The Euro-Area Government Spending Multiplier at the Effective Lower Bound*

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Abstract

We build a factor-augmented interacted panel vector-autoregressive model of the Euro Area (EA) to estimate government spending multipliers. The multipliers are contingent on the overall monetary policy stance, captured by a shadow monetary policy rate. The econometric approach deals with several technical problems highlighted in the empirical macroeconomic literature, including the issues of fiscal foresight and limited information. We find that the output response to a government spending shock when the monetary policy rate is at the effective lower bound (ELB) is significantly higher than during normal times. Depending on the specification, while in normal times the average medium-term cumulated multiplier ranges between 0.9 and 1.2, in the ELB period, the range is between 2.5 and 3.4. More generally, the multiplier is inversely correlated with the level of the shadow monetary policy rate. While we verify that EA data give support to the view that the multiplier is larger in periods of economic slack, we show that the shadow rate and the state of the business cycle have an autonomous impact on its size.

JEL classification: C32, C33, C38, E62.
Keywords: Fiscal multiplier, Zero lower bound, Panel VAR, Factor models, Euro Area.

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1 Introduction

The Global Financial Crisis (GFC) and the subsequent Great Recession pushed many governments in advanced economies—notably the US, EU member countries and Japan—to use fiscal policy as a discretionary tool to soften the adverse effects of the economic contraction. In 2008 the European Commission launched the “European Economic Recovery Plan” (EERP) to provide fiscal stimulus to the euro area (EA) economies. More recently, some EA countries experienced fiscal shocks of opposite sign as their governments went through austerity measures to put their finances back on track in response to the EA sovereign debt crisis.

In both types of situations, policymakers are confronted with a crucial question: what is the size of the government spending multiplier? While at the onset of the GFC there was already a large literature providing estimates of this multiplier, in academic and policy circles alike it became soon clear that those estimates were likely not applicable to the new economic environment. One reason is that after the GFC, monetary policy ceased to operate in the conventional way. The policy rate was lowered repeatedly until it reached its effective lower bound (ELB), and the ELB itself turned out to be a dynamic concept. While at the beginning of the GFC it was believed to coincide with the zero lower bound (ZLB), at a certain point several central banks—the European Central Bank (ECB), Danmarks Nationalbank, Sveriges Riksbank, Swiss National Bank, and Bank of Japan—adopted negative interest rate policies. Negative interest rates were not the only unconventional measures. Central banks started implementing also asset purchase programs (often dubbed as quantitative easing–QE), and forward guidance to affect long-term interest rates and boost aggregate demand.

Focusing on the EA, this paper seeks to answer the following question: when monetary policy is constrained by the ELB and operates in an unconventional manner, is the government spending multiplier different from the multiplier observed in normal times?

We tackle this important question from a purely empirical viewpoint. To overcome the difficulty of capturing the enhanced intricacy that has lately characterized monetary policy in the EA, our approach is to condition the computation of the multiplier on an indicator that summarizes the overall monetary policy stance and takes expectations on future policy into account. A prominent indicator with this desirable features has recently been the shadow rate (SR) derived by Wu and Xia (2016) and Wu and Xia (2017) using an approximation of a nonlinear term structure model. In Figure 1 we report the EA shadow rate along with the observed euro overnight index average (Eonia) rate. The two rates almost overlapped before the inception of the GFC. They started to diverge after 2008q3, a quarter made infamous in economic history by the bankruptcy of Lehman Brothers. Then, after the well-known
Figure 1: Euro Overnight Index Average (Eonia) Rate and Shadow Monetary Policy Rate in the Euro Area

Sources: European Central Bank and Wu and Xia (2017).

“Whatever it takes” speech of ECB’s President Mario Draghi, the Eonia rate was brought first to the ZLB and then turned negative, while the SR continued to sink into the negative territory.

To fully take the dynamics of the shadow rate into account, we use a factor-augmented interacted panel vector-autoregressive model purified of expectations (FAIPVAR-X), an extension of the IPVAR model by Towbin and Weber (2013) and Sá et al. (2014). Using this framework has four advantages. First, the panel dimension allows exploiting quarterly data of ten EA countries (Austria, Belgium, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal and Spain) that were part of the European Monetary Union (EMU) since its inception.\(^1\) Second, the presence of an interaction term allows capturing nonlinearities and estimating the reaction of the variables of interest to a government spending shock at each

\(^1\)In line with Auerbach and Gorodnichenko (2013), we exclude Luxembourg being it a small economy with large and volatile changes in government spending series.
percentile of the shadow rate. Third, augmenting the specification with factors extracted from a large number of macroeconomic variables addresses limited information concerns. In fact, there is likely important information that we do not explicitly include in our model, but that might have been used by economic agents in making their choices (see Bernanke et al., 2005; Stock and Watson, 2005; Fraretta and Gasteiger, 2014). Fourth, including forecasts of government spending as an exogenous variable purges government spending shocks from its anticipated component and addresses the issue of fiscal foresight (see, e.g., Forni and Gambetti (2010), among others). Addressing the issues of limited information and fiscal foresight resolves what the literature calls non-fundamentalness, a problem stemming essentially from a misalignment between the information sets of economic agents and the econometrician. Failing to rectify this problem would bias the results.

