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

MARIO DI SERIO,† MATTEO FRAGETTA†,‡,§ and EMANUEL GASTEIGER‡,¶

†Department of Economics and Statistics, Universitá degli Studi di Salerno, Via Ponte Don Melillo, 84084 Fisciano (SA), Italy (e-mail: mdiserio@unisa.it)
‡Instituto Universitário de Lisboa (ISCTE-IUL), Business Research Unit (BRU-IUL), Av.ª das Forças Armadas 1649-026,Lisboa, Portugal
§Centro di Economia del Lavoro e di Politica Economica, CELPE, Università degli Studi di Salerno, Italy
¶Institute of Statistics and Mathematical Methods in Economics, TU Wien, Wiedner Hauptstr. 8-10 1040, Wien, Austria (e-mail: emanuel.gasteiger@tuwien.ac.at)

Abstract

We estimate state-dependent government spending multipliers for the United States. We use a factor-augmented interacted vector autoregression (FAIVAR) model. This allows us to capture the time-varying monetary policy characteristics including the recent zero interest rate lower bound (ZLB) state, to account for the state of the business cycle and to address the limited information problem typically inherent in VARs. 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 in a recession range from 3.56 to 3.79 at the ZLB. Away from the ZLB, multipliers in recessions range from 2.31 to 3.05. Several robustness analyses confirm that multipliers are higher, when the interest rate is lower and that multipliers in recessions exceed multipliers in expansions. Our results are consistent with theories that predict larger multipliers at the ZLB.

I. 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

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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 that is 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.[1] 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 increase in government expenditures at the beginning of this period. This policy can be rationalized by arguing that in such an extraordinary situation as the ZLB, an increase in government spending is more effective than in normal times.[2]

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, Eichenbaum and Rebelo (2011), Eggertsson (2010), Woodford (2011), Davig and Leeper (2011) or Coenen *et al.* (2012). These models predict a government spending multiplier at the ZLB much larger than one. 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 Boneva, Braun and Waki (2016), Mertens and Ravn (2014), Aruoba, Cuba-Borda and Schorfheide (2018).

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.[3] 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 (2018) is the single paper for the United States in this literature according to our knowledge.[4] Ramey and Zubairy (2018) 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.[5] Moreover, there is a related, but distinct empirical literature quantifying state-dependent fiscal multipliers in recessions and expansions based on regime-switching VAR (see, e.g. Auerbach and Gorodnichenko, 2012, 2013), local pro-

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[1]For instance, the Fed announced three rounds of quantitative easing: in November 2008, in November 2010 and in September 2012.

[2]Consistent with the idea that fiscal multipliers are different at the ZLB, several studies find changes in macroeconomic performance at the ZLB (see, e.g. Liu *et al.*, 2019).

[3]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'.

[4]Crafts and Mills (2013) and Ramey (2011b) provide evidence for ZLB episodes suggesting multipliers below unity.

[5]Miyamoto *et al.* (2018) build on the methods used in Ramey and Zubairy (2018) and provide evidence for Japan. They find that the impact multiplier is around 1.5 at the ZLB and much larger than away from the ZLB. More recently, Amendola *et al.* (2019) build on the ideas and methods in our paper and estimate a panel version of our model for the Euro Area. Their findings are consistent with our findings.
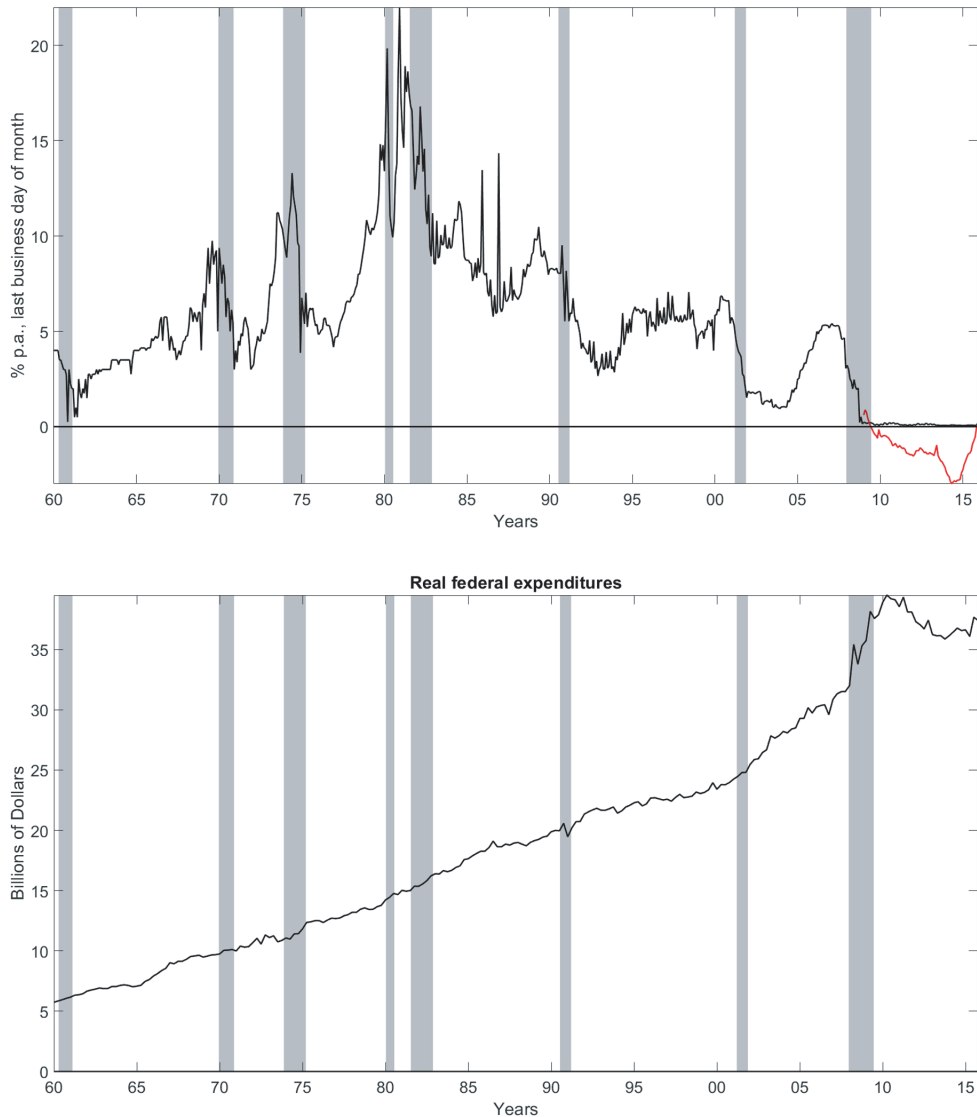

Figure 1. Monetary and fiscal policy, 1960Q1 to 2015Q4. The shaded areas indicate recessions according to NBER

jection (see, e.g. Ghassibe and Zanetti, 2020) or structural vector moving-average models (see, e.g. Barnichon, Debortoli and Matthes, 2019). However, as Figure 1 illustrates, recessions and episodes where the Federal Funds Rate is at zero or below 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 factor-augmented interacted vector autoregressive model with exogenous variables (FAIVAR-X) building on the interacted vector autore-

gression (IVAR) model in Towbin and Weber (2013) and Sá, Towbin and Wieladek (2014). We augment the IVAR model with factors from a large informational data set similar to Bernanke, Boivin and Eliasz (2005) and Fragetta and Gasteiger (2014). Incorporating exogenous variables facilitates the identification of the government spending shock.

The key advantage of building on the IVAR methodology is the interaction term, which 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 the median and low levels of the interest rate distribution, with no need to restrict the sample. Likewise, the IVAR allows us to estimate multipliers depending on the state of the business cycle. Thus, we can compute the multiplier for the case where low interest rate state coincides with a recession. This is especially important in the light of the bulk of the policy debate, which focuses on fiscal stimulus at the ZLB in recessions.

In addition, using the IVAR methodology has further benefits. For instance, one benefit compared to regime-switching approaches is that the IVAR 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 may be subject to discretion as the threshold is frequently chosen by the researcher (see, e.g. Ramey and Zubairy, 2018). In contrast, the IVAR, in principle, allows to distinguish between as many states of economy as there are observations for the variable that is used in the interaction term, for example the interest rate. A second related benefit is that the IVAR uses all the information available for the full sample, while a threshold model uses the information of each state under consideration separately. A third benefit of IVARs is that the interaction term can capture abrupt policy changes next to smooth policy changes. This is particularly important as Caggiano, Castelnuovo and Pellegrino (2017, p.11) emphasize that the change in monetary policy in times of crises is frequently abrupt and not smooth.

The second key advantage of our FAIVAR framework is that it addresses the generic limited information problem inherent in VARs. On the one side, introducing more and more variables to a VAR adds more information. However, adding additional variables to a VAR implies a loss of degrees of freedom. We handle this trade-off by augmenting the IVAR with factors, estimated as the principal components from a large informational data set.

The third 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 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.[6] Moreover, treating growth in the forecast of government spending as an exogenous variable in our FAIVAR-X framework is our way of addressing the concerns related to fiscal foresight in Leeper, Walker and Yang (2013). Our approach ensures that the shocks to government spending identified by sign restrictions are orthogonal to the information set of economic decision makers. All information regarding government spending

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[6]The sign restrictions approach is developed in Canova and De Nicolò (2002), and Uhlig (2005). Mountford and Uhlig (2009) apply it to fiscal policy.

that could have been anticipated is captured by the growth in the forecast of government spending.

For our sample from 1966Q4 to 2015Q4, the baseline FAIVAR-X specification involves government spending, GDP, the average tax rate and the 10Y treasury bond yield as endogenous variables. During recessions, the 5-year cumulative government spending multipliers at the ZLB range between 3.56 and 3.79. When monetary policy is not constrained by the ZLB, government spending multipliers during recessions are between 2.31 and 3.05.

In order to assess the robustness of these estimates, we carry out three types of analysis. First, we isolate the role played by the interaction terms, the factors and the exogenous variables in our FAIVAR-X in generating our main findings. To this end, we estimate an IVAR, FAIVAR and a IVAR-X model. Second, we consider a FAIVAR specification, where we incorporate the forecasts errors based on growth in the forecast of government spending in the vector of endogenous variables. This is an alternative way of addressing fiscal foresight. The forecast error series captures the surprise component in a broad measure of government spending and, as we show, is a relevant and strong instrument for our post-WWII sample. An alternative would be to consider the defence news series used in Ramey and Zubairy (2018). 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. Third, we identify the government spending shock via timing instead of sign restrictions to shed light on the importance of sign restrictions for our main findings.

While the range of multipliers is considerably increased over the various specifications in the robustness exercises, the bottom line result is the same: multipliers are higher when interest rates are lower. Moreover, multipliers in recessions are larger than in expansions. Thus, our results are qualitatively and quantitatively consistent with the claim that increases in government spending are more effective at the ZLB and during recessions.

