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Three Essays in Macroeconometrics

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Abstract

This dissertation comprises three self-contained chapters in macroeconometrics tackling three current macroeconomic issues.

The first chapter, “Economic Policy Uncertainty in the US: Does it matter for the Euro Area", is a single-authored paper. It studies to what extent an economic policy uncertainty shock generated in the US triggers spillover effects in the Euro Area macroeconomic activity. I estimate a two-country Structural Vector Autoregressive (SVAR) model capturing the US economic policy uncertainty shock by appealing to some indicators recently proposed by Baker, Bloom and Davis (2013). The impulse responses predict a negative and statistically significant reaction of Euro area price and quantity indicators to an unexpected increase in the US policy uncertainty. The results support the view that the effects of such shock act like a demand "type" shock. Interestingly, the Euro area variables are estimated to respond strongly to US uncertainty shock than to the European counterpart. A note from this chapter has been published in the *Economics Letters*.

The second chapter, "Estimating Fiscal Multipliers: News from a non-linear Word", is a joint work with Giovanni Caggiano, Efrem Castelnuovo and Gabriela Nodari. We estimate the effects of a US government spending news shock on output multipliers. We deal with the issue of fiscal foresight (anticipated fiscal policy changes) by appealing to the sums of revisions of expectations on future government spending. This measure is used to add more information to a standard three-variate fiscal VAR model à la Blanchard and Perotti (2002). To study the effects of anticipated fiscal policy shocks conditionally on the state of economy (in recessions and expansions), we estimate a non-linear VAR (Smooth Transition VAR). We

compute non-linear generalized impulse responses (GIRFs) à la Koop, Pesaran, and Potter (1996) to take into account the probability of smoothly switching from a regime to another due to the fiscal shock. Results show that an anticipated fiscal spending shock triggers a significant reaction of GDP. Such reaction is not significantly different between recessions and expansions. However, when we discriminate between "extreme events" (i.e., the recent great recession), we find a reaction of output significantly different among regimes. Fiscal multipliers in recessions are estimated above one and government spending news shocks are found to induce economic stabilization. This chapter is forthcoming in the *Economic Journal*.

The third chapter, "Opening the Red Budget Box: Real Effects of a Tax Shock in the UK", is a single-authored paper. It studies non-linear impulse responses to unanticipated tax shocks for output and its components. To do so, I estimate a non-linear version of the local projection technique developed by Jordá (2005). The identification of the tax shock is achieved by appealing to the measure of exogenous tax changes in the UK developed by Cloyne (2013). Tax shocks are found to affect UK macroeconomic variables depending when such shocks occur (in recessions or in expansions). An unexpected increase in the tax rate occurring in recessions triggers a large, persistent and negative reaction in output, consumption, investment, imports and government consumption. The results suggest that output tax multipliers are negative and above one (in absolute value) in recessions but not in expansions.The size and the sign of multipliers for most of the macroeconomic indicators above considered are also found to be state-dependent.

Acknowledgements

I was very lucky to have Efrem Castelnuovo as supervisor. I would like to express my gratitude to him. His experience and knowledge have allowed me to deal with "hot" macroeconomic issues. I benefited from his ideas, comments, suggestions and insightful discussion. Without his patience, guidance and encouragement this work could not have been materialized.

I am grateful to my external supervisor Timo Teräsvirta. Our weekly meeting during my visiting period at the CREATES have strengthened my knowledge on non-linear techniques. I will never forget his "you don't know if you don't try".

Special thanks go to Giovanni Caggiano whose suggestions and comments have improved my research. Thanks to the professors of my pre-defence committee and, in particular, to Giorgio Brunello for his comments and for his availability as Head of the PhD program.

I would like to thank my colleagues with whom we shared long days of study, concerns and satisfactions, but also non-academic time.

Last but not least I grateful to the patience and understanding of my boyfriend and to the support of my family.

Contents

List of Figures

List of Tables

Chapter 1

Economic policy uncertainty in the US: Does it matter for the Euro Area?

Abstract

We investigate the effects of a US economic policy uncertainty shock on some Euro area macroeconomic aggregates via Structural VARs. We model the indicators of economic policy uncertainty recently developed by Baker, Bloom, and Davis (2013) jointly with the aggregate price indexes and alternative indicators of the business cycle for the two above indicated economic areas. According to our SVARs, a one standard deviation shock to US economic policy uncertainty leads to a statistically significant fall in the European industrial production and prices of -0.12% and -0.06%, respectively. The contribution of the US uncertainty shock on the European aggregates is shown to be quantitatively larger than the one exerted by an Euro area-specific uncertainty shock.

1.1 Introduction

The attention on the macroeconomic effects of uncertainty has been recently reignited by Bloom (2009)'s highly influential paper. A number of VAR investigations have been proposed to quantify the impact of uncertainty shocks at a macroeconomic level, (see e.g., Alexopoulos and Cohen, 2009; Bloom, 2009; Baker, Bloom, and Davis, 2013; Caggiano, Castelnuovo, and Groshenny, 2013; Leduc and Liu, 2013; Nodari, 2014). Such investigations have typically followed a "within-the-US-country approach", i.e., they have focused on the reaction of a set of US variables to a shock to the level of uncertainty affecting the US economy itself. While being a somewhat natural approach, shocks hitting a leading economy such as the United States may very well spillover onto other countries. Investigations documenting the existence of spillovers include Kim (2001), who quantified the role of US macroeconomic shocks in triggering business cycles at an international level, and Favero and Giavazzi (2008) and Ehrmann and Fratzscher (2009), who look at spillover effects regarding financial markets. As for the literature dealing with uncertainty shocks, Mumtaz and Theodoridis (2012) estimate an open-economy VAR focusing on the potential impact of the volatility of shocks to US real activity on UK. They find that spillovers across these two areas may very well be important.

This paper asks the following question: *"Are there spillovers from the US economy to the Euro area due to economic policy uncertainty shocks?"* To answer this question, we model a VAR including both US and Euro area aggregates. Then, we identify a US uncertainty shock via the imposition of short-run restrictions, and focus on the responses of Euro area prices and quantities. The uncertainty shock is identified by appealing to the "economic policy uncertainty indicator" recently developed by Baker, Bloom, and Davis (2013). *The answer* provided by our empirical investigation *turns out to be positive:* a one-standard deviation shock to US economic policy uncertainty leads in the short-run to a statistically significant fall in the European industrial production and prices of -0.12% and -0.06%, respectively.

Our paper is structured as follows. Section 2 focuses on the data and the identification scheme employed in our VAR-approach. Section 3 presents our results. Section 4 concludes.

1.2 Data definition and VAR specification

We analyze the transmission of structural shock from the US to Euro area within a two-country Structural Vector Autoregressive model (SVAR). A common representation of the SVAR is:

$$B_0 y_t = B(L) y_{t-p} + \varepsilon_t \tag{1.1}$$

where $B(L)$ is an autoregressive lag-polynomia, and ε_t is the vector of structural innovations. The vector $y_t = [CPI^{US}\ IPI^{US}\ i^{US}\ News^{US}\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}\ News^{Euro}]'$ includes all the endogenous variables in our model and relies on two blocks: the first one refers to "foreign" variables (US), whereas the second one includes "domestic" variables (Euro area). Each regional block includes: the consumer price index (CPI for the US and HCPI for the Euro area), as measure of prices; the industrial production index (IPI), as a proxy for the business cycle; the short-run interest rate (indicated with "i" in the vector above), which is the Federal Funds Rate for the US and the three-month interest rate for the Euro area, as a proxy for the monetary policy instrument. To account for economic policy uncertainty in the US and the Euro area, we employ two country-specific empirical proxies carefully constructed by Baker, Bloom, and Davis (2013). The policy-related economic uncertainty for the US (EPU^{US}) relies on three components: a news-based component quantifying newspaper coverage on economic policy uncertainty ($News^{US}$); a measure of the federal tax code provisions; and a measure of disagreement among forecasters. The Euro area uncertainty index (EPU^{Euro}) relies on two components: a news-based component ($News^{Euro}$), and a measure of disagreement among forecasters. Since the overall economic policy uncertainty indexes rely on different components, we focus on uncertainty indexes based on news coverage. The correlation between the EPU indicator and its news-based component is 0.97 and 0.93 for the US and Euro area, respectively. Hence, we include in vector y_t the news-based components, $News^{US}$ and $News^{Euro}$, as proxies for the economic policy uncertainty.[1] Figure 1.1 plots the monthly time series of the overall uncertainty indexes and news components, both for the US and the Euro area.

We need to recover the structural shocks ε_t from $\varepsilon_t = B_0 u_t$, where B_0 contains the contemporaneous relationships between the reduced-form residuals u_t and the structural shocks ε_t. To identify B_0, we employ a standard Cholesky decomposition imposing a lower triangular matrix. Since we are interested in the effects of an external policy uncertainty shock (US) on the domestic macroeconomic variables (Euro area), we impose short-run restriction following a country-based exogenous approach. Because we are using a Cholesky decomposition, the ordering of the variables in our vector y_t is important. Following Favero and Giavazzi (2008), we assume that shocks hitting the Euro area exert no contemporaneous effects on the US variables. Consequently, the US block is ordered before the Euro area block in our vector. Second, within each country-block, we order uncertainty last. We do so to "purge" the uncertainty indicator in our VAR from the contemporaneous movements of our macroeconomic indicators (prices, industrial production), therefore sharpening the identification of uncertainty shocks.

