multipliers.
Consistent with this claim, Table 4 shows that multipliers are now lower in both states. Moreover, decline of multipliers in the FAIVAR compared to the IVAR is of higher magnitude in the low interest rate state. Nevertheless, multipliers in the low interest rate state range from 1.98 to 2.10, while multipliers range from 1.48 to 1.79 in the high interest rate state. Thus, as multipliers in the low interest rate state exceed the ones in the high interest rate state, we conclude that our baseline results are robust with regard to the particular specification of $Y_t$ and the generic limited information problem of (interacted) VARs.

Conclusions
This paper sheds light on the question of whether the government spending multiplier at the ZLB is larger than in normal times. To this end, we implement an Interacted VAR model and use sign restrictions to identify government spending growth forecast shocks. This framework allows us to account for fiscal foresight and to estimate state-dependent government spending multipliers at all percentiles of the nominal interest rate distribution.
In contrast to the existing state-dependent estimates, we find convincing evidence that government spending multipliers are larger in low interest rate states than in high interest rate states. For our sample from 1966 to 2015, the multipliers at the ZLB are in the range of 3.4 to 3.7. The ones away from the ZLB are between 1.5 to 2.7. Our findings are robust to several important misspecification concerns.
Thus, we conclude that the government spending multiplier at the ZLB is larger than in normal times and within the range of 3 to 5 as predicted by recent New Keynesian DSGE models.
References


Appendix 1 Data

**General Information.** Table 2 contains an overview on the data that we use. If appropriate, nominal variables are transformed into real variables by dividing by the GDP implicit price deflator. Moreover, real variables in levels, if appropriate, are normalized by dividing by real potential GDP. The forecast error that we use is the forecast error for the annualized growth rate of real government spending. We normalize this variable by subtracting the annualized growth rate of real potential GDP.

**Forecast Error.** Our measure of the forecast error, $FE_t$, builds on the annualized growth rate of real government purchases forecast for time $t$ at time $t-1$, i.e.,

$$\Delta G_{t|t-1}^F \equiv \left[ \left( \frac{G_{t|t-1}}{G_{t-1|t-1}} \right)^4 - 1 \right] \times 100,$$

(A.1)

The data source is the Mean Responses of Real Federal Government Consumption Expenditures and Gross Investment (RFEDGOV) and Real State and Local Government Consumption Expenditures and Gross Investment (RSLGOV). $G_{t|t-1}$ is the sum of RFEDGOV3 and RSLGOV3, $G_{t-1|t-1}$ is the sum of RFEDGOV2 and RSLGOV2.

As our objective is to compute a series of surprise increases in government spending, we need to control for real-time data. The forecast error for the growth rate of government spending is defined as

$$FE_t \equiv \left[ \left( \frac{G_{t|t}^{s.t.}}{G_{t-1|t}^{s.t.}} \right)^4 - 1 \right] \times 100 - \Delta G_{t|t-1}^F$$

(A.2)

Thus, for this purpose, we have downloaded first release data on real government consumption expenditures and gross investment: state and local (RGSL) from this website and real government consumption and gross investment: federal (RGF) from this website. All in quarterly vintages (Billions of real dollars, seasonally adjusted). $G_{t|t}^{s.t.}$ is the sum of RGSL and RGF.

Notice that the SPF data is only available from 1981Q4. Thus, for earlier periods, as in Auerbach and Gorodnichenko (2012), we take advantage of the fact that SPF is also quite similar to Greenbook forecasts prepared for FOMC meetings. Thus, we splice data from SPF and Greenbook forecasts and obtain a series which goes from 1966Q4 to 2015Q4.
Appendix 2: Explanatory Power of the Forecast Error

Following Ramey (2011b, pp.25-29) we examine the explanatory power of $FE_t$. In particular, we run regressions such as

$$
\Delta G_t = \beta_0 FE_t + \sum_{k=1}^{L} \beta_k FE_{t-k} + \varepsilon_t
$$

$$
\Delta G_t \equiv \left[ \left( \frac{G_t}{G_{t-1}} \right)^4 - 1 \right] \times 100 \quad \text{(B.1)}
$$

Such a regression can shed light on the question of whether $FE_t$ (or lags of it) can explain part of the variation of the growth in $G_t$. A high F-statistic is an indicator that this is the case and that $FE_t$ can be considered a relevant instrument to control for fiscal foresight. The results in the second column of Table 3 suggest that $FE_t$ is a relevant instrument and that it cannot be considered a weak instrument as the F-statistics are way above the rule-of-thumb critical value of 10.