Turning to the answer to our research question, results are as follows: (i) the output response to a government spending shock when the monetary policy rate is at the ELB is significantly higher than it is during normal times; (ii) depending on the specification, while in normal times the average medium-term (five-year) cumulated multiplier ranges between 0.9 and 1.2, in the ELB period, the range is between 2.5 and 3.4; (iii) more generally, the multiplier is inversely correlated with the level of the shadow rate. We also find the size of the multiplier to be inversely correlated with the business cycle after controlling for the level of the shadow rate, lending support to the view that the multiplier is larger in periods of economic slack (in line, e.g., with Auerbach and Gorodnichenko, 2013). However, we verify that the shadow rate and the state of the business cycle have an autonomous impact on its size.

Our paper is related to a number of contributions in the literature. The literature that uses calibrated or estimated dynamic stochastic general equilibrium (DSGE) models does not agree on whether, or in which direction, a binding ZLB for the monetary policy rate should alter the government spending multiplier. A number of contributions (Christiano et al., 2011; Coenen et al., 2012; Davig and Leeper, 2011; Eggertsson, 2010; Kilponen et al., 2015; Woodford, 2011, among others) argue in favor of a higher multiplier at the ZLB. According to these studies, the government spending multiplier at ZLB is in the range of 2 to 5. In contrast, a host of equally rigorous papers (Cwik and Wieland, 2011; Braun et al., 2013; Mertens and Ravn, 2014; Aruoba et al., 2017, among others) claim the multiplier to be small at the ZLB, sometimes even smaller than in normal times. We ascribe this disagreement to the inherent difficulty in building DSGE models encompassing all the complexity that characterized monetary policy making in the aftermath of the GFC. In the empirical literature, although with a different methodology than ours, Ramey and Zubairy (2018) estimate the U.S. government spending multiplier at the ZLB. They find mixed results that depend on
the sample period. When excluding the World War II period, they find a multiplier at the ZLB up to 1.5. To our knowledge, the literature lacks empirical estimates of the multiplier at the ELB for the EA. We fill in this gap. By conditioning the multiplier on the level of the shadow rate over virtually the entire history of the EMU, the choice of the estimation sample ceases to be an issue. Furthermore, our estimation strategy implies studying not only how the size of the multiplier is affected by the ELB, but also by the whole set of unconventional monetary policies implemented after the GFC in the eurozone.

The remainder of the paper is structured as follows. Section 2 explains the empirical methodology and the econometric specification. Section 3 reports the results. Section 4 presents robustness checks. Finally, Section 5 concludes. Data sources and details on robustness checks are appended to the paper.

2 Methodology

2.1 Empirical Model

The empirical model builds on the Interacted Panel Vector Auto-Regressive (IPVAR) framework developed by Towbin and Weber (2013) and Sá et al. (2014). This model is well suited to our purposes because the presence of interaction terms allows us to capture nonlinearities in the reaction of variables of interest to government spending shocks conditional on the whole distribution of the nominal shadow interest rate.

The model specification takes the following general structural form:

$$B_{i,t}y_{i,t} = \sum_{j=1}^{N} \kappa_j D_{j,i} + \sum_{j=1}^{N} \sum_{k=1}^{L} \Gamma_{j,k} D_{j,i} y_{i,t-k} + \sum_{j=1}^{N} \kappa_{j1} D_{j,i} x_t + \sum_{k=1}^{L} \Gamma_{1,k} x_t y_{i,t-k}$$

$$+ \sum_{j=1}^{N} v_j f(t-1:t-4) + v^1 z_{t-1} + \varepsilon_{i,t},$$

(1)