Our data are monthly and span the period 1999M1 to 2008M6. The beginning of the period is motivated by the creation of the Euro area, whereas the end is chosen to avoid possible non-linearities due to the intensification of the financial crisis. All variables are in log-levels, except for the interest rate and the uncertainty indexes, which are in levels.[2] We select the optimal number of lags in the SVAR model combining an initial lag selection based on information criteria with an LMF test for no serial correlation in the error terms.[3] Our SVAR(3) includes

[1] Our results are robust to the use of the overall indexes instead of their news components.

[2] Sims, Stock, and Watson (1990) show that VARs in log-levels provide consistent estimates of the IRFs even in presence of co-integrating vectors. We do not attempt to model co-integrating vectors given the small size of our sample.

[3] SIC and BIC information criteria suggest a VAR(1), whereas AIC a VAR(2). However, the results are robust to different lag-length choices.

equation-specific constants and linear trends. The data have been retrieved from the Federal Reserve Bank of St. Louis' database (US industrial production, price level, and federal funds rate), the European Central Bank's Statistical Warehouse (industrial production, price level, and the three-month interest rate), and the "Economic Policy Uncertainty" website (http://www.policyuncertainty.com/).

1.3 Results

Figure 1.2 depicts the impulse response functions to a one-standard deviation shock to the US uncertainty index. The responses of US industrial production and consumer price index are statistically significant, and suggest a decline in production and a deflationary phase after an increase in uncertainty. Both the industrial production and prices hit their lowest values after three months, reaching a minimum around -0.13% and -0.08%. The Federal Reserve reacts fast to the economic condition by adopting an expansionary monetary policy. As the economy settles on the recovery path, the interest rate goes back to its steady state. Our results corroborate those reported in previous contributions on the "demand" type of effects triggered by uncertainty shocks in the US economy (Alexopoulos and Cohen, 2009; Bloom, 2009; Baker, Bloom, and Davis, 2013; Caggiano, Castelnuovo, and Groshenny, 2013; Leduc and Liu, 2013; Nodari, 2014).

Moving to our research question, our VAR predicts *a negative and significant reaction of Euro area price and quantity indicators to an unexpected increase in the US policy uncertainty*. The industrial production and consumer prices drop to -0.12% and -0.06%, respectively, two months after the shock. Then, they slowly go back to their pre-shock level. One possible explanation is that increases in uncertainty lead both households and firms to postpone their consumption and investment decisions due to a precautionary saving-motive (the former) and an increase of the option-value of waiting (the latter). The fall in aggregate demand may be responsible for the temporary deflation predicted by our VARs. The monetary policy easing associated to a temporary reduction in the nominal interest rate is consistent with an inflation-targeting strategy pursued by the monetary policymakers.[4] Notably, our impulse responses suggest that, following an exogenous increase in the US economic policy uncertainty, the Euro area-related uncertainty also increases. Obviously, given the high level of contamination involving the US and the Euro area at commercial and financial levels, policy (in)decisions in the United States may very well increase the perceived uncertainty surrounding policy moves in Europe. Admittedly, our VARs do not distinguish between reactions by European aggregates due to an increase in the US uncertainty *per se* vs. reactions to an increase in the endogenous component of the Euro-area related uncertainty. This, however, does not affect our main message, i.e., US economic policy uncertainty shocks exert a significant effect on Euro area macroeconomic aggregates.

How important is a US uncertainty shock? Table 1.1 highlights the contribution of the US and European policy uncertainty shocks in explaining the short-run fluctuation in the European variables. In the short-run, the Euro area variables are estimated to respond more strongly to US uncertainty shock than to the European counterpart. At a six month horizon, the US shock explains 4% of the variation in the European industrial production whereas the European policy uncertainty accounts for 2%. The change in the European consumer prices and policy rate in response to a US uncertainty shock is six times larger than under the European counterpart. Therefore, the US policy shock explains an appreciable share of the variance of the forecast error of the Euro area variables (above all, the policy rate). More importantly, such shock appears to be more relevant on European aggregates than its European counterpart. Table 1 also reports the results obtained by estimating the impact of US uncertainty shocks with the two alternative proxies for uncertainty that compose the Economic Forecast Disagreement recently proposed

[4] Our results are robust to: i) ordering the news indexes first in each country-specific block; ii) different lag-length specifications; iii) the introduction of extra-variables in the VAR (i.e., nominal effective exchange rate, Chicago Fed National Activity Index and EuroCoin business cycle indicator, University of Michigan Consumer Sentiment Index); iv) the employment of alternative uncertainty indexes (EPU^{US}/EPU^{Euro} and $VIX/VSTOXX$); v) the inclusion of the financial crisis period in our sample. The robustness checks are available upon request.

by Baker, Bloom, and Davis (2013): the Government Spending Disagreement Forecast, and the CPI Disagreement Forecasts (CPIDF).[5] The GSDF proxy confirms the relatively larger role played by US uncertainty shocks on European variables as for industrial production and the policy rate. The CPIDF measure of uncertainty plays a milder role for both US and European uncertainty shocks, therefore suggesting that different measures of uncertainty may very well depict different contributions as for the macroeconomic dynamics of the Euro area. Finally, Table 1 (see Sample with Financial Crisis observations, line SFC) documents the reduction of the relevance of US uncertainty shocks (in the context of our baseline model), possibly due to the increased variability in the policy uncertainty index.

1.4 Robustness checks

Our results show that an expected increase in the US economic policy uncertainty has negative and statistically significant effects on the Euro Area macroeconomic aggregates. In this section, a number of robustness checks are considered.

Contemporaneous effects of economic policy uncertainty shocks. We identify the economic policy uncertainty shocks ordering last the uncertainty indicators in each country-block. This specification allows us to "purge" the uncertainty index by others US shocks that simultaneously may hit the US variables. However, such specification imposes on impact a zero-reaction of US macroeconomic aggregates to an increase in the US economic uncertainty. To check the extent to which our Cholesky identification assumption may affect the results,[6] we estimate an alternative specification in which the policy uncertainty indicators are ordered first in each country-block. This specification implies that the US policy uncertainty shock is predetermined with respect to the other US macroeconomic

[5] The US government spending disagreement forecast refers to the federal, state, and local purchases for the US, whereas the European one only concerns to the federal budget balances.

[6] In our model the innovation is orthogonized using the Cholesky decomposition of the covariance matrix, then it may be sensitive to the ordering of the variables. Choosing a different order of the variables produce different shocks and the impact of the shock on the system may depend on the way the variables are setting in vector y_t. To tackle this problem Sims (1980) suggested checking the robustness of the results to the ordering of the variables.

variables. Figure 1.3 displays the results. While in our baseline specification the US uncertainty shock acts like a demand shock on the US macroeconomic variables, such result is reversed when the uncertainty indicator is ordered first. These mixed results for the US economy are in line with the doubts arising from the recent economic literature about uncertainty. Indeed, Leduc and Liu (2013) find that a uncertainty shock acts as demand shock in the US. Conversely, Mumtaz and Theodoridis (2012), studying the spillover effect due to a volatility shock in an open economy, find that such shock acts as a supply shock. Turning to the Euro Area, this alternative specification confirms our main results. An increase in the US economic policy uncertainty has negative and statistically significant effects on the Euro Area.

Additional information. One potential concern with our baseline specification is that vector y_t does not embed sufficient information. If the SVAR is misspecified, then the IRFs may be distorted. To address this concern, we (re)estimate our SVAR including additional macroeconomic variables. To incorporate international linkages between the US and the Euro Area, we add to the vector y_t the nominal effective exchange rate of dollar to one unit of euro (henceforth, NEER). An increase in the exchange rate translates in a depreciation of the dollar. Figure 1.4 plots the results when we order the NEER last in vector y_t, as in Eichenbaum and Evans (1995). To control for some exchange rate dynamics that may affect the US economic policy uncertainty shock, we also estimate a SVAR in which the NEER is ordered first. Figure 1.5 depicts the findings. Overall, our main results are robust in these exercises.

The economic policy changes may be correlated not only to the dynamic of inflation, output and interest rate (included in our baseline specification), but also to other several macroeconomic variables. To tackle this issue, we add to our baseline model two business cycle indicators: the Chicago Fed National Activity Index (CFNAI) and the EuroCoin, for the US economy and the Euro area one respectively. Then, we estimate a FAVAR model (Bernanke, Boivin, and Eliasz, 2005) placing the business cycle indicators instead of the industrial production indices. This because the policymakers' actions may be correlated to a wider measure of real activity (Kilian, 2011) than the industrial production index. Ordering first the business factors indicators in vector y_t, we purge the US uncertainty shock from other macroeconomic effects. Moreover, since the business indicators capture also the prospective economic conditions, we "purge" the economic policy uncertainty shock from policymakers' actions depending on economic conditions in the short-run. Our robustness check is reported in figure 1.6. The findings confirm that a US economic policy uncertainty acts as an aggregate demand shock in the Euro area.

To further investigate the accuracy of the impulse responses, we purge our proxy of economic policy uncertainty from the agents' expectations that may be related to some factors relating to the future state of economy. To control for expectation, we re-estimate our baseline model adding an index of consumer expectations based on information collected via the Michigan Survey. The Consumer Expectations Index captures how changes in economic conditions affect people and the agent's expectation about future levels of economic activity. The correlation between the US economic policy uncertainty and the Consumer expectations index is -0.34. In this way we check whether our main results depend on the fact that the economic policy uncertainty reflects the consumer confidence. Figure 1.7 shows the results. This exercise confirms our main results.