The paper proceeds as follows: section II outlines the FAIVAR-X model, our baseline specification and data, our inference and identification approach and how we calculate the multipliers; section III discusses the main results; section IV addresses robustness concerns; section V concludes.

II. Methodology

Empirical model

We use a FAIVAR-X based on Towbin and Weber (2013) and Sá *et al.* (2014). The recursive form is given by

$$B_t Y_t = \kappa + \sum_{k=1}^{L} \Gamma_k Y_{t-k} + \nu Z_{t|t-1} + \sum_{m=1}^{N} \kappa_m^1 X_{t,m} + \sum_{k=1}^{L} \sum_{m=1}^{N} \Gamma_{k,m}^1 X_{t,m} Y_{t-k} + \varepsilon_t, \qquad (1)$$

where $t = 1, \ldots, T$ denotes time, $k = 1, \ldots, L$ denotes the lag length and $m = 1, \ldots, N$ denotes the number of interaction terms. Y_t is a $q \times 1$ vector which contains explanatory variables. κ is the intercept, Γ_k is a $q \times q$ matrix of autoregressive coefficients and $\varepsilon_t \sim \mathrm{N}(0, \Sigma)$ is the vector of residuals. $Z_{t|t-1}$ is an exogenous variable accounting for fiscal foresight.

Moreover, X_t is a $m \times q$ matrix, which denotes the interaction terms. These interaction terms can influence both the dynamic relationship between endogenous variables and their level, trough $\Gamma^1_{k,m}$ and κ^1_m 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 qth-ordered variable on the wth-ordered variable. It is constructed as follows:

$$
B_t(w,q) = \begin{cases} 0 & \text{for } q > w \\ 1 & \text{for } q = w \\ B(w,q) + \sum_{m=1}^{N} B^1_m(w,q) X_{t,m} & \text{for } q < w, \end{cases}
$$

where $B^1_m(w,q)$ are regression coefficients capturing the relation with the contemporaneous marginal effects of a change in the interaction terms. The recursive form of the matrix B_t implies that the covariance matrix of the residuals, Σ, is diagonal (for more details on the interacted VAR framework, see for example Sá *et al.*, 2014).

Baseline specification

Our data set consists of US quarterly data and goes from 1966Q4 to 2015Q4.[7] In our baseline specification, (1) the vector of endogenous variables $Y_t = [G_t, \text{GDP}_t, T_t, i^{10y}_t, F_t]'$ includes mostly 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. GDP_t stands for real gross domestic product. Moreover, T_t denotes the net (of transfers) average tax rate computed as nominal current tax receipts divided by nominal GDP. Variables G_t and GDP_t are considered in levels and have been normalized with an estimate of real potential GDP.[8]

We also include the 10Y treasury bond yield, denoted by i^{10y}_t. We use this variable to account for potential effects of the shock to government spending on the central bank interest rate. Ideally, we would use the (shadow) Federal Funds Rate for this purpose. However, as we discuss below, we will use the latter as an interaction term, which prevents us from using it as an endogenous variable at the same time. Nevertheless, according to the *expectations hypothesis of the term structure*, the long-term bond yield should be a valid measure of controlling for monetary policy responses. Movements in expected short-term interest rates should transmit into movements in long-term interest rates. For instance, Gürkaynak, Sack and Swanson (2005), Roush (2007) or Favero and Giavazzi (2008) provide empirical evidence that supports this claim. Roush (2007) also finds that the term premium does not play a significant role in the United States.

Next, we augment Y_t with the 4×1 vector F_t, which captures the first four principal components of an informational data set.[9] Our motivation is twofold. First, the choice of

[7]The choice of this time period is motivated by the availability of the Greenbook and Survey of Professional Forecasters (SPF) real government spending forecasts, which we use in our identification strategy detailed further below and the end of quantitative easing in the United States.

[8]Appendix A and Table 2 contain detailed information on the data. We normalize G_t and GDP_t to avoid biases in the government spending multiplier calculation as discussed in detail below. Normalizing the net average tax rate by potential GDP does not affect results.

[9]We apply the principal components method by using the same informational data set as used in Fragetta and Gasteiger (2014). Their informational data set comprises 61 publicly available time series from the Federal Reserve

variables in Y_t is subject to discretion. Thus, one may argue that any results obtained are due to a particular choice of variables in Y_t. Second, given the considerations and results in Fragetta and Gasteiger (2014), an interacted VAR (IVAR) model is potentially affected by a generic limited information problem. Applied econometricians have to preserve degrees of freedom and have to specify parsimonious models. Thus, they can only specify a limited number of variables. However, when economic agents make their decisions, they may use all available information at the time. This misalignment in information sets may render the government spending shock that we identify below as non-fundamental and also bias our estimates (see, Lippi and Reichlin, 1994). This can be seen as a *generic* limited information problem of VARs. By augmenting Y_t with F_t, both of these concerns can be addressed. 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 factor-augmented model allows us to overcome the generic limited information problem.[10]

Moreover, $Z_{t|t-1}$ in equation (1) is the growth in the forecast of real government spending used in Auerbach and Gorodnichenko (2012) and based on the Greenbook and SPF forecasts.[11] In this way, we address fiscal foresight, which is a *specific* limited information problem that causes a misalignment of information sets. As a matter of fact, agents can forecast the future fiscal stance and change their behaviour even before its implementation. Thus, by adding $Z_{t|t-1}$ to our model, we account for this limited information problem (for a detailed discussion of fiscal foresight see Leeper *et al.*, 2013).

Finally, we use the US Shadow Federal Funds Rate developed by Wu and Xia (2016), that is $X_t = sr_{t-1}$ as interaction term. This 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 sr_{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 *low* interest rate state, as the 13th percentile coincides with a value of the interest rate equal to 0.25. Values below this value are conventionally accepted by the literature as the lower bound for monetary policy in using the Federal Funds Rate as instrument (see, e.g. Ramey and Zubairy, 2018). Results for the 25th percentile and above are associated with the *high* interest rate 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.

A potential concern to quantifying the fiscal multiplier with such an interaction term is that the low interest rate state may coincide with recessionary episodes in the data, for example the Great Recession. In consequence, the results may be interpreted as

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potentially coming from the recession effect instead of being attributed to the low interest rate state. Indeed, the findings by Auerbach and Gorodnichenko (2012) indicate that the fiscal multiplier is higher in recessions. Therefore, in order to separate the effects attributable to the interest rate state on the one hand and the business cycle state on the other, we add a second interaction term indicating the state of the business cycle: the official NBER recession indicator.

Regarding the use of the US Shadow Federal Funds some further remarks are in order. The rate is considered 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 the ZLB Wu and Xia (2016) use a Gaussian Affine Term Structure Model 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 these policies. Moreover, using the US Shadow Federal Funds Rate has the big advantage that our low interest rate state also includes negative short-term policy rates that were targeted by central banks during the great recession (see, e.g. Swanson, 2018).

Notice also that we use the first lag of the shadow rate to address potential endogeneity concerns. Specifying sr_{t-1} in X_t implies that the monetary policy instrument is not endogenous to Y_t. If we were to specify sr_t in X_t, reversed causality could be a problem for part of the information set contained in the endogenous variables. Moreover the four factors in F_t contain information on the monetary policy stance, which does not need to be identified in our case. Therefore the lagged shadow rate does not create particular concerns. Finally, based on the Hannan–Quinn information criterion, we choose a lag length of order 2.

Inference and identification

As in Uhlig (2005) and Sá *et al.* (2014), to capture parameter uncertainty, we use Bayesian estimation by setting an uninformative normal-Wishart prior. We start with the estimation of the structural recursive model described in equation (1). Since the covariance matrix Σ is diagonal by construction we can proceed by estimating the model equation by equation. We draw the recursive-form parameters from the posterior.[12] We evaluate them at a prespecified value of the interaction term and compute reduced form parameters by inverting the matrix B_t.

Given the reduced form, we use a sign restriction strategy to identify an unexpected government spending shock. More specifically, we follow the same procedure of Sá *et al.* (2014), by using the algorithm developed by Rubio-Ramírez, Waggoner and Zha (2010). Defining $V_{x,d}$ as the Cholesky decomposition of the reduced form variance–covariance matrix $\Sigma_{x,d}$, we draw an orthonormal matrix Q such that $Q'Q = I$, from which follows $B_d = V_{x,d}Q$ and $\Sigma_{x,d} = B'_d B_d = V'_{x,d}Q'QV_{x,d}$, where d indicates a stable draw from the posterior distributions. To achieve identification, the impulse responses implied by B_d have to satisfy

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

TABLE 1

Sign restrictions for identifying the government spending shock

Baseline			Robustness		
Variable	Sign	Periods	Variable	Sign	Periods
			FE_t	+	1
G_t	+	4	G_t	+	4
GDP_t	+	4	GDP_t	+	4
T_t	*		T_t	*	

the following restrictions: a government spending shock should raise GDP_t and G_t for at least four quarters, see Table 1.

For every 100 draws of the Q matrix which meet our sign restrictions, we save its median value.[13] We make 20.000 draws from the posterior distribution and use the median over the 10.000 medians obtained as our central estimate of interest.[14] We account for parameter uncertainty by saving the 16th and 84th percentile of the distribution of the median as error or confidence bands.[15,16]

Multipliers

We estimate the model with G_t and GDP_t in levels and normalized with an estimate of real potential GDP. This is particularly important as our main objective is to provide estimates for the government spending multiplier. Ramey and Zubairy (2018) show that the usual approach of using log levels requires an *ex post* conversion to dollar equivalents of the estimated elasticities can produce serious bias. The problem is even more acute in nonlinear models and in particular in our model, where it is possible to calculate several multipliers with the exception of the quarters which have the same interest rate, since the *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, there is no need to carry out the *ex post* conversion that is typically applied in the existing literature (see, e.g. Ramey, 2011b).

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[13] Note that the algorithm excludes the possibility to have a multiple shock problem (see, Fry and Pagan, 2011). Thus, it discards the draws that give rise to more than one identical shock.

[14] Note that we consider the first 10.000 parameter draws as burn-in draws.

[15] Parameter uncertainty is due to the limited amount of data. As in Sá *et al.* (2014), we use confidence bands to address parameter uncertainty, because we are mainly interested on the differences between the low and high interest state. Notice also that in our approach there is another source of uncertainty, which is identification uncertainty. It reflects the lack of information we have about the true properties of the structural shock and is intrinsic in our sign restriction approach. For further details, see Sá *et al.* (2014).

[16] The method to derive IRFs used herein has indeed been criticized by Fry and Pagan (2011), who claim that considering the median response as the central estimate of interest may be inaccurate since the median contains information from different identified models, following from the different accepted draws of the rotation matrix (identification uncertainty). For robustness purpose, we also consider and compute IRFs using the median target approach described in Fry and Pagan (2011). The results based on this method are very similar to ones reported in the paper and are available from the authors upon request.