where \( t = 1, \ldots, T \) denotes the time dimension; \( j = 1, \ldots, N \) denotes the country dimension; and \( k = 1, \ldots, L \) represents the lag structure. The vector of endogenous variables is denoted by \( y_{i,t} \); the interaction term is represented by \( x_t \); while the vectors of two sets of exogenous variables are denoted by \( f(t-1:t-4) \) (discussed in Subsection 2.2) and \( z_{t-1} \) (foreign exogenous variables, also discussed in Subsection 2.2). Furthermore, coefficient \( \kappa_j \) is the country-specific intercept of country \( j \); \( \kappa_{j1} \) is the country-specific coefficient of the interaction term; \( \Gamma_{j,k} \) is the matrix of autoregressive coefficients attached to the endogenous variables; \( v_j \) is the matrix of country-specific coefficients attached to the first set of exogenous variables; \( v^1 \)
represents the pooled estimated coefficients of the another set of exogenous variables; \( D_{j,i} \) is an indicator variable for each country (equal to 1 if \( i = j \), and 0 otherwise); and, lastly, \( \varepsilon_{i,t} \) is a vector of i.i.d. residuals, which are uncorrelated across countries by assumption. It is noteworthy that the interaction term, \( x_t \), affects both the level and the dynamic relationship across endogenous variables through \( \kappa^j \) and \( \Gamma^k \) (for more technical details on the IPVAR framework see, e.g., Sá et al., 2014).

To allow for as much heterogeneity as possible, we utilize a panel model with fixed effects and heterogeneous slopes, which we estimate using the mean group estimator. This estimator has been shown to perform better than alternative estimators in dynamic panels with heterogeneous slopes (see, e.g., Pesaran and Smith, 1995 and Canova and Ciccarelli, 2013, among the others). Due to data availability constraints we estimate homogeneous slopes of lagged interacted terms, \( x_t y_{i,t-k} \), and foreign exogenous variables, \( z_{t-1} \).

Matrix \( B_{i,t} \) is a \((q \times q)\) lower triangular matrix with ones on the main diagonal. The recursive structure imposed on matrix \( B_{i,t} \) implies that the covariance matrix of the residuals, \( \Sigma \), is diagonal. Given that the FAIPVAR-X model requires the estimation of a large number of parameters, for the sake of parsimony, we produce the baseline results with a uniform lag structure on one quarter \((L = 1)\). However, we re-run the estimation also with two lags for robustness (Section 4).

### 2.2 Data and Baseline Specification

Our dataset is composed of quarterly data and covers the period from 2002q2 to 2017q4.\(^2\) We consider ten of the eleven countries that joined the EA when it came into existence: Austria, Belgium, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal and Spain. In line with Auerbach and Gorodnichenko (2013), we exclude Luxembourg being a small economy, which exhibits large and volatile changes in government spending series. For details on the construction of the dataset, see Appendix A.

To examine the macroeconomic effects of fiscal shocks, the VAR literature has traditionally used variations of the following vector of endogenous variables:

\[
y_{i,t} = [G_{i,t}, GDP_{i,t}, T_{i,t}]'
\]

where \( G_{i,t}, GDP_{i,t} \) and \( T_{i,t} \) represent real government purchases (the sum of government gross fixed capital formation and government consumption), real gross domestic product and real net taxes (the sum of government receipts of direct and indirect taxes minus transfers to

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\(^2\)The beginning of our sample is dictated by the availability of the Economist Intelligence Unit forecasts of government spending, the use of which is explained below.
businesses and individuals), respectively. We make a number of modifications to this simple specification to overcome a series of issues we discuss below.

First, to simplify the procedure related to the computation of government spending multipliers, we divide all endogenous variables by the real potential GDP of the corresponding country. This way there is no need to take the logarithm of the variables and perform *ex-post* conversions of the estimated elasticities to dollar equivalents, avoiding potential biases. In fact, *ex-post* conversion requires the use of constant sample averages of the ratios of fiscal variables to GDP, which may instead vary over time, potentially biasing the size of the multipliers. This problem is even more acute in nonlinear models, such as that adopted in this paper (for more details on this issue see, e.g., Gordon and Krenn, 2010 and Ramey and Zubairy, 2018, among others). We compute real potential GDP using the filter recently proposed by Hamilton (2018), which avoids the spurious persistence in the cyclical component implied by the traditional Hodrick-Prescott (HP) filter.

Second, to the basic set of endogenous variables listed in vector $y_{i,t}$, we add common factors, via principal components, extracted from a large number of macroeconomic times series. In fact, VAR models are characterized by a trade-off between parsimony and omission of relevant variables, which can give rise to *nonfundamentalness* of the identified shocks (see, e.g., Forni et al., 2009). Extracting information from a large set of macroeconomic variables overcomes the limited information problem because the principal components proxy the unobserved factors affecting most macroeconomic variables (see Fragetta and Gasteiger, 2014 for further details). Similar to Bernanke et al. (2005), we implement a two-step estimation procedure. As a first step, we extract five common factors, as established by the Bai and Ng (2007) $IC_{p,2}$ information criterion. The second step is adding the five factors to our vector of endogenous variables, which reads as follows:

$$y_{i,t} = [G_{i,t}, GDP_{i,t}, T_{i,t}, F_t]'$$

where $F_t$ is a $1 \times 5$ vector common to all countries, but that may have a different impact in each country, capturing also potential spillovers across countries.