Alternative identifications of uncertainty shocks. In our baseline specification, the uncertainty shocks are identified by the "news" components of the economic policy uncertainty index proposed by Baker, Bloom, and Davis (2013). To check the sensitivity of our results to the identification of the uncertainty structural shock, we re-estimate our baseline. We substitute our measure of US and Euro area policy uncertainty, $News^{US}$ and $News^{Euro}$, with the overall economic policy uncertainty indexes, EPU^{US} and EPU^{Euro} index. Figure 1.8 displays the results. Our main results are robust to this alternative specification.

A proxy for uncertainty widely used in the literature (Bloom, 2009; Baker, Bloom, and Davis, 2013; Leduc and Liu, 2013) is the VIX index. It captures expected volatility of S&P500 index option price on the next 30 days period. Figure 1.9

shows the results, when our US and Euro Area economic policy uncertainty indexes are replaced by the VIX and VSTOXX indexes, respectively. The transmission mechanism of such financial shock on the US industrial production turns out to be different from the economic policy uncertainty one. Indeed, there is no significant reaction of the US industrial production to a financial uncertainty. Notice that the correlation between the US economic policy uncertainty index and the VIX is 0.66. Thus, this finding can be explained by the fact that the news component may be correlated to other variables which have not been captured by the VIX index. It means, as in Alexopoulos and Cohen (2009), that the Wall Street concerning may be in some way different from the Main Street ones. Turning to the Euro Area, the response of the European industrial production and inflation to a financial shock is statistically significant and negative. This exercise confirms our main results on the Euro area.

Alternative lag specification. We select the optimal number of lags in the SVAR model combining an initial lag selection based on information criteria with an LMF test for no serial correlation in the error terms. We estimate a SVAR(3). However, the lag specification matters for the accuracy of the impulse response functions.[7] To address this concern, we (re)estimate a SVAR(5). Figure 1.10 displays the results. Except the "jaggedness" of variable responses the pattern of the variables are in line with those of our baseline model estimated combining the parsimonious result of information criteria and LMF test.

Our results holds to alternative lag specification, measures of uncertainty, ordering, and additional macroeconomic variables.

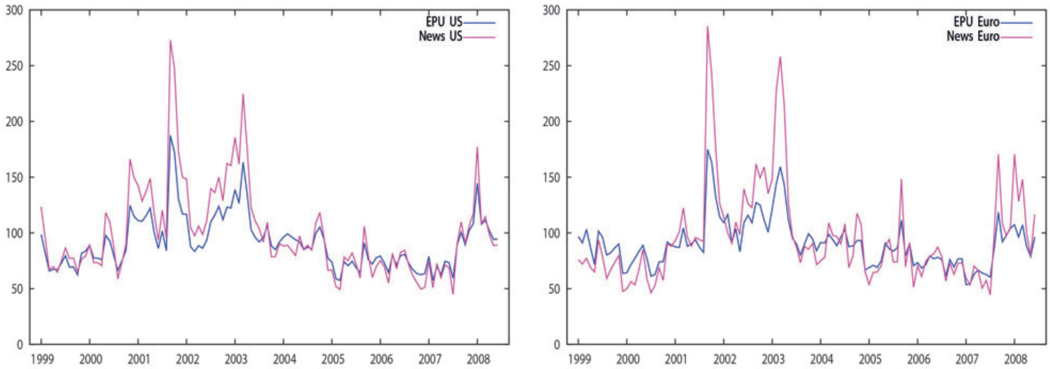
1.5 Conclusions

We investigate to what extent US economic policy uncertainty shock may trigger reactions at a macroeconomic level in the Euro area. Our VARs find a negative

[7] A large lag length relating to the number of observations may cause inefficient estimates of the parameters. Conversely, a too short lag length may arise a spurious significance of the parameter (Bjørnlan, 2000).

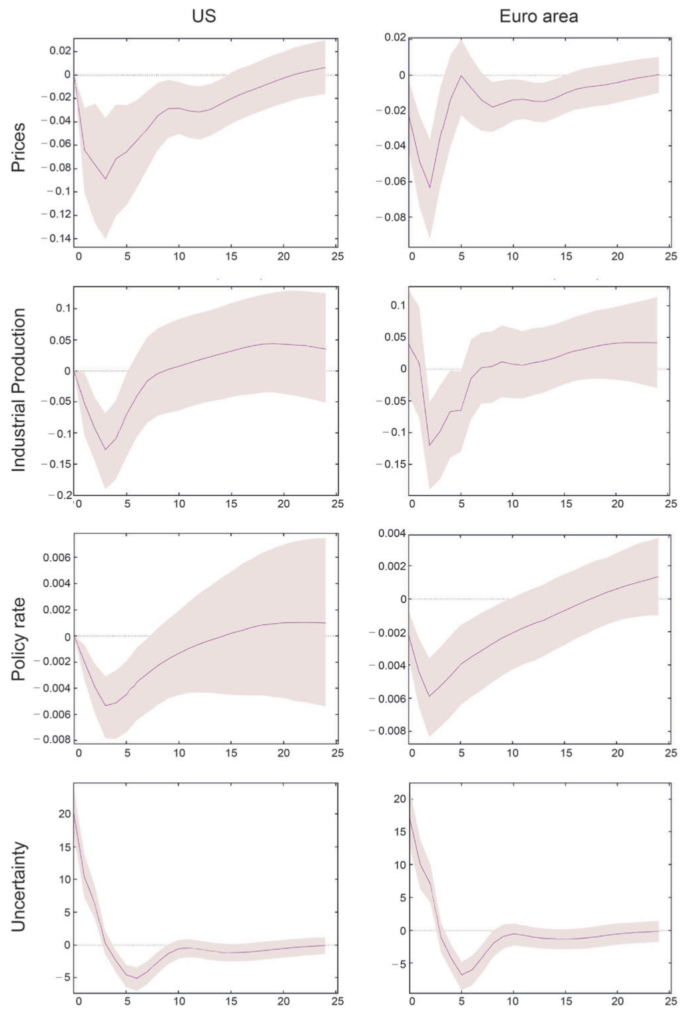
and significant reaction of Euro area price and quantity indicators to such shock. We find the contribution of exogenous variations of the US uncertainty indicator to be larger than that induced by its European counterpart.

FIGURE 1.1: Plots of time series of EPU and news policy uncertainty indexes for US and Euro (1999M1-2008M6).



Notes: Fugures plot the monthly time series of the overall uncertainty indexes and news components, both for the US (on the left) and the Euro area (on the right).

FIGURE 1.2: Empirical Impulse Responses to a US Economic Policy Uncertainty Shock



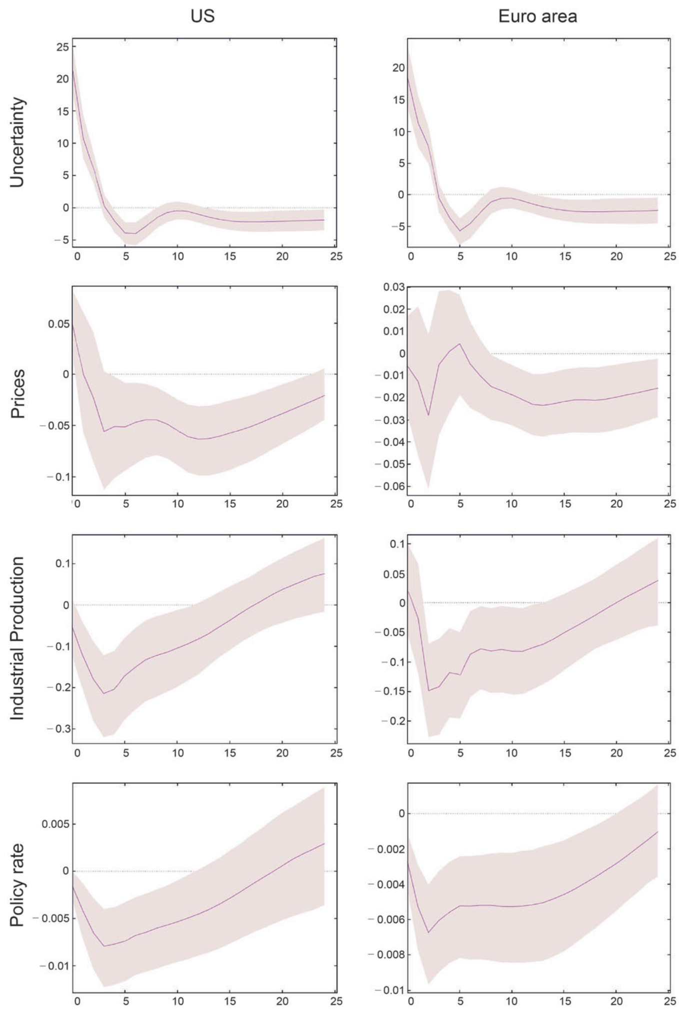
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap Kilian (1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

TABLE 1.1: Forecast error variance decomposition of the European variables due to US and European economic policy uncertainty shock (percentage)

Horizon	Consumer Prices		Industrial Production		Policy rate	
(in months)	News^{US}	$\text{News}^{\text{Euro}}$	News^{US}	$\text{News}^{\text{Euro}}$	News^{US}	$\text{News}^{\text{Euro}}$
1	2	0	0	0	7	0
6	7	1	4	2	18	3
12	6	1	3	2	11	2
18	6	1	2	2	7	2
24	6	1	2	2	6	2
6 (GSDF)	0	2	8	1	4	0
6 (CPIDF)	2	2	2	0	1	1
6 (SFC)	1	3	3	3	2	2