Thus, government spending multipliers can be computed directly.[17] Our IRFs represent the change in the variable of interest to a surprise change in government spending. Therefore, throughout the paper, we compute and report cumulative multipliers similar to Ramey (2011b), who makes a discrete approximation of the integral of the median IRFs over time horizon $h = 0, 1, \ldots, H$ given by

$$\mathcal{M} = \frac{\sum_{h=0}^{H} y_h}{\sum_{h=0}^{H} g_h},$$

where y_h and g_h denote the value of the IRF of GDP and government spending at horizon h.[18] As is common in the related literature, we compute IRFs for a horizon of up to $H = 20$ quarters.

III. Main results

In this section, we present the macroeconomic effects of a one standard deviation government spending shock obtained for our baseline specification. For illustrative purposes, we report IRFs for the 5th and 50th percentile. The percentiles are chosen to be representative for the low and high interest rate state respectively. Moreover, given that a significant share of the policy debate is centred around fiscal stimulus in recessions, the discussion in this paper focuses on a comparison of the 5th and 50th percentile in the recession state. Whenever appropriate, we also highlight differences in the expansion state.[19] To make the IRFs in different states comparable, we scale the IRFs such that the IRFs of government spending are equal among states on impact.

The left-hand side panel in Figure 2 shows that the behaviour of government spending is similar among interest rate states in recessions. Government spending peaks shortly after impact and is persistently different from zero for the first 9–12 quarters of the considered time horizon.

What are the effects on GDP in the low and high interest rate state in recessions? GDP has a hump-shaped IRF, peaks also shortly after impact, and, has a persistently positive IRF in both interest rate states. However, the IRF is larger in the low interest rate state for the first 10 quarters and similar to the IRF in the high interest rate state thereafter.

The net average tax rate closely follows the pattern of real GDP in both interest rate states, which implies, by construction, that nominal current tax receipts increase by more than nominal GDP. This behaviour suggests that we are, to some extent, identifying a tax-financed spending expansion in both the low and high interest rate state.[20]

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[17] For more details on the bias caused by the *ex post* conversion of the elasticities, see Ramey and Zubairy (2018) and Gordon and Krenn (2010).

[18] We obtain similar results if we compute the multipliers by using numerical integration, through the use of the *Trapezoidal* and *Simpson's* rule respectively. The goal of these two rules is to give more accurate approximations of the integrals in $\mathcal{M} = (\int_0^H y(h)dh)/(\int_0^H g(h)dh)$.

[19] The full set of IRFs is available on request.

[20] An alternative would be to impose further restrictions on the net average tax rate in order to identify entirely balanced budget or tax-financed spending expansions (see, e.g. Mountford and Uhlig, 2009).
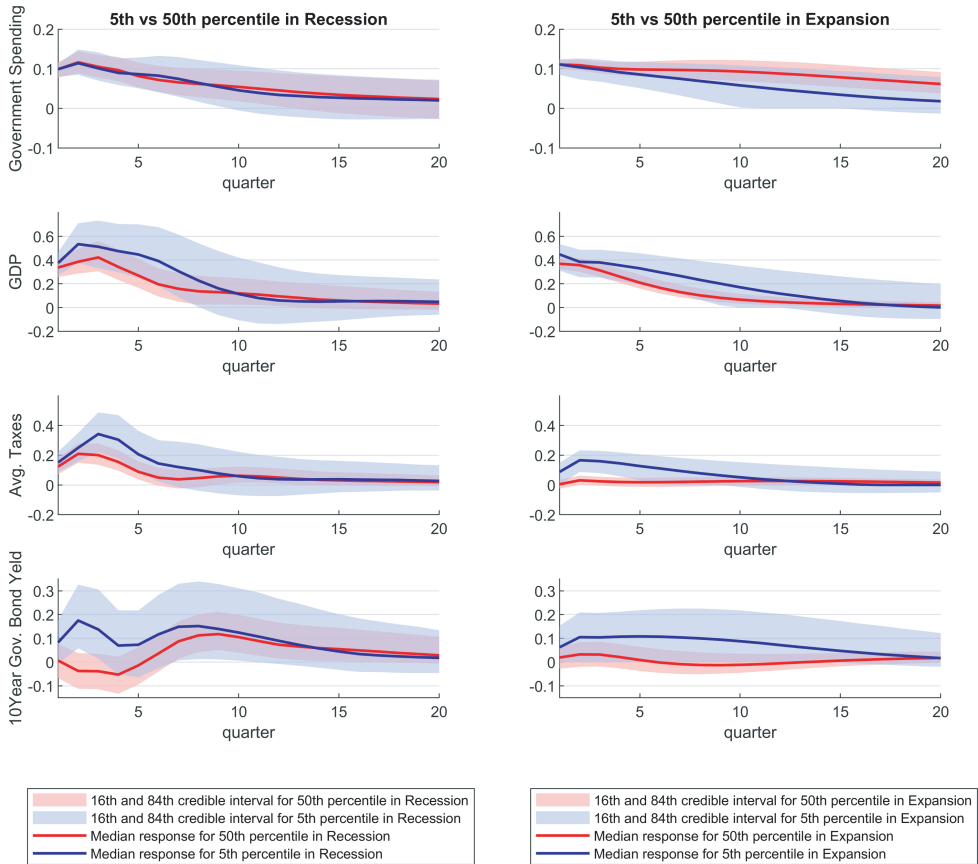
Figure 2. IRFs to a one standard deviation government spending shock for the baseline specification, FAIVAR-X. The blue solid lines represent the median of the median distribution of IRFs for each parameter draw, and the red dotted lines report the 16th and 84th of the set of accepted impulse-response functions for all parameter draws

The IRF of the 10Y treasury bond rate at the 5th percentile increases on impact and the quarters thereafter. In contrast, at the 50th percentile, the 10Y treasury bond rate shows only a modest increase after 7 quarter.

Taking the behaviour of GDP and government spending together, the IRFs suggest that, when the interest rate is at the ZLB and the economy is in a recession, a comparable exogenous increase in government spending is more effective in stimulating GDP. Next, the right-hand side panels show the IRFs for the 5th and 50th percentile of the interest rate distribution during expansion. Qualitatively, these IRFs show by and large a similar picture, except for the missing hump-shape in government spending and GDP. Moreover, the figures suggest quantitative differences, which can be assessed by computing multipliers.

The implied multipliers are consistent with the above observations. Panel 3a-3e in Figure 3 show multipliers for the different percentiles of the interest rate distribution for several time horizons. For instance, the 5-year multipliers in Panel 3e are in the range of 3.79–2.66 in the low interest rate state and around 3.05 and 1.01 in the high interest rate state when we ignore the business cycle. Moreover, two key findings stand out: first, multipliers
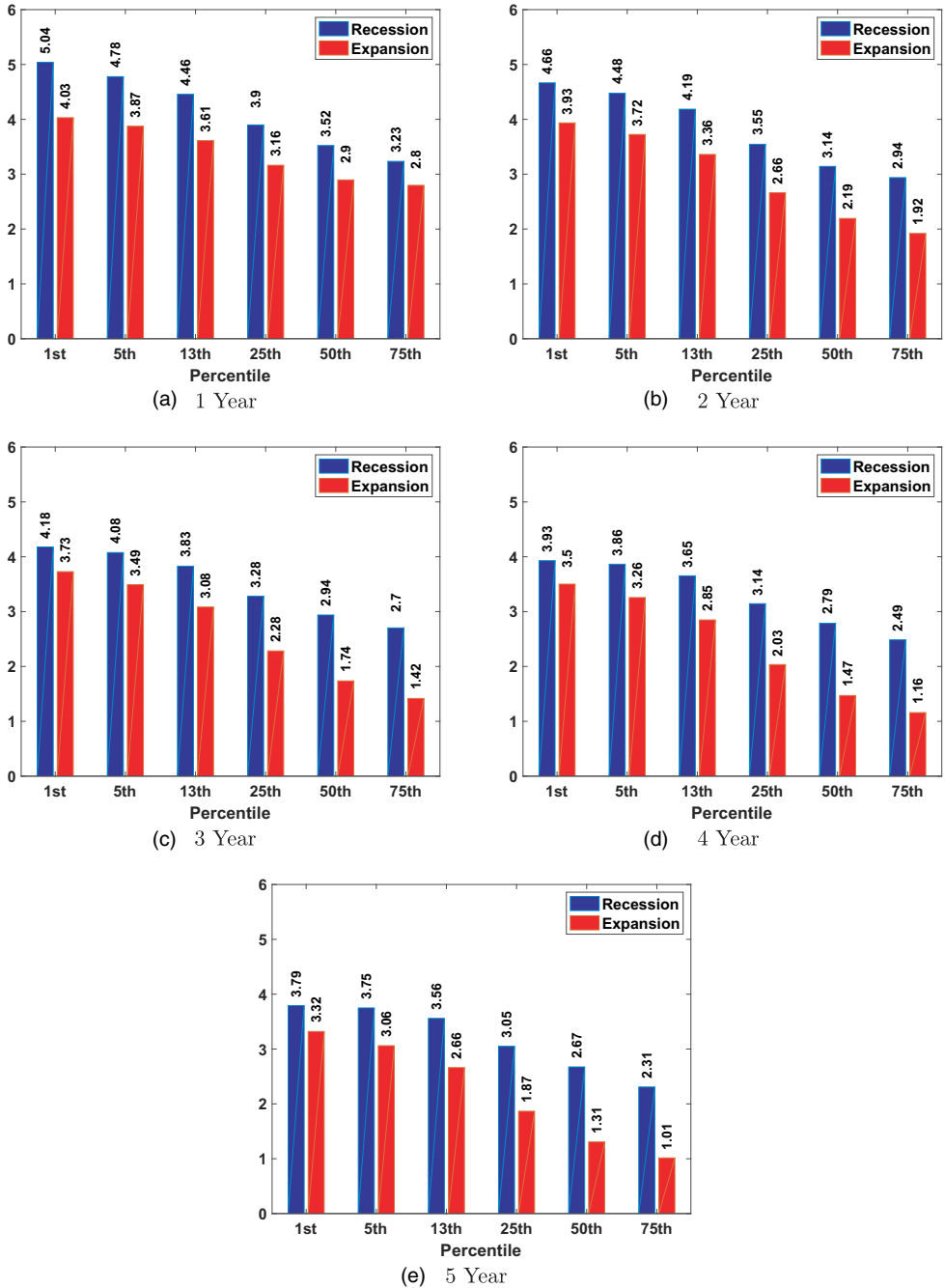
Figure 3. Cumulative government spending multipliers at several time horizons and probabilities for differences in multipliers for the FAIVAR-X model: (a) 1 year; (b) 2 year; (c) 3 year; (d) 4 year; (e) 5 year; (f) probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile; (g) probability that the cumulative multiplier in a recession is higher than in an expansion

(f) Probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile.