Third, among the exogenous variables, we add the $f_{(t)}|_{t-1:t-4}$ series. This represents the forecast of time-$t$ government spending over the past 12 months (four quarters), published by the Economist Intelligence Unit. The addition of this variable represents a way to purge our structural government spending shocks from the change in government spending already anticipated by economic agents, and hence to solve the problem known in the literature as *fiscal foresight*. Fiscal foresight is the phenomenon by which private agents, mainly due to legislative and implementation lags, can anticipate future movements in government spending. Failing to account for them in the identification of what are meant to be unanticipated
government spending shocks may give rise to endogeneity and bias the results (see, e.g., Forni and Gambetti, 2010 and Leeper et al., 2013 among others for further details).

Fourth, we use as interaction term, $x_t$, the European Central Bank’s shadow monetary policy rate developed by Wu and Xia (2017) (discussed in Section 1), which allows us to control for the overall monetary policy stance in the eurozone, and study the effects of a government spending shock at each percentile of the shadow rate distribution. Since this rate is available from 2004Q3 onward, for the very beginning of the sample, we complement it with the Main Refinancing Operations (MRO) rate, given that the two, until 2008, are virtually indistinguishable. To avoid potential reversed causality issues, we use the first lag of the shadow rate (i.e. $x_t = sr_{t-1}$), such that it is predetermined relative to the endogenous variables. It must be said, however, that the five factors do contain information on the contemporaneous monetary policy stance although, for our purposes, there is no need to explicitly identify a monetary policy shock.

Lastly, in order to account for international factors which may influence our variables of interest, we add as exogenous variables also a set of U.S. variables, $z_{t-1}$, including the U.S. output gap, U.S. inflation and the U.S. shadow monetary policy rate developed by Wu and Xia (2016).

All the abovementioned modifications made to the traditional fiscal VAR specification are implemented within the IPVAR model. Therefore we label the model used in this paper as a factor-augmented interacted panel vector-autoregressive model purified of expectations (FAIPVAR-X).

### 2.3 Inference, Identification and Computation of Cumulated Government Spending Multipliers

In line with Sá et al. (2014), we estimate the FAIPVAR-X model presented in equation (1) and compute cumulated government spending multipliers adopting the following eight steps:

1. Estimate the structural model equation by equation using ordinary least squares (OLS) and adopt a Bayesian strategy for inference utilizing an uninformative independent Normal–Wishart prior, which in turn uses a Montecarlo simulation to recover the posterior distribution of the structural parameters.

2. Make a draw of the posterior distribution and evaluate it at pre-specified values of the interaction term $x_t$.

3. Derive the model’s corresponding reduced form, by pre-multiplying equation (1) by $B_{i,t}^{-1}$. 


Table 1: Sign Restrictions for Identifying the Government Spending Shock

<table>
<thead>
<tr>
<th>Variable</th>
<th>Sign</th>
<th>Periods</th>
</tr>
</thead>
<tbody>
<tr>
<td>$G_{it}$</td>
<td>+</td>
<td>4</td>
</tr>
<tr>
<td>$\text{GDP}_{it}$</td>
<td>+</td>
<td>4</td>
</tr>
<tr>
<td>$T_{it}$</td>
<td>None</td>
<td>None</td>
</tr>
<tr>
<td>$F_t$</td>
<td>None</td>
<td>None</td>
</tr>
</tbody>
</table>

4. The reduced form and the covariance matrix may depend on the order given in the structural recursive form. Given legislative and implementation lags in fiscal policymaking, order government spending as the first variable using standard practice dating back to Blanchard and Perotti (2002). Note that changes in the order of the explicit variables (not in the factors) do not materially change the results.

5. Use a sign restriction strategy to identify an unexpected government spending shock and compute the resulting impulse response functions (IRFs). More specifically, follow the same procedure of Sá et al. (2014), which in turn uses the algorithm developed by Rubio-Ramírez et al. (2010): after defining $V_x^d$ as the Cholesky decomposition of the reduced form variance-covariance matrix $\Sigma_x^d$, draw an orthonormal matrix $Q$ such that $Q'Q = I$, from which it follows that $B^d = V_x^d Q$ and $\Sigma_x^d = B^{dt} B^d = V_x^{dt} Q' Q V_x^d$ where $d$ indicates a stable draw from the posterior distributions. To achieve identification, the impulse responses implied by $B^d$ have to satisfy the following restrictions: a government spending shock should raise $\text{GDP}_{it}$ and $G_{it}$ for at least four quarters (Table 1).