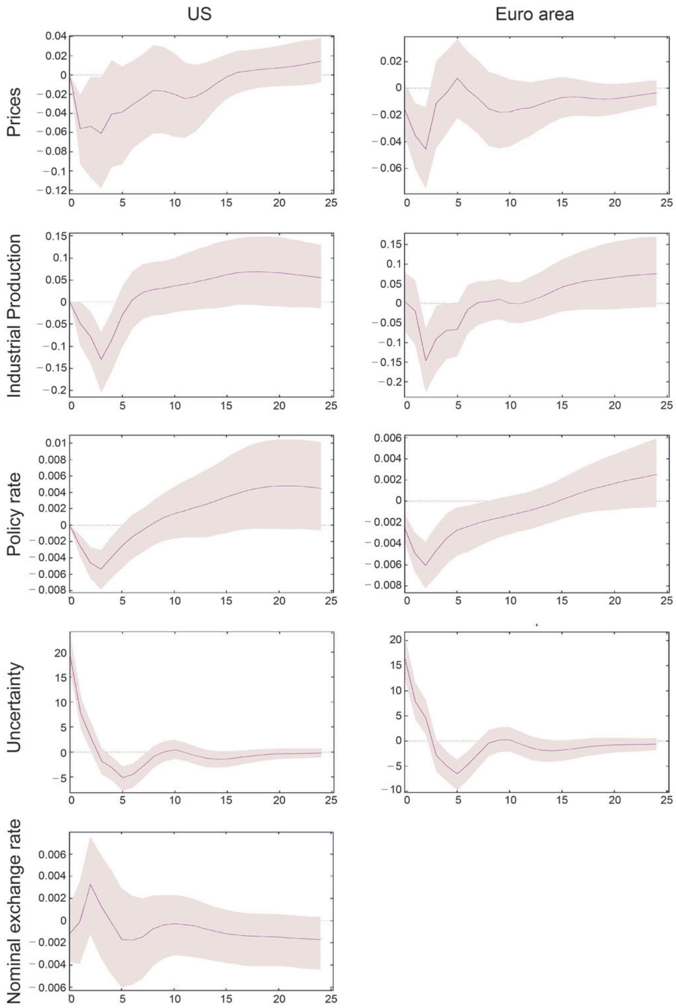
Notes: GSDF: Government Spending Disagreement Forecasts, CPIDF: CPI Disagreement Forecasts, SFC: Sample with Financial Crisis.

FIGURE 1.3: Empirical Impulse Responses to a US Economic Uncertainty Shock (trying a different ordering)



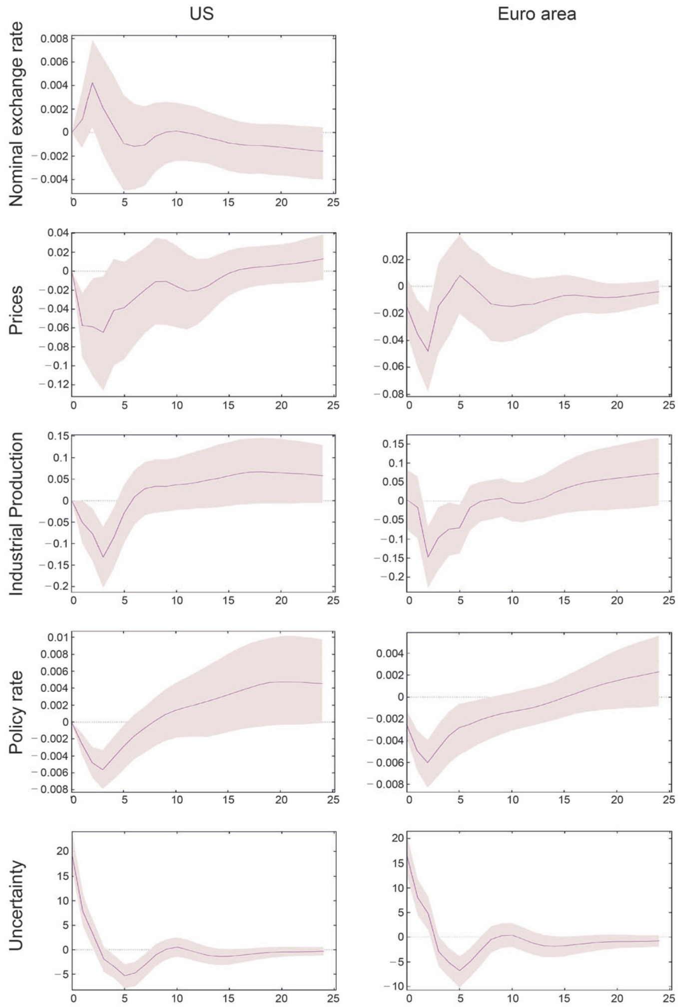
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We order the policy uncertainty indexes first in each country-block. We estimate the following vector $y_t = [News^{US}\ CPI^{US}\ IPI^{US}\ i^{US}\ News^{Euro}\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.4: Empirical Impulse Responses to a US Economic Uncertainty Shock (with the Nominal exchange rate ordered last)



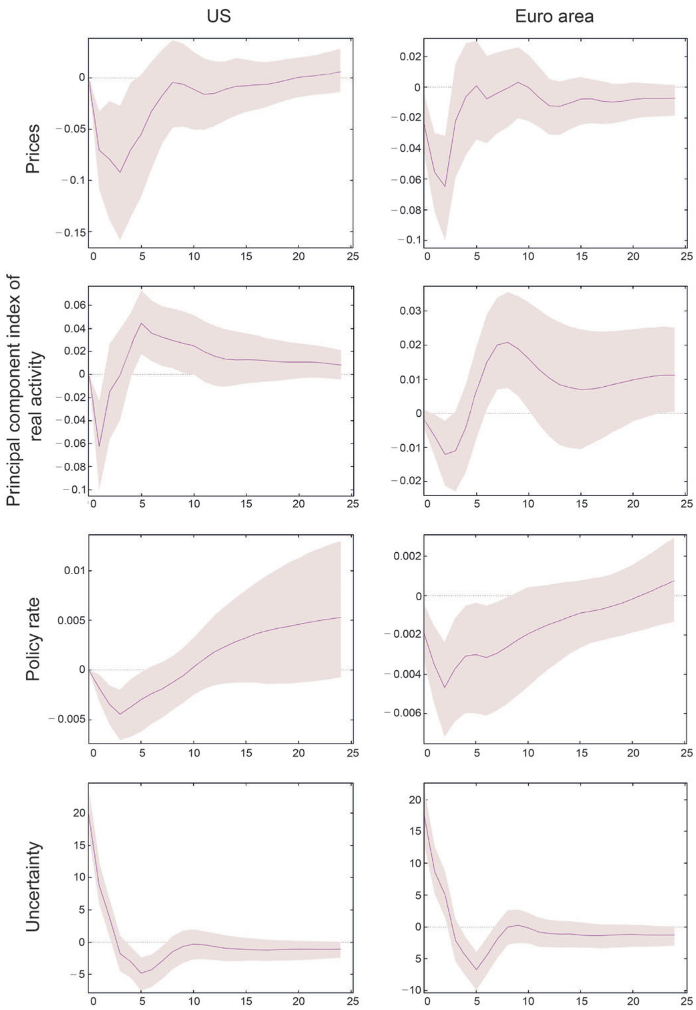
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We add to our baseline model the Nominal effective exchange rate (USD/EUR). We estimate $y_t = [CPI^{US}\ IPI^{US}\ i^{US}\ News^{US}\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}\ News^{Euro}$ $Exchange\ rate]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.5: Empirical Impulse Responses to a US Economic Uncertainty Shock (with the Nominal exchange rate ordered first)



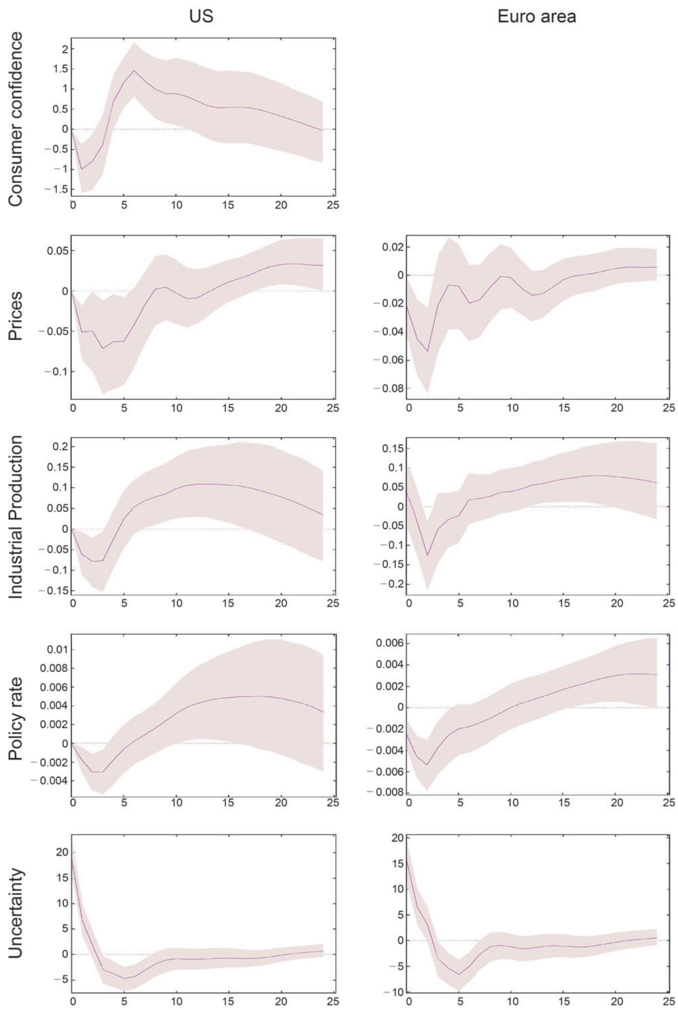
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We add to our baseline model the Nominal effective exchange rate (USD/EUR). We estimate $y_t = [Exchange\ rate\ CPI^{US}\ IPI^{US}\ i^{US}\ News^{US}\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}\ News^{Euro}]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.6: Empirical Impulse Responses to a US Economic Uncertainty Shock (with Business cycle indicators)