(g) Probability that the cumulative multiplier in a recession is higher than in an expansion.
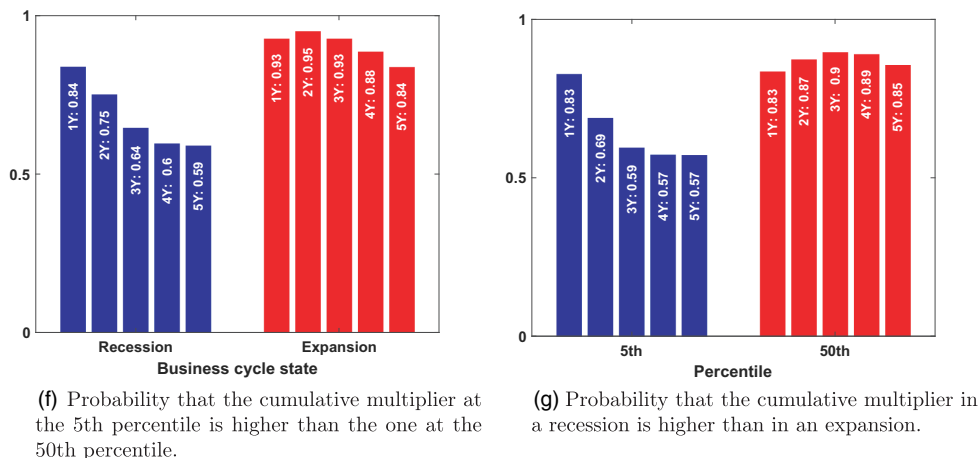
Figure 3.   (Continued)

monotonically decline with the increasing interest rate percentile in both recession and expansion; second, multipliers in a recession are higher than in an expansion, independent of the interest rate percentile.

The first key finding, similar to the IRFs above, also suggests that government spending increases are more effective in the low interest rate state. In order to further examine this point, we compute the probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile at various time horizons.[21] To this end, we construct distributions of the difference between multipliers conditional on specific shadow rate percentiles. The difference between multipliers is computed for each of the 10.000 parameter draws from the posterior distribution. The results are presented in Panel 3f and indicate that the probability that the multipliers are higher at the 5th percentile is in the range of 59–95%, depending on the time horizon and whether the economy is in a recession or expansion. Moreover, notice that the multipliers for both states are relatively large compared to the VAR literature in general (see, e.g. Ramey, 2011a) and compared to the findings of Ramey and Zubairy (2018), who report multipliers of at most 1.5 at the ZLB and multipliers below unity away from the ZLB.

The second key finding can also be investigated to a further extent by computing the probability that the cumulative multiplier in a recession is higher than in an expansion at a certain percentile. Panel 3g reports these probabilities at the 5th and 50th percentile for various time horizons. The probabilities range from 83% to 57% at the 5th percentile and from 83% to 90% at the 50th percentile. This suggests that it is highly likely that the cumulative multiplier in a recession is higher, which is in line with the findings in Auerbach and Gorodnichenko (2012, 2013) and Fazzari, Morley and Panovska (2015), but contradicts with the findings in Ramey and Zubairy (2018). This also makes clear that the timing of fiscal stimulus is important. If a stimulus package is intended to lift the economy out of a recession, but the state of the business cycle has already changed, the package may be less effective than policymakers are hoping for.

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[21] While Bayesian estimation precludes the frequentists-style hypothesis testing, our approach can be seen as a way of summarizing the dispersion of the posterior distributions of the multipliers.

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. Boneva *et al.*, 2016; Mertens and Ravn, 2014; Aruoba *et al.*, 2018). 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 strong negative wealth effect and low multipliers due to crowding out of consumption.[22]

In contrast, our results can be reconciled with New Keynesian DSGE models that predict government spending multipliers at the ZLB in the range of 2–5 (see, e.g. Christiano *et al.*, 2011; Eggertsson, 2010; Woodford, 2011; Davig and Leeper, 2011; Coenen *et al.*, 2012). For instance, in models such as Christiano *et al.* (2011), the negative wealth effect of a government spending stimulus is weakened by assuming certain model features. As a consequence, co-movement in consumption, investment and real wages due to counter-cyclical markups is possible.[23] An increase in government spending raises aggregate output, marginal cost and expected inflation. Furthermore, the key channel to explain the higher multipliers at the ZLB is related to the real interest rate. As expected inflation increases and the nominal interest rate is zero, the real interest rate must fall.[24] In consequence, private consumption and investment increases, raises aggregate output, marginal cost and expected inflation once more. Thus, the ZLB amplifies the effects of government spending on output.

In order to underline this point, we have estimated augmented versions of our baseline specification following the 'intermediate strategy' suggested by Burnside, Eichenbaum and Fisher (2004) p. 94) and followed, for instance, in Ramey (2011b) or Fragetta and Gasteiger (2014). The strategy is to specify $Y_t = [G_t, GDP_t, T_t, i_t^{10y}, \bullet_t, F_t]'$, where $\bullet_t$ is a stand-in for variables of interest that we rotate into Y_t one at a time. The rotation variables are: normalized real private consumption (non-durables and services), C_t; normalized real private non-residential fixed investment, I_t; annualized rate of inflation based on the CPI for all urban consumers: all items, π_t; normalized real average hourly earnings of production and non-supervisory employees in manufacturing, w_t; and normalized total hours of wage and salary workers on non-farm payrolls, N_t.
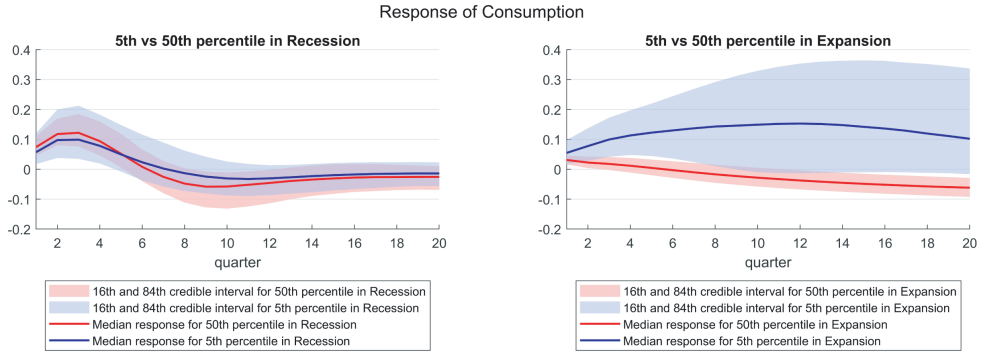
The results are depicted in Figure 4. Again we focus on the recession state that is the IRFs on the LHS.[25] The IRFs of G_t, GDP_t, and T_t, are qualitatively similar to the IRFs discussed above. Moreover, we observe that real private consumption is stimulated to a

[22]An increase in government spending lowers the present value of after-tax income. As a consequence, agents lower consumption and increase labour supply. The latter decreases the real wage and higher employment can raise investment.
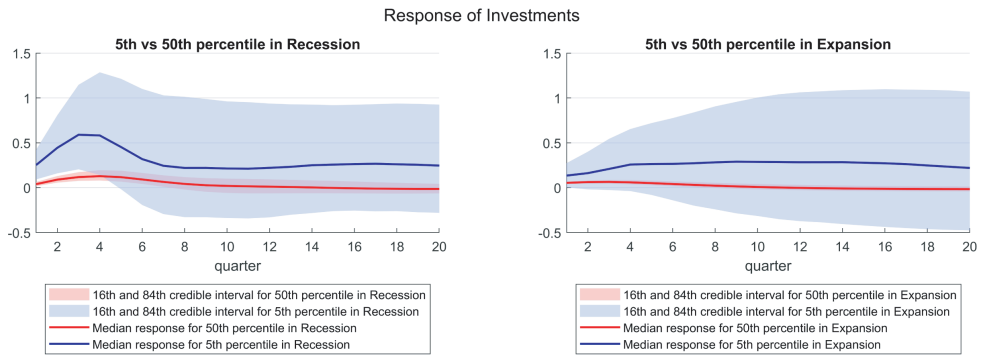
[23]Thus, in such models, multipliers can be large even without considering the ZLB (see, Galí, López-Salido and Vallés, 2007).

[24]Notice that the increase in the 10Y treasury bond rate that we find in our baseline analysis is consistent with the New Keynesian DSGE model in general (see, e.g. Bekaert, Cho and Moreno, 2010). However, the difference in predictions between the low and high interest rate state in recession or expansion is an open question, both theoretically and empirically. We rationalize the stronger increase in IRFs at the low interest rate state by the delayed increase in the nominal interest rate, when the ZLB is binding in a New Keynesian DSGE model. Thus, the increase of more distant expected short-term nominal interest rates may translate into increases in the long-term bond yield. In contrast, when the ZLB is not binding, then the nominal interest rate responds on impact, which may have less or no effect on long-term bond yields.
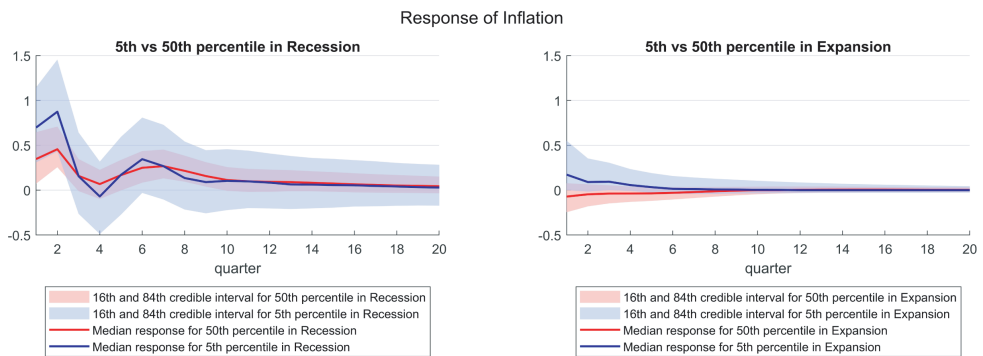
[25]The IRFs during expansions in the RHS panels of Figure 4 overlap for most variables, which makes it difficult to reconcile the IRFs with a theoretical model.