6. For every 100 draws of the $Q$ matrix satisfying the sign restrictions, save its median value.

7. Make 20000 draws from the posterior distribution and use the median over the 10000 medians as the central estimate of interest.

8. Compute cumulated government spending multipliers following the approach proposed by Gordon and Krenn (2010) and Ramey and Zubairy (2018). As already discussed, having normalized the variables of interest by real potential GDP, circumvents any concerns related to ex-post conversion. Thus, cumulated multipliers are computed

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3 As in Cogley and Sargent (2005); Primiceri (2005); Sá et al. (2014), we discard any explosive draws from the unrestricted posterior.

4 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.

5 Note that we consider the first 10000 parameter draws as burn-in draws.
simply as the ratio of discrete approximations of the integral of the median IRFs of real output and government purchases over a given time horizon \( h = 0, 1, \ldots, H \):

\[
\mathcal{M} = \frac{\sum_{h=0}^{H} dGDP(h)}{\sum_{h=0}^{H} dG(h)}. \tag{4}
\]

In the remainder of the paper, we use a time horizon of five years (\( H = 20 \)), as this conventionally represents the medium term. We account for parameter uncertainty by saving the 16th and 84th percentile of the distribution of the median as error bands.\(^6\)

3 Results

In this section, first we report impulse response functions (IRFs) of government spending, output, and net taxes—all in real terms—to an unexpected shock to government spending. One of the advantages of the FAIPVAR-X model is that it allows conditioning the IRFs on a specific percentile of the distribution of the shadow interest rate. In other words, we can interpret the IRFs as the dynamic reaction of macroeconomic variables to a shock to government spending occurring when the shadow rate takes a value corresponding to a given percentile of its own distribution.

For expository ease we report IRFs conditional on two percentiles that are representative of two euro-area monetary policy regimes. We label the first regime normal times. This regime corresponds to the period between the beginning of our sample (2002q2) and the bankruptcy of Lehman Brothers (2008q3). In this period the shadow rate almost coincided with the official Eonia rate and the two were clearly positive (see Figure 1). We label the second regime ELB. This regime is comprised between 2012q4, the quarter following ECB President Mario Draghi’s famous ‘Whatever it takes’ speech, to the end of the sample (2017q4). During this period, the ECB lowered the monetary policy rate first to the ZLB and then to negative values. In other words, this period is characterized by a binding but time-varying ELB. The systematically negative shadow rate captures a series of unconventional measures including the Asset Purchase Program (APP) and forward guidance.\(^7\)

\(^6\)We can distinguish between identification uncertainty and parameter uncertainty: the former reflects the lack of information we have about the properties of the structural shock; the latter accounts for the limited amount of data. While our identification strategy entails both identification and parameter uncertainty, we only show parameter uncertainty. It can be shown to be sufficient to establish whether there are differences among states (for further details see Paustian (2007); Sá et al. (2014))

\(^7\)We do not report IRFs for the intermediate period (2008q4-2012q3) as it is a hybrid period in which the monetary policy rate was quickly lowered but did not reach the ELB, while the shadow rate started to depart from the Eonia rate and fluctuated around zero, crossing the zero line three times. However, below we report cumulated government spending multipliers also for the intermediate period. IRFs (available upon
The choice of regimes translates into a number of choices as far as shadow rate percentiles are concerned. We pick the 77th percentile (2.75 percent; 2003q2) as a representative of the normal times regime, being the closest to the average shadow rate for that period (2.82 percent); and the 16th percentile (-2.23 percent; 2015q3) as a representative of the ELB regime, again being the closest to the average shadow rate for that period (-2.29 percent).

Figure 2 contrasts the IRFs for the normal times regime with those for the ELB regime. A few remarks are in order. First, in all cases a shock to government spending keeps spending itself persistently above baseline and it takes about ten quarters to die out. Second, output and tax revenues respond positively and significantly to the shock. Third, comparing the results for normal times against those for ELB unveils that, when the economy is at the ELB, the responses of output are for the most part much larger, and the difference is statistically significant in many quarters after the occurrence of the shock (quarters 0-1 and 7-16). The responses of net taxes also differ across the two regimes, but to a lesser extent relative to those of output.