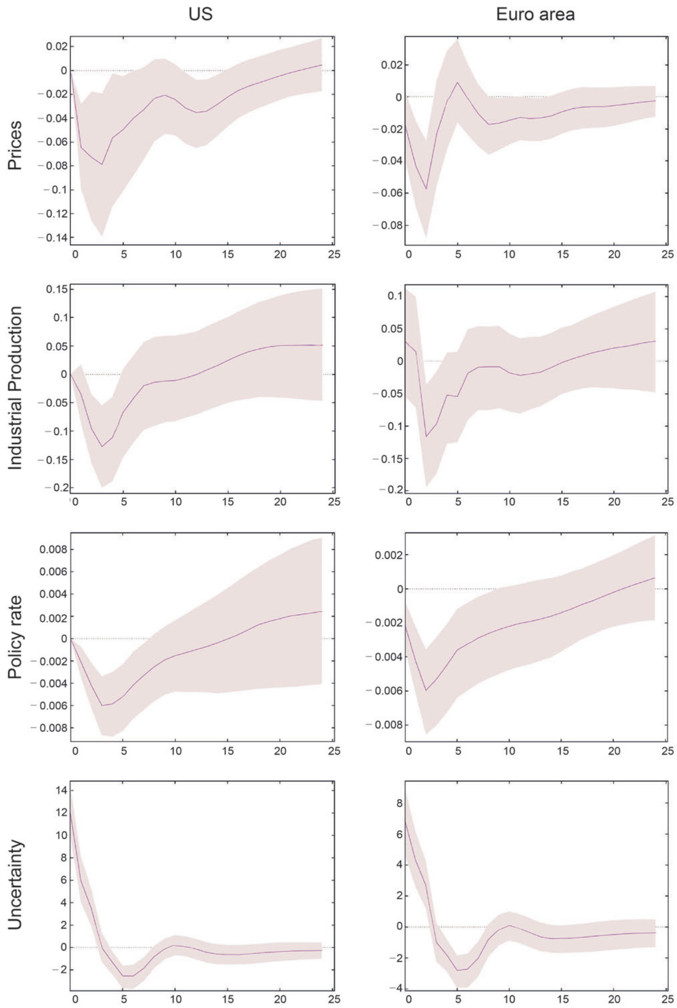
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We set two principal component indexes of real activity, the CFNAI and the EuroCoin business cycle (source: Datastream), instead of the US and Euro area industrial production. We estimate the following vector $y_t = [CPI^{US}\ CFNAI\quad i^{US}\ News^{US}\ HCPI^{Euro}$ $EuroCoin\ i^{Euro}\ News^{Euro}]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.7: Empirical Impulse Responses to a US Economic Uncertainty Shock (with the US consumer confidence)



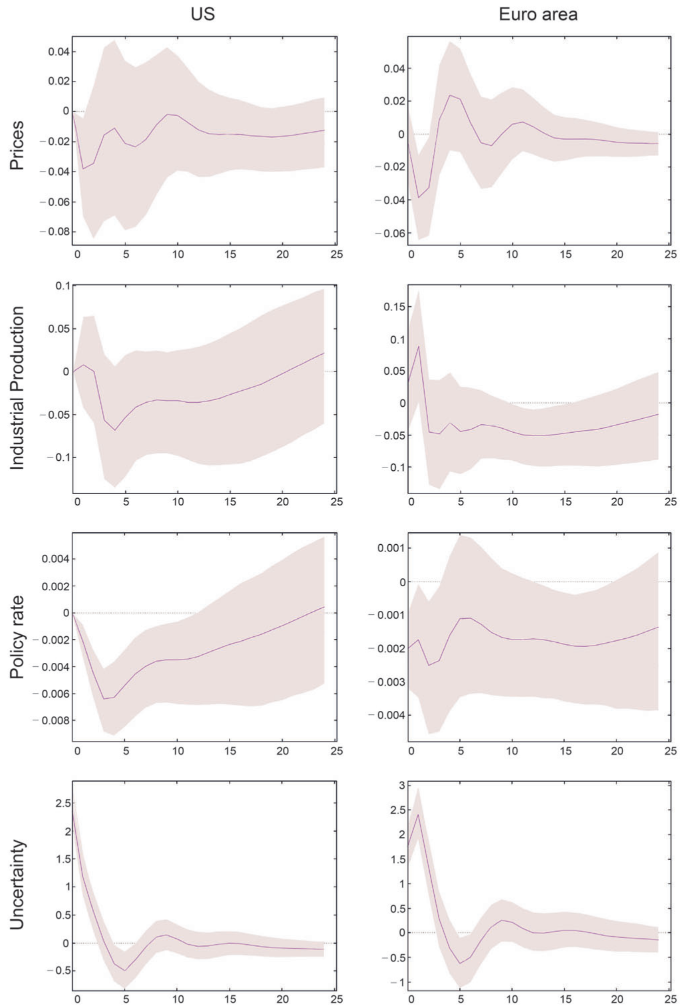
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We add to our baseline model the University of Michigan Consumer Sentiment Index. We estimate $y_t = [Cons.\ Conf\ CPI^{US}\ IPI^{US}\ i^{US}\ News^{US}\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}\ News^{Euro}]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.8: Empirical Impulse Responses to an Uncertainty Shock (substituting the economic policy uncertainty indexes)



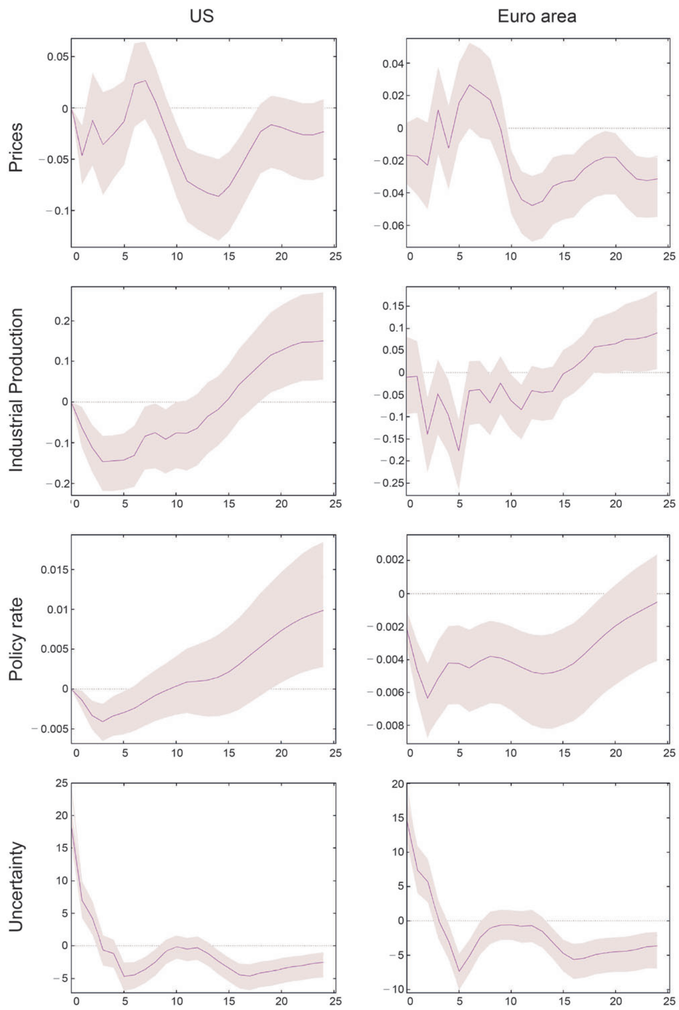
Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We set the EPU^{US} and the EPU^{Euro} instead of the US and Euro area news component and we estimate $y_t = [CPI^{US}\ IPI^{US}\ i^{US}\ EPU^{US}\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}\ EPU^{Euro}]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.9: Empirical Impulse Responses to a US Economic Uncertainty Shock (substituting the economic policy uncertainty indexes)



Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We set the VIX (Leduc and Liu, 2013; Bloom et al., 2013) and the VSTOXX instead of the US and European policy uncertainty index, respectively. We estimate the following vector $y_t = [CPI^{US}\ IPI^{US}\ i^{US}\ VIX\ HCPI^{Euro}\ IPI^{Euro}\ i^{Euro}\ VSTOXX]'$. The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

FIGURE 1.10: Empirical Impulse Responses to a US Economic Uncertainty Shock (with a different lag specification)



Notes: The figure reports orthogonalized impulse responses to an unanticipated US economic policy uncertainty shock. We estimate our baseline model with a different lag specification, a SVAR(5). The columns on the left and on the right report the IRFs for the US and European variables, respectively. The solid lines denote the median IRFs. The shaded areas identify the bootstrap-after-bootstrap (Kilian, 1998) confidence intervals at 90% level (2,000 replications). The economic policy uncertainty indexes are expressed in levels, whereas all the other variables are expressed in percent deviations with respect to their steady state. The horizontal axis identifies months.