Response of Consumption

5th vs 50th percentile in Recession

5th vs 50th percentile in Expansion

16th and 84th credible interval for 50th percentile in Recession	16th and 84th credible interval for 50th percentile in Expansion
16th and 84th credible interval for 5th percentile in Recession	16th and 84th credible interval for 5th percentile in Expansion
Median response for 50th percentile in Recession	Median response for 50th percentile in Expansion
Median response for 5th percentile in Recession	Median response for 5th percentile in Expansion

(a) Consumption

Response of Investments

5th vs 50th percentile in Recession

5th vs 50th percentile in Expansion

16th and 84th credible interval for 50th percentile in Recession	16th and 84th credible interval for 50th percentile in Expansion
16th and 84th credible interval for 5th percentile in Recession	16th and 84th credible interval for 5th percentile in Expansion
Median response for 50th percentile in Recession	Median response for 50th percentile in Expansion
Median response for 5th percentile in Recession	Median response for 5th percentile in Expansion

(b) Investment

Response of Inflation

5th vs 50th percentile in Recession

5th vs 50th percentile in Expansion

|	
16th and 84th credible interval for 50th percentile in Recession	16th and 84th credible interval for 50th percentile in Expansion
16th and 84th credible interval for 5th percentile in Recession	16th and 84th credible interval for 5th percentile in Expansion
Median response for 50th percentile in Recession	Median response for 50th percentile in Expansion
Median response for 5th percentile in Recession	Median response for 5th percentile in Expansion

(c) Inflation

Figure 4. IRFs to a one standard deviation government spending shock for the augmented baseline specification, FAIVAR-X. 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: (a) consumption; (b) investment; (c) inflation; (d) *ex post* real interest rate; (e) hours (f) real wage
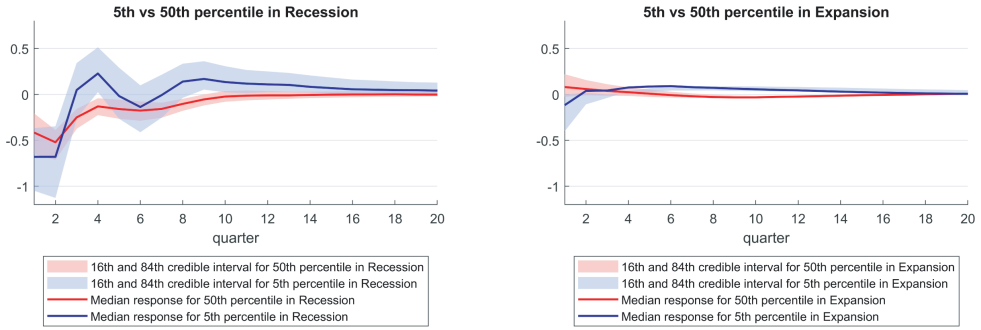
(d) *Ex-post* real interest rate
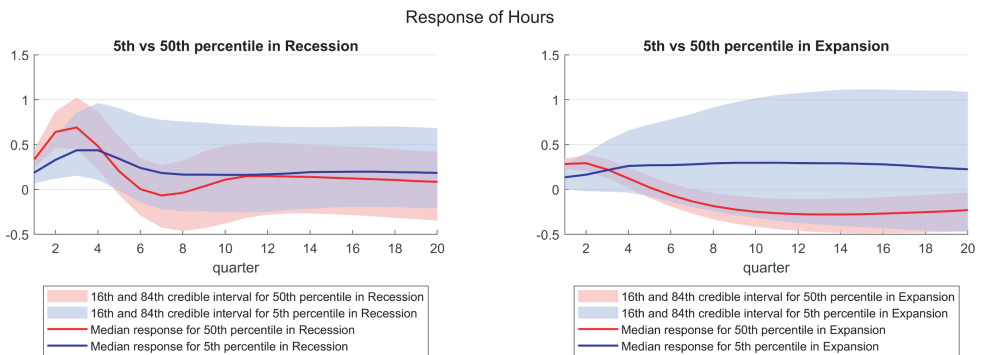
(e) Hours
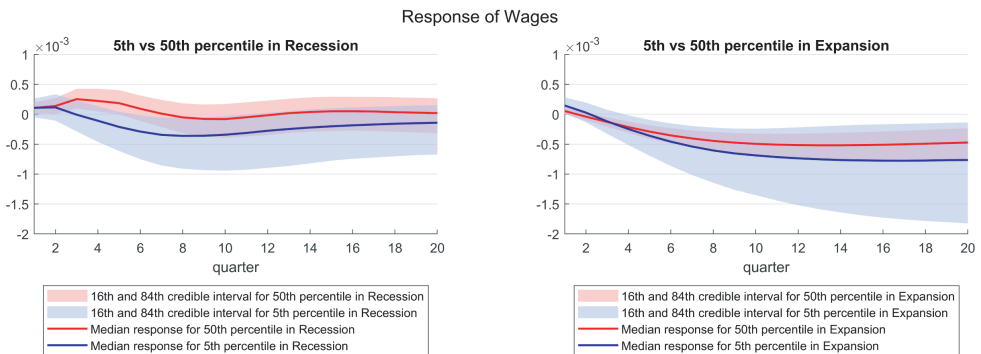
(f) Real wage

Figure 4. (Continued)

similar extent in both the low and high interest rate state. However, real private investment increases far more in the low interest rate state. The differences may be due to a different response of inflation and therefore also of the real interest rate. Panel 4c and 4d depict these responses. One can see that the increase in inflation is much higher on impact in the

low interest rate state. Likewise, the *ex post* real interest rate, computed as the difference of the 10Y treasury bond yield less inflation, shows that on impact and the subsequent two quarters the *ex post* real interest rate declines more in the low interest rate state.[26]

Next, also the IRFs of the real wage and hours are remarkable. First, in both interest rate states the real wage responds positively on impact but the IRFs are not different from zero. Second, in both states hours respond positively on impact with a hump-shaped IRF. Although the impact and peak response is higher in the high interest state, the IRF remains persistently higher for longer in the low interest rate state.

Thus, overall the IRFs of the low interest rate state in a recession are fairly in line with New Keynesian DSGE models that predict large multipliers at the ZLB: the real interest rate declines, consumption and especially investment increase and at the same time hours respond positively while the real wage does hardly respond. In a New Keynesian DSGE model, the latter can be due to a nominal wage rigidity.

One may be concerned about the extent to which the IRFs from our augmented version can be compared to the predictions from a theoretical model. In many theoretical analyses, the effects of government spending at the ZLB crucially hinge on whether the monetary policy instrument is responding. Moreover, in practice, there are good reasons to expect that the effective ZLB is not exactly at zero, but below or above zero, see for example the discussions in Swanson (2018), Swanson and Williams (2014), Bernanke and Reinhart (2004) or Woodford (2011). In addition, this effective ZLB may be difficult to estimate. Thus, in light of these considerations, we think that our framework is able to capture realistic ZLB episodes by considering the lower percentiles of the interest rate distribution in recessions. Therefore, the augmented versions of our baseline specification are one available tool to assess the support of theoretical predictions in the data from our point of view.

IV.   Robustness

The purpose of this section is to assess the robustness of our results with regard to our baseline specification described in section II above. We follow three lines of inquiry. First, we maintain the identification approach with sign restrictions detailed in section II and assess the role of limited information. In particular, we discuss results for variations of our baseline specification. Second, we use an alternative strategy to account for fiscal foresight based on forecast errors instead of growth in the forecast of government spending. Finally, we use our baseline specification, but identify the government spending shock via timing instead of sign restrictions. It turns out that our baseline results are robust to almost all of these modifications of the research design. The exception is the findings for the alternative strategy to account for fiscal foresight based on forecast errors. In this case, we do not find a notable difference between multipliers in the low and high interest rate state for time horizons of 3 years or longer. All other findings are confirmed.

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[26]Ideally, one would like to compute the IRFs of expected inflation and/or the *ex ante* real interest rate to be fully consistent with the New Keynesian DSGE model. However, this is not feasible within our framework. Nevertheless, the above reported IRFs are approximations that are feasible within our framework.

The role of interaction terms and limited information

The key distinguishing feature of our model relative to linear VAR models widely used in the literature estimating the fiscal multiplier are the interaction terms that allow for state dependence in government spending multipliers. However, our model has additional features that are commonly used in the literature to account for the limited information problem: principal components F_t used to proxy latent unobserved factors driving macroeconomic variables and the growth in government spending forecasts $Z_{t|t-1}$. In order to assess the role of each of these features, we carry out three exercises.

IVAR. We examine the role of the interaction terms in generating our main results by excluding F_t and $Z_{t|t-1}$ from our baseline specification. Nevertheless, we do think that excluding these two features from the model has the potential to bias the estimates. The IVAR multipliers reported in Figure 5 are obtained in the same way as for the baseline specification. Overall, compared to the baseline results (Figure 3), one can observe that multipliers, independent of the state, are lower at the 1- and 2-year horizon, but higher at longer horizons. Moreover, on average, multipliers are higher at the low interest rate state and in recessions. Remarkably, while multipliers show a monotonic decline from the 1st to the 75th percentile in expansions, they show a non-monotonic behaviour in recessions, being largest at the 13th percentile, which we consider to be part of the low interest rate state. Thus, even if the IVAR does not address the limited information problem in general or fiscal foresight in particular, our baseline results for the state-dependent government spending multiplier are confirmed qualitatively.

FAIVAR. Next, we examine the contribution of accounting for the general limited information problem by adding principal components F_t to the IVAR specification. We continue to exclude the growth in government spending forecasts $Z_{t|t-1}$. As one can see from Figure 6, independent of the state, the multipliers are lower at all horizons relative to the baseline specification (Figure 3), but higher relative to the IVAR (Figure 5). Therefore, not accounting for fiscal foresight in the specification appears to generate a downward bias in our baseline estimates for the government spending multiplier. Moreover, one can observe that multipliers based on the FAIVAR are higher at the low interest rate state and in recessions. Finally, independent of the state of the business cycle, there is a monotonic decline from the 1st to the 75th percentile as in the baseline specification. The latter suggests that adding the principle components removes the non-monotonicity of multipliers in recessions observed for the IVAR. In sum, the findings for the FAIVAR confirm our findings for the baseline specification.