With the FAIPVAR-X model, IRFs can be computed conditional on virtually any level of the shadow rate in the sample, which can be easily reconducted to a specific quarter in the history of the EMU. This allows us to compute the cumulated government spending multipliers explained in Subsection 2.3. Table 2 reports the level of the shadow rate and the multiplier for each feasible quarter.\footnote{There were eight levels of the shadow rate for which the FAIPVAR-X procedure did not yield stable IRFs, therefore we discarded them from the analysis.}

Figure 3 allows visually inspecting this pattern. An inverse association between the shadow rate and the size of the cumulated government spending multiplier is easily detectable, with a correlation coefficient between the two of -0.82. In normal times, characterized by a clearly positive shadow rate, the cumulated government spending multiplier was on average 0.88; in the ELB regime, characterized by a binding ELB and a negative shadow rate, the multiplier was on average 3.39; while in the intermediate period, when the shadow rate fluctuated around zero, the average multiplier was 2.54. In other words, when the EA monetary policy rate was at the ELB, the government spending multiplier was, on average, almost four times higher than in normal times. These results are in line with a strand of the theoretical (DSGE) literature that claim the fiscal multiplier to be much higher at the ZLB (Christiano et al., 2011; Coenen et al., 2012; Davig and Leeper, 2011; Eggertsson, 2010; Kilponen et al., 2015; Woodford, 2011, among others). However, it has to be stressed that while in DSGE models it is possible to quantify the effects of the ZLB in isolation, in the data this task is much more difficult. In fact, when the policy rate reached the ELB, mon-
Figure 2: Impulse Responses to a Government Spending Shock in Normal Times and at the ELB

Notes: Impulse responses in percent to a shock of size one standard deviation. Bold lines represent median responses. Shadowed areas and dashed lines represent 16th and 84th percentile confidence bands.

Economic policy did not simply cease to operate; it continued to function in an unconventional manner. By conditioning the computation of the multiplier on the shadow monetary policy rate, we simultaneously capture not only the effects of the time-varying ELB in the EA, but
also those of all unconventional monetary policies. It must be said, however, that at the beginning of the ELB regime (before the start of QE in 2015q1), multipliers were already much higher than in normal times (2.90 on average), suggesting a role for the ELB independent of QE.

There is a strand of the literature that finds government spending multipliers to be dependent on the business cycle in advanced economies, and in particular to be higher in recessions relative to expansions (see Auerbach and Gorodnichenko, 2012, 2013; Batini et al., 2012, among others).\(^9\) If there were a perfect correlation between the shadow rate and the EA business cycle, then one possibility would be our results to be driven by the state of the business cycle, rather than by monetary policy. Figure 4, however, shows that the business cycle indicator for the countries in our sample (see Appendix A for its construction) is uncorrelated with the shadow rate (correlation coefficient of -0.02).

We also conduct a more formal check, that is, we run a simple regression of the multiplier on the business cycle indicator and compute the correlation between the residuals of this

\(^9\)Ramey and Zubairy (2018) do not confirm this result for the U.S. using a longer times series when they construct the fiscal multipliers as we do (see Subsection 2.3).
regression and the shadow rate. This amounts to -0.84 and it is statistically different from zero at a 1 percent level, giving further assurance that the role that the shadow rate has on the size of the government spending multiplier is autonomous from that of the business cycle. Nonetheless, EA data still give support to the view that the multiplier is larger in periods of economic slack. We arrive at this conclusion by running a second regression of the multiplier on a constant and the lagged shadow rate, and by computing the correlation between the residuals of this regression and the business cycle indicator. The correlation coefficient amounts to -0.36 and it is statistically different from zero at a 1 percent level. In other words, after controlling for the level of the shadow rate, the multiplier is still negatively correlated with the business cycle. In sum, both the shadow rate and the state of the business cycle have an autonomous impact on the size of the fiscal multiplier in the eurozone.
4 Robustness Checks

In this section we present the results of three robustness checks addressing issues commonly discussed in the literature, which may be applicable also to the analysis presented in this paper:

1. Impulse response function à la Fry and Pagan (2011). The method we used to derive IRFs closely follows Sá et al. (2014) and is widespread in the literature. However, this has been recently criticized by Fry and Pagan (2011), who claim that considering the median response as the point estimate of the exactly identified model may be inaccurate. Therefore, we deem appropriate to check if our results are robust to the use of the median target approach, a different IRFs computation procedure proposed by Fry and Pagan (2011) themselves. In practice, relative to the procedure used to obtain the baseline results, this procedure differ in that, out of the 100 rotation matrices $Q$ that satisfy the sign restrictions, we save that implying the model with the impulse
response functions closest to the median IRFs. Then, out of all collected models, for every draw of the posterior distribution, we consider again only the model producing the IRFs nearest to the median IRFs.