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Chapter 2

Estimating Fiscal Multipliers: News From a Nonlinear World

Abstract

We estimate nonlinear VARs to assess to what extent fiscal spending multipliers are countercyclical in the United States. We deal with the issue of non-fundamentalness due to fiscal foresight by appealing to sums of revisions of expectations of fiscal expenditures. This measure of anticipated fiscal shocks is shown to carry valuable information about future dynamics of public spending. Results based on generalized impulse responses suggest that fiscal spending multipliers in recessions are greater than one, but not statistically larger than those in expansions. However, nonlinearities arise when focusing on "extreme" events, i.e., deep recessions vs. strong expansionary periods.

2.1 Introduction

How large is the fiscal spending multiplier? Following the lead of Blanchard and Perotti (2002), several VAR models featuring fiscal aggregates have been estimated to answer this question (for a survey, see Ramey, 2011a). However, the quantification of fiscal multipliers with standard VARs is controversial for two reasons. First, as stressed by Parker (2011), the effects of fiscal policy shocks may very well be countercyclical. Fiscal multipliers may be larger in periods of slack because of a milder crowding out of private consumption and investment due to less responsive prices (see the textbook IS-LM-AD-AS model), a constrained reaction of nominal interest rates due to the zero-lower bound (Eggertsson, 2010; Christiano, Eichenbaum, and Rebelo, 2011; Woodford, 2011; Leeper, Traum, and Walker, 2011; Fernández-Villaverde, Gordon, Guerrón-Quintana, and Rubio-Ramírez, 2012), higher returns from public spending due to countercyclical financial frictions and credit constraints (Canzoneri, Collard, Dellas, and Diba, 2011), and lower crowding out of private employment due to a milder increase in labor market tightness (Michaillat, 2014; Roulleau-Pasdeloup, 2014). Empirical evidence in favor of state-dependent fiscal multipliers is provided by, among others, Tagkalakis (2008), Auerbach and Gorodnichenko (2012, 2013a, 2013b), Bachmann and Sims (2012), Batini, Callegari, and Melina (2012), Mittnik and Semmler (2012), Baum, Poplawski-Ribeiro, and Weber (2012), Fazzari, Morley, and Panovska (2014).[1] Second, anticipation effects are likely to be of great relevance in the transmission of fiscal policy shocks, a phenomenon often referred to as "fiscal foresight" (see, among others, Yang, 2005; Fisher and Peters, 2010; Mertens and Ravn, 2011; Ramey, 2011b; Gambetti, 2012a; Gambetti, 2012b; Kriwoluzky, 2012; Favero and Giavazzi, 2012; Leeper, Walker, and Yang, 2013; Ellahie and Ricco, 2013). Modeling a standard set of U.S. variables with a medium-scale

[1]Other forms of state-dependence have been identified in the literature. Corsetti, Meier, and Müller (2012) investigate the sensitivity of government spending multipliers to different economic scenarios. They find fiscal multipliers to be particularly high during times of financial crisis. Rossi and Zubairy (2011) and Canova and Pappa (2011) show that fiscal multipliers tend to be larger when positive spending shocks are accompanied by a decline in the real interest rate. Perotti (1999) shows that fiscal multipliers may depend on the debt-to-GDP ratio in place when fiscal shocks occur. For a DSGE-based quantification of fiscal multipliers in presence of normal vs. abnormal debt-to-GDP ratios, see Cantore, Levine, Melina, and Pearlman (2013).

structural model that allows for foresight up to eight quarters, Schmitt-Grohe and Uribe (2012) find that about sixty percent of the variance of government spending is due to anticipated shocks. Unfortunately, in presence of fiscal foresight, standard VARs - which rely on current and past shocks to interpret the dynamics of the modeled variables - are typically "non-fundamental", in that they do not embed the information related to "news shocks", i.e., future shocks anticipated by rational agents.[2] Leeper, Walker, and Yang (2013) work with a variety of fiscal models and show that the anticipation of tax policy shocks severely affects VAR exercises aiming at identifying fiscal shocks. Forni and Gambetti (2010a) and Ramey (2011b) show that government spending shocks estimated with standard fiscal VARs are predictable, i.e., they are non-fundamental.

This paper estimates *state-dependent fiscal multipliers* by explicitly addressing the issue of *fiscal foresight*. We tackle the issue of non-fundamentalness by jointly modeling a measure of *anticipated ("news") fiscal spending shocks* along with a set of standard macro-fiscal variables. Such a measure of fiscal news is the *sum of revisions of expectations about future government spending* collected by the Survey of Professional Forecasters. As shown by Gambetti (2012a, 2012b) and Forni and Gambetti (2014a), this measure of fiscal shocks is particularly powerful to capture the effects of fiscal spending shocks when the implementation lag of fiscal policy is larger than one quarter, a very plausible assumption as for U.S. fiscal policy decisions.[3] We include this measure of fiscal news in a nonlinear Smooth Transition Vector AutoRegressive (STVAR) model, which we use to discriminate dynamic responses to fiscal shocks in bad and good times (i.e., recessions vs. expansions). Our multipliers are computed as the integral of the impulse response of output (up to a chosen horizon) divided by the integral of the response of fiscal expenditure (up to the same horizon) and rescaled by the sample mean value of

[2]For a recent discussion on non-fundamentalness in the VAR context and a survey of the main contributions in this area, see Beaudry and Portier (2014).

[3]Yang (2005) shows that the average implementation lag for major postwar U.S. income tax legislation is about seven months. Mertens and Ravn (2011) find that the median implementation lag is six quarters. Leeper, Richter, and Walker (2012) calibrate tax foresight and government spending foresight to range between two and eight quarters (the former) and between three and four quarters (the latter).

the output-public spending ratio.[4] To assess the effects of public spending shocks on output and estimate fiscal multipliers in recessions and expansions, we compute Generalized Impulse Response Functions (GIRFs), which model the endogeneity of the transition from a state to another after a fiscal shock. Importantly, as explained by Koop, Pesaran, and Potter (1996), GIRFs allow us to scrutinize the role played by different initial conditions. We then isolate *"extreme" events*, i.e., deep recessions and strong expansions, with the aim of understanding if *fiscal multipliers are larger in very severe economic conditions*. To our knowledge, this key policy-relevant question has not been previously studied in the empirical literature on fiscal multipliers.

Our results are the following: i) anticipated fiscal expenditure shocks trigger a significant reaction of output; ii) such a reaction is not statistically different across different phases (recessions/expansions) of the U.S. business cycle; iii) the reaction becomes statistically different for extreme phases of the business cycle, i.e., deep recessions vs. strong expansions; iv) fiscal multipliers in recessions are statistically larger than one; v) spending shocks in recessions have a noticeable stabilization effect and substantially reduce the probability that the economy will remain slack. These results are robust to a wide battery of checks, including i) the employment of a "purged" measure of fiscal news, which is constructed using information available to survey respondents when they formulate their expectations over future public spending, to account for potential identification issues; ii) the use of the fiscal news constructed by Ramey (2011b), which allows us to extend our sample back to 1947, to control for small-sample biases that may affect our data-intensive estimator; iii) the role of debt, to account for the role played by fiscal strains in computing multipliers; iv) several different VAR specifications.

Our paper represents a novel contribution under several respects. First, our VAR

[4] Our results are robust to the employment of an alternative way of computing fiscal multipliers, i.e., the ratio of the "peak" value of the impulse responses of output and public spending rescaled by the sample mean ratio of the levels of ouput over public spending. Our Appendix (available upon request) documents the results obtained with this alternative way of computing fiscal multipliers.

jointly accounts for two relevant issues for the quantification of fiscal multipliers: fiscal foresight and state dependence. Second, we estimate the response of economic aggregates to fiscal shocks via GIRFs, which allow us to endogenize the possibly stabilizing effects of fiscal policy. Third, the use of GIRFs allows us to address a previously unexplored issue, i.e., the role played by business cycle conditions for the quantification of fiscal multipliers, which we investigate by distinguishing between "extreme" and "moderate" business cycle phases. As a result, we are able to establish some new stylized facts about government spending multipliers in the U.S., in particular the fact that firm evidence of state dependence arises only when looking at extreme phases of the business cycle.

The closest papers to ours are Auerbach and Gorodnichenko (2012, 2013a), Owyang, Ramey, and Zubairy (2013), and Ramey and Zubairy (2014). Auerbach and Gorodnichenko (2012, 2013a) employ a STVAR model and find evidence of countercyclical fiscal multipliers.[5] There are substantial differences between Auerbach and Gorodnichenko's contributions and ours. First, they investigate the role of unanticipated fiscal spending shocks. Differently, we focus on anticipated changes in fiscal spending. Second, their impulse responses are conditionally linear, i.e., expansionary fiscal spending shocks are, by construction, not allowed to drive the economy out of a recession. As pointed out by the same authors, this assumption provides an "upper bound" for their estimates of the fiscal multiplier in recessions, because it does not allow the returns from fiscal spending to be decreasing as the economy exits a recession. Our approach links the evolution of the variables in our STVAR to the probability of being in a recession, which is then endogenously modeled. Third, our focus is on "extreme" events, i.e., realizations on the tails of the distribution of our business cycle indicator (like the 2007-09 crisis). Our main result is that, while fiscal multipliers may be acyclical when recessions and expansions are considered all alike (i.e., they may be similar when considering the average effect in recessions vs. expansions), they are likely to be large in presence of particularly severe economic conditions. Owyang, Ramey, and Zubairy (2013) and Ramey and Zubairy (2014) employ local-projection methods á la Jordá (2005)

[5] For a similar exercise focusing on the role of business confidence, see Bachmann and Sims (2012).

to investigate the nonlinearity of fiscal multipliers. They find no evidence of larger fiscal multipliers during downturns as for the United States. The comparability between our exercises and theirs is not immediate due to a number of different modeling choices (construction of the news shocks, length of the sample, construction of the impulse responses, among others). We notice that our results are similar to theirs in that we also do not find larger fiscal multipliers in recessions on average. However, when it comes to deep recessions vs. strong expansions, we find such larger multipliers to arise.