IVAR-X. Finally, we focus on the effect of controlling for fiscal foresight by adding government spending forecasts $Z_{t|t-1}$ in our specification, but excluding principal components F_t. The multipliers in Figure 7 are lower than the baseline multipliers in Figure 3 for the 1- and 2-year horizon, but larger at longer horizons. At the same time, these multipliers are higher compared to the IVAR (Figure 5). Thus, not addressing fiscal foresight in our specification implies an upward bias in our estimates at longer horizons. In addition, the IVAR-X multipliers are again higher at the low interest rate state and in recessions. However, there is again a non-monotonicity in multipliers in recessions, which suggests that this non-monotonicity emerges due to not accounting for the general limited information problem.
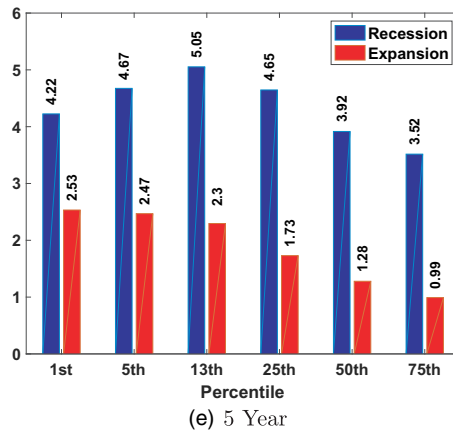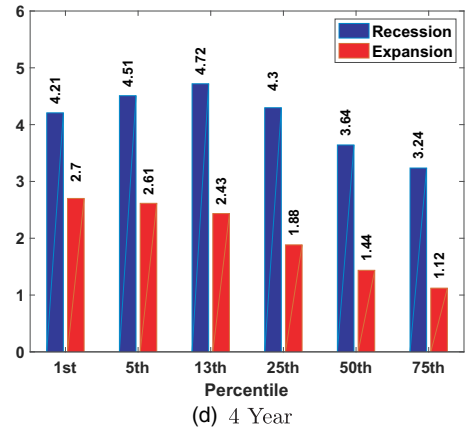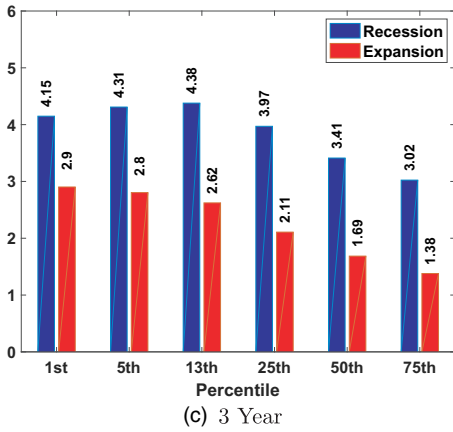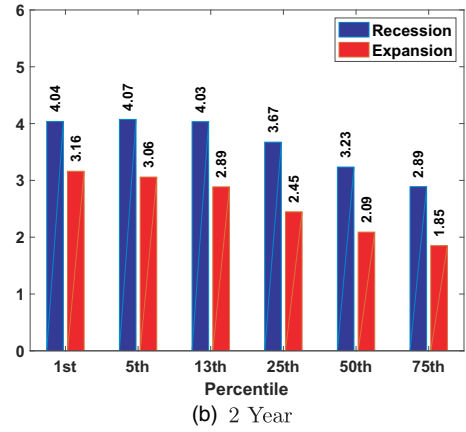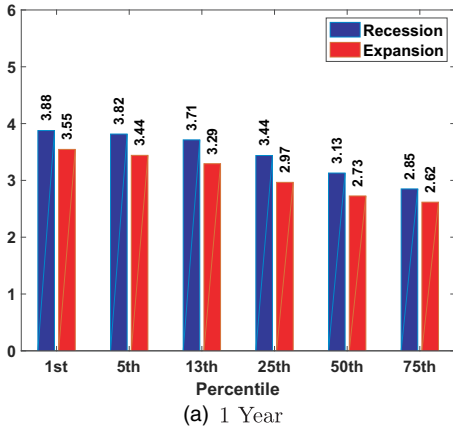
Figure 5. Cumulative government spending multipliers at several time horizons for the IVAR model: (a) 1 1year (b) 2 year; (c) 3 year; (d) 4 year; (e) 5 year
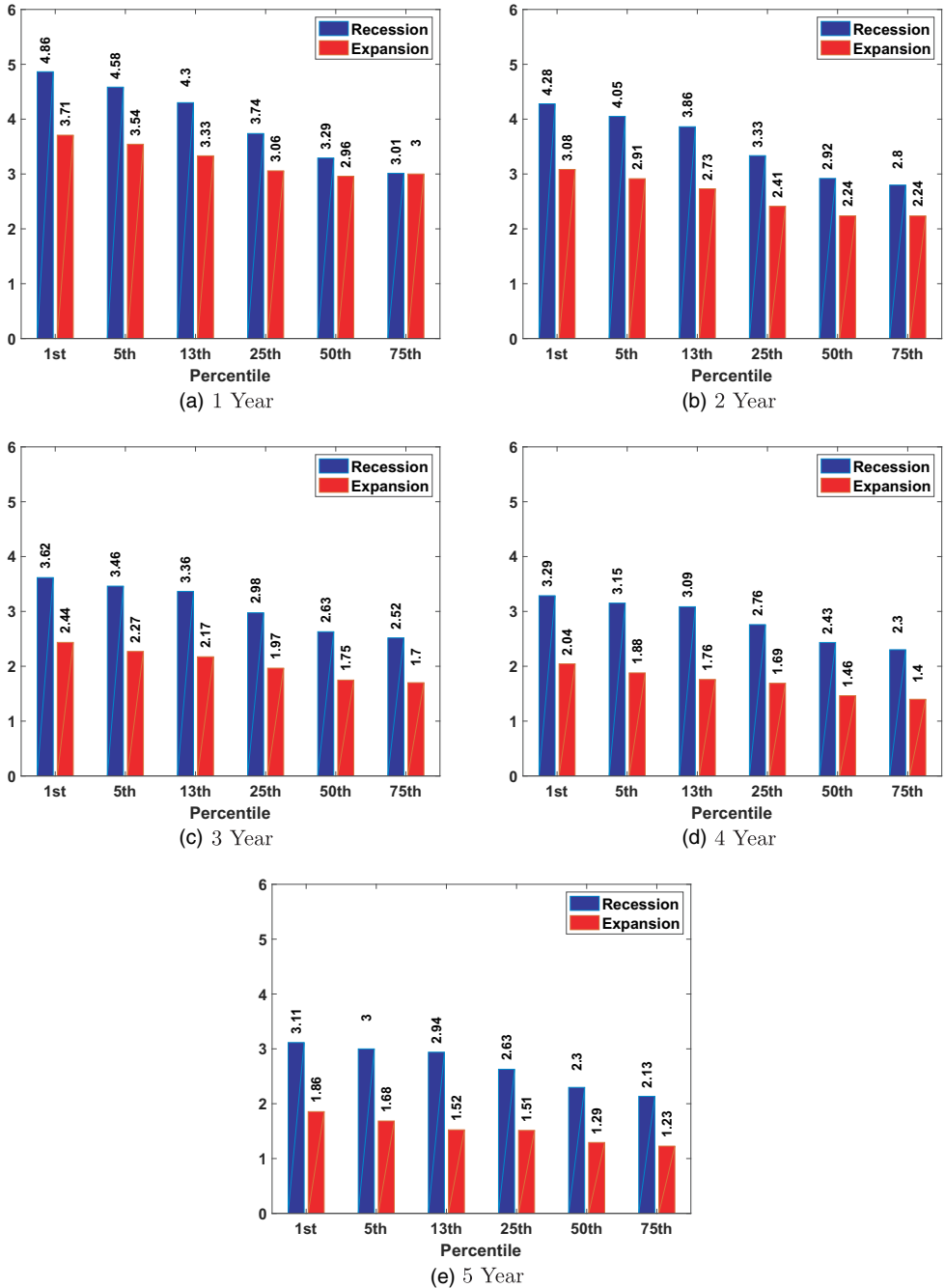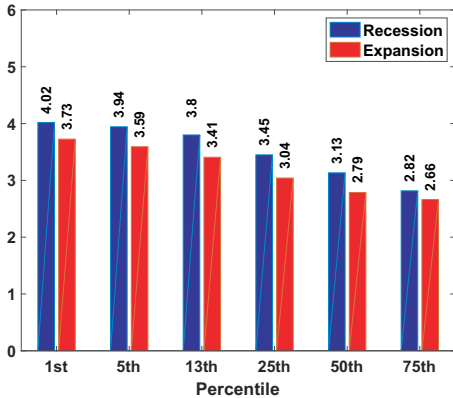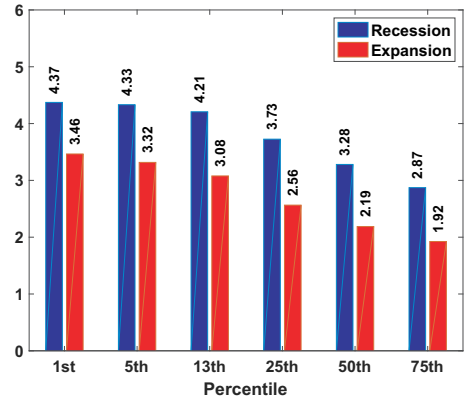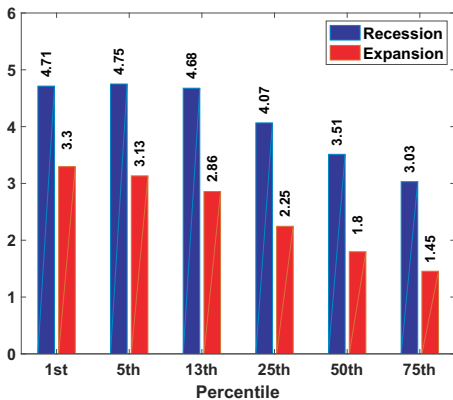
Figure 6. Cumulative government spending multipliers at several time horizons for the FAIVAR model: (a) 1 year; (b) 2 year; (c) 3 year; (4) 4 year; (e) 5 year

Figure 7. Cumulative government spending multipliers at several time horizons for the IVAR-X model: (a) 1 year; (b) 2 year; (c) 3 year (d) 4 year; (e) 5 year

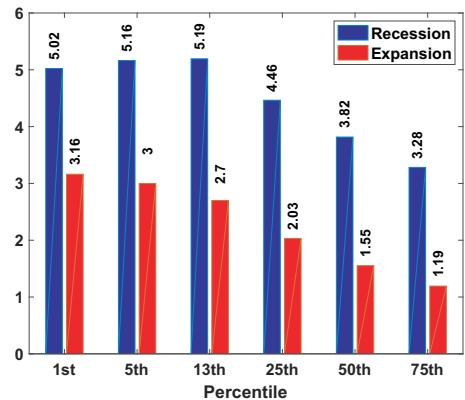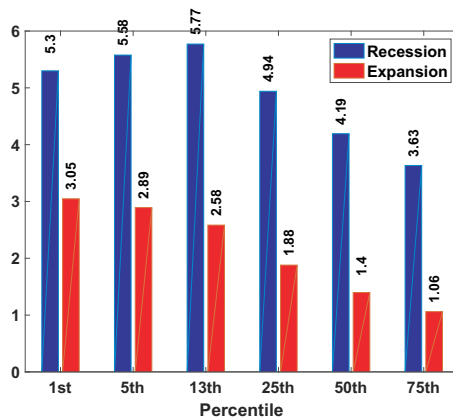### Forecast errors

While the FAIVAR specification in the robustness exercise above did not account for fiscal foresight on purpose, it is possible to do so within the FAIVAR model, but without relying on the exogenous variable $Z_{t|t-1}$. Doing so enables us to assess robustness of our main results to an alternative strategy of accounting for fiscal foresight. Following Auerbach and Gorodnichenko (2012), we construct the forecast error variable, FE_t, use it as endogenous variable and abandon the exogenous SPF forecast. Thus, we turn the FAIVAR-X model described in equation (1) into a FAIVAR model, by suppressing the vector of exogenous variables $Z_{t|t-1}$. The vector of endogenous variables is now given by $Y_t = [FE_t, G_t, GDP_t, T_t, i_t^{10y}, F_t]'$, where the FE_t series is constructed as the difference between the forecast made at previous quarter $t-1$ for the contemporaneous quarter t minus the actual government spending. Intuitively, the forecast error represents the surprise experienced by private agents about the actual volume of the government spending policy.[27] We also 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.