2. **Sign restrictions imposed for two periods.** To produce the baseline results, we applied positive sign restrictions to the IRFs of real GDP and government spending for 4 quarters, with the purpose of identifying a temporary government spending shock, given that governments’ budgets normally cover a year. Given that the literature has shown that results may be sensitive to the choice of number of quarters for which restrictions are imposed, we check the robustness of our results to imposing the minimum set of restrictions useful to reach identification. That is, we restrict the IRFs of real GDP and government spending to be positive for at least two quarters. Note that if we were to restrict the IRFs to be positive for at least one quarter, we would face the so-called multiple shocks problem (Fry and Pagan, 2011). This problem arises when the econometrician does not include enough information to discriminate across shocks.

3. **Lag structure of two quarters.** Given that the FAIPVAR-X model requires the estimation of a large number of parameters, for the sake of parsimony, we produce the baseline results with a uniform lag structure of one quarter. Bearing in mind that the use of a very long lag structure would not be feasible as we would run out of degrees of freedom, we check whether results are robust to the use of a lag structure of two quarters ($L = 2$).

As shown in Figure 5, these robustness checks, to a an extent, mostly dampen the average five-year cumulated government spending multipliers, but in some cases they also magnify them. Importantly, all the conclusions drawn from the analysis of the baseline results survive these checks. First, the negative association between the level of the shadow monetary policy rate and the size of the multiplier is still clearly detectable (with the correlation coefficients ranging between -0.82 and -0.65). Next, the average multiplier during the ELB regime is still much higher than in normal times regardless of the specification. Table 3 compares the average five-year cumulated government spending multipliers between the baseline and the alternative specifications in normal times and during the ELB regime. Although the average multipliers change across specifications, the ELB regime has multipliers 2.3 to 4.4 times higher than normal times. In fact, while in normal times the average multiplier ranges between 0.9 and 1.2; in the ELB period, the range is between 2.5 and 3.4. Lastly, running the auxiliary regressions described at the end of Section 3 confirms that, in all cases, the role that the shadow rate has on the size of the government spending multiplier is autonomous.
Figure 5: Five-Year Cumulated Government Spending Multipliers and the Shadow Monetary Policy Rate in the Euro Area: Robustness Checks

(a) Impulse response functions *à la* Fry and Pagan (2011)

(b) Sign restrictions imposed for two periods

(c) Lag structure of two quarters
Table 3: Average Five-Year Cumulated Government Spending Multipliers in the Euro Area: Robustness Checks

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<th>Average Five-Year Cumulated Government Spending Multipliers</th>
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from that of the business cycle and that, nevertheless, EA data give support to the view that the multiplier is larger in periods of economic slack.

5 Conclusions

Since the 1930s, especially in periods of low aggregate demand, the fiscal multiplier has been at the forefront of the macroeconomic debate. Policymakers always face the practical challenge of having to make assumptions on fiscal multipliers when designing macroeconomic adjustment scenarios. Academics have often questioned whether the multiplier is indeed a multiplier, that is, whether it is greater or smaller than one. Therefore, many theoretical and empirical contributions proposed models and empirical strategies to estimate the size of the multiplier. The economic profession is still far from achieving a consensus on actual estimates. However, there is at least a convergence on the idea that the size of the multiplier heavily depends on the macroeconomic policies prevailing when fiscal measures are implemented.

This paper focuses on one aspect of macroeconomic policy that became one of the defining features of the period that followed the GFC in most advanced economies: the ELB. Because of the ELB, many argued that fiscal authorities needed to do more to provide the necessary stimulus to mitigate output and job losses. Central banks, on their part, had to become creative in how to conduct monetary policy, often exploring unchartered territories. Macroeconomists started to argue that the size of the fiscal multiplier would likely be different when the economy is at the ELB. This assertion was initially only circumstantiated using theoretical model, as unfortunately there were not enough data points to conduct formal empirical estimations.
Ten years after the beginning of the GFC, both data and econometric tools allow us to explore this question empirically. We focus on the EA because, besides being one of the largest advanced economy with no available estimates of the fiscal multiplier at the ELB, it is a very interesting laboratory for this research question. The ECB made the ELB time-varying, pushing the monetary policy rate into negative territory and then complemented it with quantitative easing and forward guidance. In addition, the eurozone experienced both fiscal stimulus measures following the GFC, and austerity to deal with the sovereign debt crisis, in sum substantial discretionary fiscal shocks.

Cognizant of the challenges of isolating the effect of the ELB on the computation of the fiscal multiplier, we prefer to take an approach that controls for the overall monetary policy stance, including the time-varying ELB. In other words, we condition the calculation of the government spending multiplier on a shadow monetary policy rate. The most recent literature highlighted many econometric problems that should be taken into consideration when dealing with fiscal multipliers. We use these lessons to devise our econometric approach, the FAIPVAR-X model, which tackles problems of limited information and fiscal foresight, among other matters.