Notice that even with two lags, $L = 1$, $FE_t$ has considerable predictive power. This is surprising as, by construction, one would expect that it has only predictive power for $L = 0$. The reason for the latter is that $FE_t$ represents a measure for the unpredictable component of $\Delta G_t$. Therefore, our results for $L > 0$ imply that the unpredictable components in $\Delta G_t$ have some persistence.

The third column in Table 3 reports the marginal F-statistic for a regression of the growth rate of $G_t$ on the explanatory variables used in the baseline specification. However, $FE_t$ is excluded, i.e.,

$$
\Delta G_t = \sum_{k=1}^{L} \beta_{k,G} G_{t-k} + \sum_{k=1}^{L} \beta_{k,GDP} GDP_{t-k} + \sum_{k=1}^{L} \beta_{k,T} T_{t-k} + \varepsilon_t. \quad \text{(B.2)}
$$

Table 3 reports low marginal F-statistics and values for R-squared, which suggests that $FE_t$ is a relevant instrument.
Appendix 3 Figures

Figure 1  Monetary and fiscal policy, 1960Q1 to 2015Q4. The shaded areas indicate recessions according to NBER
Figure 2  
IRFs to a one unit government spending growth forecast shock for the baseline specification with \( X_t = st_{t-1} \) in the low interest rate state. The blue solid lines represent the median of the median distribution of IRFs for each parameter draw, and the red dotted lines report the 16th and 84th of the set of accepted impulse-response functions for all parameter draws.
Figure 3  IRFs to a one unit government spending growth forecast shock for the baseline specification with $X_t = st_{t-1}$ in the high interest rate state. The blue solid lines represent the median of the median distribution of IRFs for each parameter draw, and the red dotted lines report the 16th and 84th of the set of accepted impulse-response functions for all parameter draws.
Figure 4: IRFs to a one unit government spending growth forecast shock for the specification with $X_t = s r_{t-1}$ and $I_t$ in the low interest rate state. The blue solid lines represent the median of the median distribution of IRFs for each parameter draw, and the red dotted lines report the 16th and 84th of the set of accepted impulse-response functions for all parameter draws.
Figure 5  IRFs to a one unit government spending growth forecast shock for the specification with $X_t = r_{t-1}$ and $F_t$ in the high interest rate state. The blue solid lines represent the median of the median distribution of IRFs for each parameter draw, and the red dotted lines report the 16th and 84th of the set of accepted impulse-response functions for all parameter draws.
### Table 1  Sign Restrictions for Identifying the Government Spending Shock

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<td>Forecast Error of the Annualized Growth Rate of Real Government Purchases</td>
<td>All the above variables are used for the computation, see Appendix A.</td>
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<td>Normalized</td>
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<td>Nominal Federal Government Current Receipts</td>
<td>US Bureau of Economic Analysis</td>
<td>FGRECPT</td>
<td>Real, Average w.r.t. GDP</td>
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<td>Real Gross Domestic Product</td>
<td>US Bureau of Economic Analysis</td>
<td>GDPC1</td>
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<td>Gross Domestic Product: Implicit Price Deflator</td>
<td>US Bureau of Economic Analysis</td>
<td>GDPDEF</td>
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<td>Shadow Federal Funds Rate</td>
<td>Wu and Xia (2016)</td>
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<td>Real Potential Gross Domestic Product</td>
<td>US Congressional Budget Office</td>
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<td>Lag Length</td>
<td>R-squared</td>
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<td>0.0473</td>
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Table 3  Explanatory Power of $F_{E_t}$.\(^a\)

Note  \(^a\) For each lag length L the first line reports results for regression (B.1). The second line reports results for regression (B.2). In the case of L = 0, (B.2) uses contemporaneous values.

### Baseline: $X_t = sr_{t-1}$

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<th>1(^{st})</th>
<th>5(^{th})</th>
<th>13(^{th})</th>
<th>25(^{th})</th>
<th>50(^{th})</th>
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<td>$m_2$</td>
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<td>$m_3$</td>
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<td>2.53</td>
<td>2.07</td>
<td>1.64</td>
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</table>

### Robustness: $X_t = sr_{t-1}$ and $F_t$

<table>
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<td>1.76</td>
<td>1.60</td>
<td>1.48</td>
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<tr>
<td>$m_3$</td>
<td>1.98</td>
<td>2.08</td>
<td>2.04</td>
<td>1.76</td>
<td>1.60</td>
<td>1.48</td>
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</tbody>
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Table 4  Multipliers identified with $F_{E_t}$.\(^a\)

Note  \(^a\) Multipliers $m_i \in \{1, 2, 3\}$ are calculated as outlined in Section 1.4.
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