Other strands of the literature have dealt with fiscal foresight and anticipated fiscal spending shocks in VARs. Mertens and Ravn (2010) recover the non-fundamental responses to an anticipated fiscal policy shock via economic theory-driven restrictions to gauge information about economic agents' anticipation rate. Such a rate is then used as an input in Blaschke matrices to flip the roots that cause the non-invertibility of the VMA representation of fiscal spending and output. Kriwoluzky (2012) recovers reduced-form innovations by estimating a VARMA model using the Kalman filter. Then, he identifies anticipated fiscal shocks via theoretically-supported sign restrictions. Ramey and Shapiro (1998) follow a narrative approach to identify exogenous changes in military spending related to wars. Ramey (2011b) constructs a measure of changes in the expected present value of government spending. Fisher and Peters (2010) construct a measure of excess returns of large U.S. military contractors which is shown to anticipate future military spending shocks. Ben Zeev and Pappa (2014) identify U.S. defense news shocks as the shocks that best explain future movements in defense spending over a five year horizon and are orthogonal to current defense spending. All these contributions show that, at least qualitatively, anticipated positive fiscal shocks induce a significant increase in output.[6] Perotti (2007, 2011), Ramey (2011b), Gambetti (2012a, 2012b), Blanchard and Leigh (2013), Alesina, Favero, and Giavazzi (2014), Forni and Gambetti

[6] Another interesting approach to account for fiscal foresight rests on the use of municipal bond spreads. This bond spread is well-known to have predictive power for tax changes and can therefore be used to control for anticipated tax changes (see, among others, Poterba (1989), Fortune (1996), and Kueng (2014)). Leeper, Richter, and Walker (2012) show that spreads with maturity lengths of 1 and 5 years are very informative about future tax events. Our paper deals with anticipated fiscal spending shocks. We leave the analysis of anticipated tax shocks to future research.

(2014a), and Ricco (2014) work with expectations revisions in different modeling frameworks. Our paper complements these contributions, in that it quantifies the effects of anticipated fiscal spending shocks with a nonlinear model focusing on extreme events.[7]

The structure of the paper is the following. Section 2 deals with the issue of non-fundamentalness in the macro-fiscal context due to the presence of fiscal foresight, and explains why the sums of revisions of fiscal expectations variable employed in our analysis helps solving the issue. Section 3 offers statistical support to the role of nonlinearities in this context and presents the Smooth Transition VAR model employed in our analysis. Our main results are shown in Section 4, which deals with the computation of fiscal multipliers in recessions and expansions, and Section 5, which focuses on extreme events. Section 6 documents a battery of robustness checks. Concluding remarks are provided in Section 7.

2.2 Non-fundamentalness and expectations revisions

The role of expectations revisions. As anticipated in our Introduction, standard fiscal VARs may return severely biased impulse responses in presence of news shocks. Consider the model

$$y_t = \delta E_t y_{t+1} + g_t + \omega_t \quad (2.1)$$

$$g_t = \varepsilon_{t-h} + \phi_1 \varepsilon_{t-h-1} + \ldots + \phi_{q-h-1} \varepsilon_{t-(q-1)} + \phi_{q-h} \varepsilon_{t-q} = \Phi(L)\varepsilon_t \quad (2.2)$$

[7] Admittedly, the theoretical papers modeling nonlinearities cited in this Introduction mainly consider models in which government spending is implemented without lags. As for the zero lower bound, however, Christiano, Eichenbaum, and Rebelo (2011) conduct an exercise in which they model implementation lags in their framework featuring the zero lower bound. They find that a key determinant of the size of the multiplier is indeed the state of the world in which new government spending comes on line. Our conjecture is that such asymmetric effects may be present also when anticipated fiscal shocks hit economic systems characterized by state-dependent financial constraints and labor market downward rigidities.

where $|\delta| < 1, \phi_i > 0\ \forall i, h \geq 0, q \geq h$, and $\phi_0 = 0$. The forward-looking process y_t - say, output measured as log-deviations from its trend - is affected by the exogenous stationary process g_t - say, a fiscal shock - plus a random shock ω_t, which is assumed to capture non-fiscal spending shocks affecting output and which is assumed to be *i.i.d.* with zero mean and unit variance. The process (2.2) features $q - h + 1$ moving average terms. If $h = 0$ and $q > 0$, the process (2.2) features an unanticipated, ε_t, as well as anticipated shocks ε_{t-q} for $q > 0$. For $h > 0$, the process (2.2) would feature only unanticipated shocks, where h is the number of periods of foresights. The process g_t is a news-rich process if $|\phi_i| > 1$ for at least one $i > 0$ (Beaudry and Portier, 2014). In all cases, $\{\varepsilon_{t-j}\}_{j=h}^{q}$ is said to be fundamental for g_t if the roots of the polynomial $\Phi(L)$ lie outside the unit circle (Hansen and Sargent, 1991). Importantly, if the g_t process is non-fundamental, its structural shock is not recoverable by employing current and past realizations of g_t only. Consequently, its impulse response to an anticipated shock as well as the dynamic responses of other variables – in this example, y_t – will not be correctly recovered by estimating a VAR in y_t and g_t.

We assume that agents have rational expectations and observe news shocks without noise.[8] It can be shown that, if the period of foresight $h \geqslant 1$ is known, the problem of non-fundamentalness in model (2.1)-(2.2) can be solved by alternatively including: i) the h-step-ahead expectation, $E_t g_{t+h}$, if $h = q$; ii) the h-step-ahead expectation revision, $E_t g_{t+h} - E_{t-1} g_{t+h}$, if $h < q$. However, if $h > 1$ is unknown, expectation revisions are not of help. To solve this issue, Gambetti (2012a) proposes to use a news variable defined as

[8] Forni, Gambetti, Lippi, and Sala (2013) investigate the case in which economic agents deal with noisy news. Agents are assumed to receive signals regarding the future realization of TFP shocks. Since such signals are noisy, agents react not only to genuinely informative news, but also to noise shocks that are unrelated to economic fundamentals. They find that such noise shocks explain about a third of the variance of output, consumption, and investment. We leave the quantification of the role of noise shocks in the fiscal context to future research.

$$\eta_{1J}^{g}=\sum_{j=1}^{J}\left(E_t g_{t+j}-E_{t-1}g_{t+j}\right)=\begin{cases}\left(1+\phi_1+...+\phi_{J-h}\right)\varepsilon_t \text{ if } J<q\\ \left(1+\phi_1+...+\phi_{q-h}\right)\varepsilon_t \text{ if } J\geqslant q\end{cases}, \tag{2.3}$$

which correctly identifies the news shock if $J \geqslant h$.[9] Our Appendix provides further discussions and derivations as regards this news variable.

The News13 variable. We will then consider a fiscal VAR augmented with a measure of news constructed by summing up revisions of expectations as follows:

$$\eta_{13}^{g}=\sum_{j=1}^{J}(E_t g_{t+j}-E_{t-1}g_{t+j}) \tag{2.4}$$

where $E_t g_{t+j}$ is the forecast of the growth rate of real government spending from period $t+j-1$ to period $t+j$ based on the information available at time t. Hence, $E_t g_{t+j}-E_{t-1}g_{t+j}$ represents the "news" that becomes available to private agents between time $t-1$ and t about the growth rate of government spending j periods ahead. We use data coming from the Survey of Professional Forecasters (SPF), which collects forecasts conditional on time $t-1$ of variables up to time $t+3$. This is the reason why our baseline analysis will be conducted by considering the variable η_{13}^{g}.[10]

Information content of expectations revisions. To assess the statistical relevance of our news variable for the dynamics of public expenditure, we regress public spending on a constant and three lags of the dependent variable, public receipts, real GDP, and one lag of the measure of news η_{13}^{g} (a detailed description of the

[9] If $J<h$, the news variable would have no predictive content about fiscal shocks, and would be equal to zero. In our sample, however, this never happens. This is consistent with the evidence in Leeper, Richter, and Walker (2012), who report an average implementation lag of about three quarters. In our example above, h should be interpreted as the minimum temporal gap between the announcement of the implementation of future fiscal spending and the realization of the spending itself (which may take more than one quarter), rather than the mean value. Hence, also the effects of the announcement of future spending whose full implementation would take more than J quarters would be captured by our news, as long as the minimum lag h is less than J.