Our results agree with those theoretical predictions pointing to a larger size of the government spending multiplier at the ELB versus so-called normal times. Following the most recent literature, we compute cumulated multipliers for the medium run (five years). Depending on the specification, we find that while in normal times the average multiplier ranges between 0.9 and 1.2, in the ELB period, the range is between 2.5 and 3.4. The FAIPVAR-X model allows us to compute a multiplier for virtually all percentiles of the shadow rate. It turns out that the multiplier is inversely correlated with the level of the shadow monetary policy rate, suggesting a clear impact of the monetary stance on the multiplier itself. Lastly, we verify that EA data are in line with the view that the multiplier is larger in periods of economic slack. Importantly, however, we show that the shadow rate and the state of the business cycle have an autonomous impact on its size.

These results are important at least for two reasons. First, using data of a large advanced economy, they add empirical backing to a strand of theoretical contributions. The second reason is perhaps more important from a policy perspective. At a time of phasing out of the ECB’s asset purchase program and prospective normalization of monetary policy, this paper may provide an educated guess of the multiplier policymakers in the EA should use in the forthcoming normal times. Our estimates suggest a value around one.
References


Appendix

A Data

A.1 Endogenous Variables

Our variables of interest are gross domestic product, net taxes and government spending. As standard in the literature we construct net taxes as the sum of government receipts of direct and indirect taxes minus transfers to businesses and individuals. The government spending series is constructed as the sum of government gross fixed capital formation and government consumption. All the variables are downloaded from the Eurostat database available on the Thomson Reuters Datastream Economics database and are transformed in real terms using the implicit GDP price deflator. Then they are normalized by diving by real potential GDP.

A.2 Exogenous Variables

We use as exogenous variables the forecast of the annualized growth rate of total government expenditure over GDP produced by the Economist Intelligence Unit. In particular, we create a series where in each quarter we compute the average forecast for the current year over the past 12 months. Our interaction term is the European Central Bank’s shadow rate developed by Wu and Xia (2017). The other exogenous variables are the U.S. output gap, and the U.S. inflation downloaded from the Federal Reserve Bank of St. Louis database, and the U.S. Shadow Rate developed by Wu and Xia (2016).

A.3 Informational Dataset

The informational dataset we used to extract common factors is composed by 250 series downloaded from the Eurostat database available on the Thomson Reuters Datastream Economics database. Specifically we downloaded the following variables for each country considered: Change in Inventories; Domestic Demand; Early Estimates of Labor Productivity - Total Economy; Employees Domestic Concept; Export of Goods and Services; Final Consumption Expenditure of Households; Government Consolidated Gross Debt: Central Govt; Gross Capital Formation; Harmonized Government 10-Year Bond Yield; Household and NPISH FCE; Imports of Goods and Services; Industrial Production Index: Manufacturing; Industrial Production Index: Mig-Intermediate Goods; Money Supply: M1 - Contribution to Euro M1; Money Supply: M2 - Contribution to Euro M2; Money Supply: M3 - Contribution to Euro M3; NEER: 28 Trading Partners; NEER: 37 Trading Partners; Nominal Unit Labor
Cost based on persons; Official Reserve Assets; Production - Total Industry Excl. Construction; Production of Total Construction; S&P BMI - Price Index; Unemployment: Total; Wages and Salaries. The following series are converted from monthly to quarterly frequency: Harmonized Government 10-Year Bond Yield; Industrial Production Index: Manufacturing; Industrial Production Index: Mid-Intermediate Goods; Money Supply: M1 - Contribution to Euro M1; Money Supply: M2 - Contribution to Euro M2; Money Supply: M3 - Contribution to Euro M3; Official Reserve Assets. Moreover, the S&P BMI - Price Index series is converted from daily to quarterly frequency. Where appropriate we transform variables to guarantee stationarity tested by the Dickey and Fuller (1979) and Kwiatkowski et al. (1992) tests.

A.4 Business Cycle Indicator

Our aggregate indicator of business cycle is computed as \( \sum_{i=1}^{10} \frac{GDP_i}{GDP_i^{trend}} \), where \( i \) are the countries considered in our sample, and \( GDP_i, GDP_i^{trend} \) are the real GDP and the real potential GDP of each country, respectively. In general, if our aggregate indicator is large and positive (large and negative), it means that most countries are in expansion (recession).
### Extended Results of Robustness Checks

Table B.1: IRFs à la Fry and Pagan (2011): Five-Year Cumulated Government Spending Multipliers

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Table B.2: Sign Restrictions Imposed for Two periods: Five-Year Cumulated Government Spending Multipliers

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Table B.3: Lag Structure of Two Quarters: Five-Year Cumulated Government Spending Multipliers

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