The procedure to derive government spending multipliers does not change in comparison to section II. However, the set of sign restrictions that we apply to derive IRFs has to be modified as illustrated in Table 1. In order to derive an unexpected government spending shock, we restrict the forecast error to be positive for at least 1 quarter and, Government Spending and GDP to be positive for at least 4 quarters. Figure 8 shows the multipliers that we obtain with this specification.

In comparison to the baseline results (Figure 3), the estimated multipliers behave similar along several dimensions: independent of the state, multipliers decline with the horizon and, on average, are higher at the low interest rate state. However, remarkable differences between the low and high interest rate state can only be found for the 1- and 2-year horizon, but not at longer horizons. In contrast, the pattern that multipliers are higher in recessions than in expansions is more robust as it is evident at all considered time horizons.

The computed probabilities that the cumulative multipliers at the 5th percentile are higher than the ones at the 50th percentile at various time horizons in Panel 8f are consistent with this interpretation of results. They are lower or equal to their counterparts in Panel 3f, and, while the probability is above 50% at the 1- and 2-year horizon, it is equal to or below 50% at longer horizons. Likewise, the probabilities that the cumulative multipliers in recession are higher than the ones in expansion at various time horizons are displayed in Panel 8g. Again, these probabilities are lower or equal to their counterparts in Panel 3g and range from 89% to 64%, depending on the time horizon and interest rate state.

In sum, with this alternative method to account for fiscal foresight, our main results for the state-dependent government spending multiplier are fully confirmed at the 1- and 2-year horizon. Over this short horizon, multipliers are higher in the low interest rate state relative to the high interest rate state. However, at longer horizons we can only confirm our findings regarding the business cycle: multipliers are higher in recessions independent of the interest rate state.
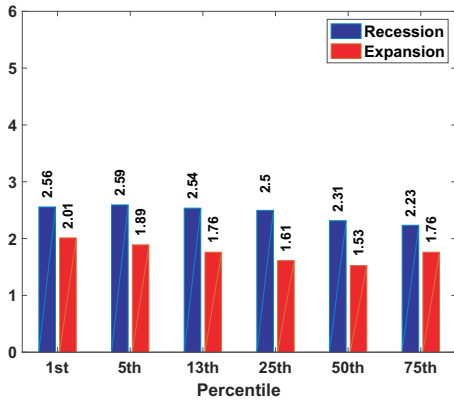
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[27]Appendix A contains further information on the computation of this variable.
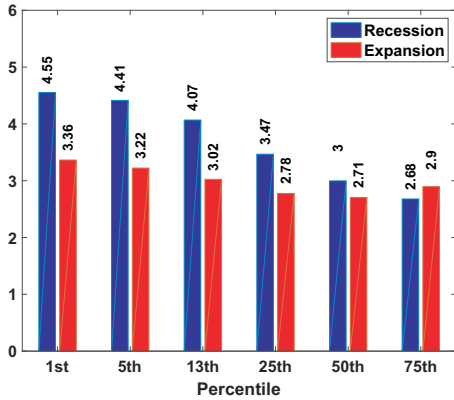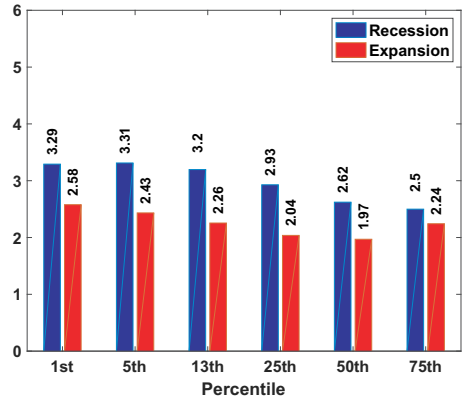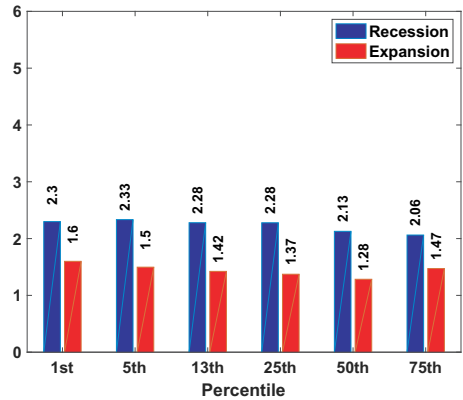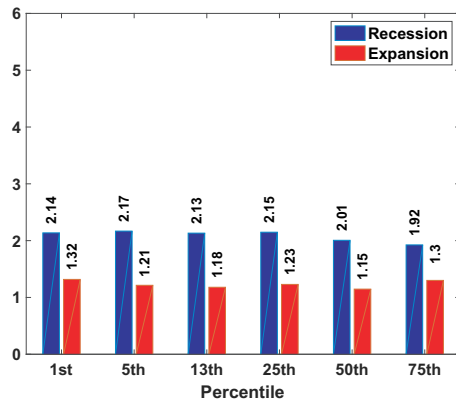
Figure 8. Cumulative government spending multipliers at several time horizons and probabilities for differences in multipliers for the FAIVAR model with forecast errors, FE$_t$: (a) 1 year; (b) 2 year; (c) 3 year; (d) 4 year; (e) 5 year; (f) probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile; (g) probability that the cumulative multiplier in a recession is higher than in an expansion
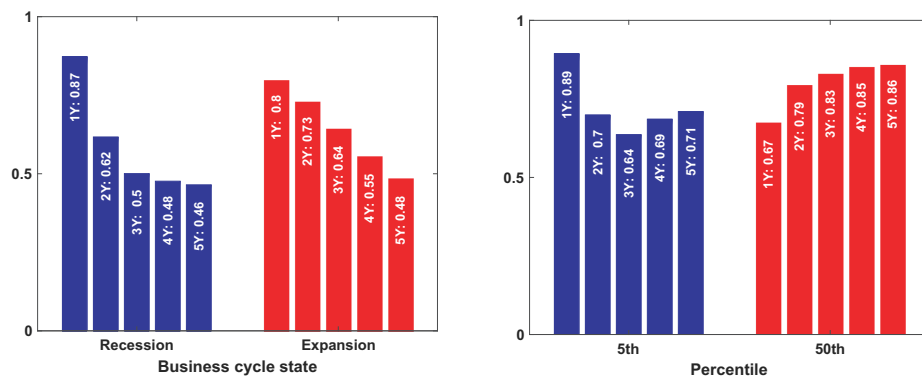
(f) Probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile.

(g) Probability that the cumulative multiplier in a recession is higher than in an expansion.

Figure 8.   (Continued)

The role of sign restrictions

The findings for the IVAR model above suggest that the non-linearity introduced by the interaction term in the IVAR can explain the discrepancy between our main findings of rather large multipliers in comparison to the rather low multipliers found in linear VARs that utilize sign-restrictions.[28] For instance, our IVAR results are not directly comparable to estimates based on the linear structural VAR model with sign-restrictions in the important paper by Mountford and Uhlig (2009). Their set-up differs in many ways from ours: they consider a different sample, identify additional shocks and identify three policy scenarios (deficit spending, deficit-financed tax cuts and a balanced budget spending expansion). In contrast, in line with Ramey and Zubairy (2018) and others, we just identify the government spending shock. However, the most relevant difference between the approach in this paper and Mountford and Uhlig (2009) may be that we not only restrict government spending to increase for four quarters, but also that output has to do so. Therefore, one may be concerned whether our findings critically hinge on this particular identification approach. In order to address this concern, we re-estimated our baseline specification, the FAIVAR-X model, but identify the government spending shock with timing restrictions by applying the Cholesky decomposition. This strategy follows Blanchard and Perotti (2002). The government spending shock is identified by ordering government spending as the first variable, that is, one assumes that government spending does not respond contemporaneously to any of the other variables in our model due to implementation and legislation lags. More important, the sign and duration of the response of output to a shock in government spending is not restricted.

Figure 9 reports multipliers for this exercise. A first observation is that independent of the time horizon, multipliers are lower compared to our main results (Figure 3). At the low interest rate state, multipliers are higher relative to the high interest rate state and decline monotonically from the 1st to the 75th percentile of the interest rate distribution. Both of

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[28]Note that an alternative way of capturing the potential nonlinearity emerging from the ZLB would be to use a regime-switching VAR with sign-restrictions. An example of such an approach is Liu *et al.* (2019).
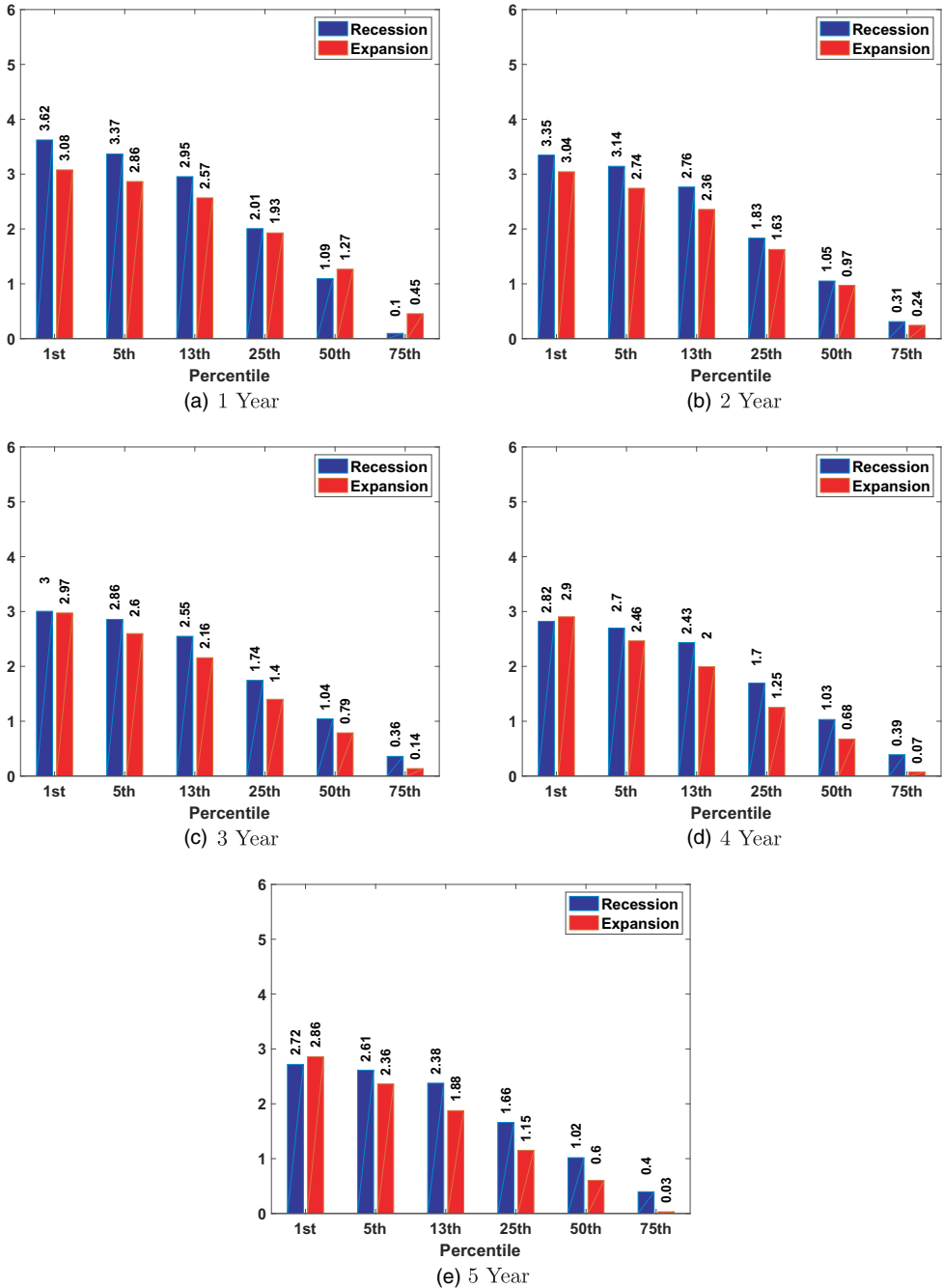
Figure 9. Cumulative government spending multipliers at several time horizons and probabilities for differences in multipliers for the FAIVAR-X model with identification via timing-restrictions: (a) 1 year; (b) 2 year; (c) 3 year; (d) 4 year; (e) 5 year; (f) probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile; (g) probability that the cumulative multiplier in a recession is higher than in an expansion
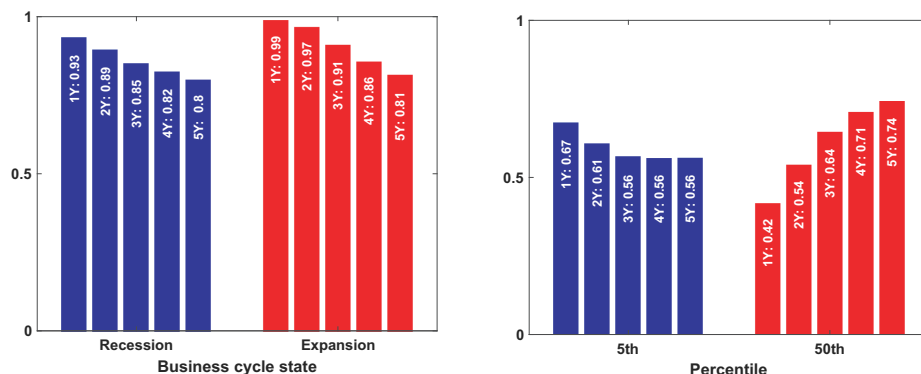
(f) Probability that the cumulative multiplier at the 5th percentile is higher than the one at the 50th percentile.

(g) Probability that the cumulative multiplier in a recession is higher than in an expansion.

Figure 9.   (Continued)

these findings are independent of the state of the business cycle. In addition, consistent with our baseline findings, multipliers in recessions exceed multipliers in expansions for most percentiles at all time horizons, however, the differences are very small.

Moreover, also for this exercise, we have re-calculated the probabilities that the cumulative multipliers at the 5th percentile exceed their counterparts at the 50th percentile at various time horizons, see Panel 9f. Relative to our baseline results in Panel 3f, this probability is higher at almost each time horizon in both recession and expansion. It ranges from 80% to 93% in recession and from 81% to 99% in expansion.

Finally, Panel 9g reports the probability that the cumulative multiplier in a recession exceeds the one in an expansion at a certain percentile. In comparison to the probabilities for the main results in Panel 3g, the probabilities appear to be equal or lower, depending on the time horizon. The probability ranges between 67% and 56% for the 5th percentile and between 42% and 74% for the 50th percentile. Nevertheless, consistent with our main results, the probabilities suggest that it is likely that the cumulative multipliers in a recession are higher than the ones in an expansion. Overall, these findings suggest that the choice of sign-restrictions that generated our main results is not crucial for our findings regarding the size of the government spending multiplier at the ZLB.

V.   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 a FAIVAR model and use sign restrictions to identify government spending shocks. This framework allows us to account for fiscal foresight as well as the generic limited information problem inherent in VARs and to estimate state-dependent multipliers at all percentiles of the nominal interest rate distribution, both during recessions and expansions.

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. The multipliers during recessions are also larger than during expansions and

largest during recessions in the low interest rate state. For our sample from 1966 to 2015, the 5-year cumulative multipliers in a recession are in the range of 3.56 to 3.79 at the ZLB. The corresponding ones away from the ZLB are between 2.31 and 3.05. These results are robust along several important dimensions including modifications of the baseline specification, alternatives of accounting for fiscal foresight, and identification via timing restrictions.

We also estimate augmented versions of our baseline specification with consumption, investment, inflation, hours worked, and the real wage in order to obtain a more complete picture of the transmission mechanism of a government spending shock. We find significant evidence for a decline in the real interest rate, an increase in consumption and especially investment, and at the same time hours respond positively while the real wage does hardly respond. These predictions are typical for New Keynesian DSGE models. Thus, we conclude that the government spending multiplier at the ZLB is larger than in normal times as predicted by many recently developed New Keynesian DSGE models.

Appendix A: 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.

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 $t-1$, that is

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

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

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

Thus, for this purpose, we have downloaded first release data on real government consumption expenditures and gross investment: state and local (RGSL) from this (https://www.philadelphiafed.org/research-and-data/real-time-center/real-time-data/data-files/rgsl) website and real government consumption and gross investment: federal (RGF) from this (https://www.philadelphiafed.org/research-and-data/real-time- center/real-time-data/data-files/rgf) website. All in quarterly vintages (Billions of real dollars, seasonally adjusted). G_t^{1st} is the sum of RGSL and RGF.

TABLE 2

Data

Series	Source	Mnemonic code	Transformation
Forecasts of Real Federal Government Consumption Expenditures & Gross Investment	Federal Reserve Bank of Philadelphia	RFEDGOV	
Forecasts of Real State and Local Government Consumption Expenditures and Gross Investment	Federal Reserve Bank of Philadelphia	RSLGOV	
Greenbook projections of Real Federal Government Consumption and Gross Investment	Federal Reserve Bank of Philadelphia	gRGOVF	
Greenbook projections of Real State and Local Government Consumption and Gross Investment	Federal Reserve Bank of Philadelphia	gRGOVSL	
Real Government Consumption and Gross Investment: Federal	Federal Reserve Bank of Philadelphia	RGF	
Real Government Consumption and Gross Investment: State and Local	Federal Reserve Bank of Philadelphia	RGSL	
Forecast Error of the Annualized Growth Rate of Real Government Purchases	All the above variables are used for the computation, see Appendix A.		Normalized
Nominal Government Consumption Expenditures and Gross Investment	US Bureau of Economic Analysis	GCE	Real, Normalized
Federal Government Current Tax Receipts	US Bureau of Economic Analysis	W006RC1Q027SBEA	Real, Average w.r.t. GDP
State and Local Government Current Tax Receipts	US Bureau of Economic Analysis	W070RC1Q027SBEA	Real, Average w.r.t. GDP

(Continued)

TABLE 2

(Continued)

Series	Source	Mnemonic code	Transformation
Real Gross Domestic Product	US Bureau of Economic Analysis	GDPC1	Normalized
Gross Domestic Product: Implicit Price Deflator	US Bureau of Economic Analysis	GDPDEF	
Long-Term Government Bond Yields: 10-year: Main (Including Benchmark) for the United States	Organization for Economic Co-operation and Development	IRLTLT01USQ156N	
Shadow Federal Funds Rate	Wu and Xia (2016)		
NBER based Recession Indicators for the United States from the Peak through the Trough	Federal Reserve Bank of St. Louis	USRECQM	
Real Potential Gross Domestic Product	US Congressional Budget Office	GDPPOT	
Personal Consumption Expenditures: Non-durable Goods	US Bureau of Economic Analysis	PCND	Normalized
Personal Consumption Expenditures: Services	US Bureau of Economic Analysis	PCESV	Normalized
Private Non-residential Fixed Investment	US Bureau of Economic Analysis	PNFI	Normalized
Hours of Wage and Salary Workers on Non-farm Payrolls: Total	US Bureau of Labor Statistics	TOTLQ	
Average Hourly Earnings of Production and Non-supervisory Employees: Manufacturing	US Bureau of Labor Statistics	CES3000000008	Normalized
Consumer Price Index for All Urban Consumers: All Items in US City Average	US Bureau of Labor Statistics	CPIAUCSL	

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 B: 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

$$\Delta G_t = \beta_0 FE_t + \sum_{k=1}^{L} \beta_k FE_{t-k} + \varepsilon_t, \quad \Delta G_t \equiv \left[\left(\frac{G_t}{G_{t-1}} \right)^4 - 1 \right] \times 100. \tag{B.1}$$

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 ΔG_t. Therefore, our results for $L > 0$ imply that the unpredictable components in ΔG_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 that is

$$\Delta G_t = \sum_{k=1}^{L} \beta_{k,G} G_{t-k} + \sum_{k=1}^{L} \beta_{k,GDP} GDP_{t-k} + \sum_{k=1}^{L} \beta_{k,T} T_{t-k} + \sum_{k=1}^{L} \beta_{k,i^{10y}} i_{t-k}^{10y} + \varepsilon_t. \tag{B.2}$$

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

TABLE 3

Explanatory power of FE_t[a]

	R-squared	F-statistic	Marginal F-statistic
$L = 0$			
1966Q4-2015Q4	0.261	68.71	
1966Q4-2015Q4	0.060		3.05
$L = 1$			
1966Q4-2015Q4	0.266	35.19	
1966Q4-2015Q4	0.071		3.65

Notes: [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.

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