[10] SPF data are affected by frequent changes in the base years. Forecast errors on the growth rates are not affected by these changes. Hence, they are preferable to forecast errors computed with SPF levels. About this point, see also Perotti (2011).

data is provided in Section 3). This regression augments the public spending equation of a trivariate VAR system modeling the "usual suspects" (public spending, tax receipts, output) with our news variable lagged one period.[11] Public spending shocks are often identified with a Cholesky decomposition of the covariance matrix of the VAR residuals. Hence, the (orthogonalized) residuals of the public spending equation are interpreted as public spending shocks. As shown in Table 2.1 - which collects the p-values for our η^g_{13} variable in the equation described above - news shocks are found to carry significant information about the future evolution of public spending. This implies that the trivariate fiscal VAR without news is non-fundamental. Digging deeper, we find that all the three components (forecast revisions) included in η^g_{13} have some predictive power. Overall, this empirical exercise highlights the significant contribution of news revisions regarding *future* realizations of public expenditure. Differently, revisions of expectations based on nowcasting, i.e., $E_t g_t - E_{t-1} g_t$, turn out to be insignificant at the 90% confidence level (see Table 1, last column). In line with Ricco (2014), this result suggests that revisions based on "nowcasts" (revision of expectations at time t of contemporaneous public expenditures) are possibly of help in identifying truly *unanticipated* fiscal shocks, rather than *anticipated, news* shocks.[12]

Overall, our results i) show that, from a statistical standpoint, residuals typically employed in a standard trivariate fiscal VAR cannot be interpreted as fiscal shocks; ii) suggest that the components of the variable η^g_{13}, which we interpret as a measure of anticipated fiscal shocks, can augment the information content of our VAR system. These results are consistent with the outcome of the Granger-causality tests conducted by Gambetti (2012b), who shows that η^g_{13} Granger-causes fiscal spending at different horizons.[13]

[11]The regression includes variables in (log-)levels and the news η^g_{13} variable in cumulated sums to preserve the same order of integration. This is consistent with the modeling choices of our baseline VAR analysis (specified in the next Section).

[12]These results are conditional on news variables constructed as revisions of the mean predicted values of the levels of future government spending as collected by the Survey of Professional Forecasters. Similar results were obtained by employing median values of such forecasts, as well as variables expressed in growth rates.

[13]In a recent paper, Perotti (2011) questions the use of the SPF forecast errors employed by Ramey (2011a) to isolate fiscal spending anticipated shocks. In particular, he shows that the one-step-ahead predictive power of the forecast revisions as for federal spending is quite

Extreme realizations of the news spending variable: An interpretation. Figure 2.1 plots our news variable (an updated version of Gambetti, 2012b). The standardized variable η^g_{13} conveys useful information about fiscal policy shocks in the United States. To see this, we isolate the seven realizations which exceed two in absolute value, and provide an interpretation based on the recent U.S. fiscal history. The 1983Q1 positive realization is associated to Ronald Reagan's "Evil Empire" and "Star Wars" speeches, with which the U.S. President announced a forthcoming increase in military spending. The 1986Q1 negative spike reflects the speech given in January 1986 by Mikhail Gorbachev, who proposed decommissioning all nuclear weapons by 2000 in the early stage of the "Perestrojka" period. The 1987Q1 positive forecast revisions might be due to the mid-term Senate elections won by the Democrats in November 1986 plus the questioned constitutionality of the Gramm-Rudman-Hollings Balanced-Budget Act. The 1987Q4 forecast revisions are due to announcements about spending cuts for the Pentagon. The fall of the Berlin Wall in November 1989 is behind the negative spike in 1989Q4. The war in Afghanistan rationalizes the positive peak in 2001Q4. Finally, the upward spike in 2009Q1 can be associated to Obama's stimulus package.

Comparison with Ramey's (2011b) news variable. Figure 2.1 also plots the military spending news variable constructed by Ramey (2011b), and extended up to 2010Q4 by Owyang, Ramey, and Zubairy (2013).[14] It appears that the η^g_{13} variable anticipates changes in Ramey's, or at least it is not anticipated by the latter. To corroborate this statement, we run Granger-causality tests based on an estimated bivariate VAR with one lag involving the military spending news proposed by Ramey (2011b) (as well as its updated version by Owyang, Ramey, and Zubairy, 2013) and the η^g_{13} variable. Table 2.2 collects the outcome (p-values associated to testing the null hypothesis that the column variable does not Granger-cause the

modest, since such revisions are shown to be noisy. Our results are fully consistent with Perotti (2011) analysis, in that we also reject the relevance of very short-term SPF forecast revisions on future fiscal spending. This evidence suggests the need of searching for anticipation effects beyond one-quarter relative to the moment in which predictions are formulated, and supports the employment of a variable like η^g_{13}.

[14] Ramey (2011b) employs *Business Week* and other newspaper sources to construct an estimate of changes in the expected present value of goverment spending (nominal spending divided by nominal GDP one period before).

alternative news measure) of this exercise for our benchmark sample and a shorter sample to account for the fact that, for the first five years in the benchmark sample, Ramey (2011b) variable is equal to zero. While the contribution of our news shock variable finds large statistical support, Granger-causality running from Ramey's shock to ours is clearly rejected by the data. The same evidence emerges when employing the news variable by Owyang, Ramey, and Zubairy (2013), which includes observations related to the 2007-2009 recession. Again, these results are in line with those reported in Gambetti (2012b), who also finds Ramey's news shock to be predicted by forecast revisions over one quarter.

2.3 Econometric approach: A STVAR macro-fiscal model

Modeling choices. We assess the state-dependence of fiscal spending multipliers to news shocks by estimating a Smooth-Transition VAR model (for an extensive presentation, see Terësvirta, Tjostheim, and Granger, 2010). Our STVAR framework reads as follows:

$$\boldsymbol{X}_t = F(z_{t-1})\boldsymbol{\Pi}_R(L)\boldsymbol{X}_t + (1 - F(z_{t-1}))\boldsymbol{\Pi}_E(L)\boldsymbol{X}_t + \varepsilon_t, \tag{2.5}$$

$$\varepsilon_t \sim N(0, \boldsymbol{\Omega}_t), \tag{2.6}$$

$$\boldsymbol{\Omega}_t = F(z_{t-1})\boldsymbol{\Omega}_R + (1 - F(z_{t-1}))\boldsymbol{\Omega}_E, \tag{2.7}$$

$$F(z_t) = \exp(-\gamma z_t)/(1 + \exp(-\gamma z_t)), \gamma > 0, z_t \sim N(0, 1). \tag{2.8}$$

where $\boldsymbol{X}_t$ is a set of endogenous variables which we aim to model, $F(z_{t-1})$ is a transition function which captures the probability of being in a recession, γ regulates the smoothness of the transition between states, z_t is a transition indicator, $\boldsymbol{\Pi}_R$ and $\boldsymbol{\Pi}_E$ are the VAR coefficients capturing the dynamics of the system during recessions and expansions (respectively), ε_t is the vector of reduced-form residuals having zero-mean and whose time-varying, state-contingent variance-covariance matrix is $\boldsymbol{\Omega}_t$, and $\boldsymbol{\Omega}_R$ and $\boldsymbol{\Omega}_E$ stand for the covariance structure of the residuals in recessions and expansions, respectively. The modeling assumption is that the variables can be described with a combination of two linear VARs, one suited to describe the economy during recessions and the other during expansions. The transition from a state to another is regulated by the standardized transition variable z_t. The smoothness parameter γ affects the probability of being in a recession $F(z_t)$, i.e., the larger the value of γ, the faster the transition from a state to another. Notably, the model (2.5)-(2.8) allows for nonlinearities to arise from both the contemporaneous and the dynamic relationships of the economic system. Our baseline analysis refers to the vector $\boldsymbol{X}_t = [G_t, T_t, Y_t, \eta^g_{13,t}]'$, where G is the log of real government (federal, state, and local) purchases (consumption and investment), T is the log of real government receipts of direct and indirect taxes net of transfers to business and individuals, and Y is the log of real GDP.[15] The construction of G and T closely follows Auerbach and Gorodnichenko (2013).[16] The variable η^g_{13} is the public expenditure news variable (2.4). The variables are expressed in levels because of possible cointegration relationships. Consistently, the variable η^g_{13} is considered in cumulated sums to preserve the same order of integration as the other variables included in the vector. Our sample of U.S. data spans the period 1981Q3-2013Q1, 1981Q3 being the first available quarter to construct the news variable.[17]

The choice of the transition variable z_t and the calibration of the smoothing parameter γ are justified as follows. As in Auerbach and Gorodnichenko (2012),

[15] Our fiscal aggregates are constructed using the Bureau of Economic Analysis' NIPA Table 3.1. Current tax receipts are constructed as the difference between current receipts and government social benefits. Fiscal expenditure is the sum of consumption expenditure and gross government investment from which we subtract the consumption of fixed capital. Data on real GDP and the implicit GDP deflator (which we use to deflate all nominal series) are provided by the Federal Reserve Bank of St. Louis.

[16] Auerbach and Gorodnichenko (2013) check and verify the robustness of the results in Auerbach and Gorodnichenko (2012) to the employment of a different definition of the net tax series that avoids the double-counting of mandatory Social Security contributions.

[17] Our interpretation of the news variable here is that of an instrument to gauge the real effects of anticipated changes in fiscal spending. We recall that different identification approaches may very well lead to the construction of different, but in principle equally valid, instruments. For an elaboration of this point, see Favero and Giavazzi (2012).

Bachmann and Sims (2012), Caggiano, Castelnuovo, and Groshenny (2014), and Berger and Vavra (2014), we employ a standardized moving average of the real GDP quarter-on-quarter percentage growth rate.[18] We calibrate the smoothness parameter γ to match the observed frequencies of the U.S. recessions as identified by the NBER business cycle dates, i.e. 15% in our sample. Then, we define as "recession" a period in which $F(z_t) \geqslant 0.85$, and calibrate γ to obtain $\Pr(F(z_t) \geqslant 0.85) \approx 15\%$. This metric implies a calibration $\gamma = 2.3$. The choice is consistent with the threshold value $\overline{z} = -0.75\%$ discriminating recessions and expansions, i.e., realizations of the standardized transition variable z lower (higher) than the threshold will be associated to recessions (expansions).[19] Figure 2.2 plots the transition function $F(z_t)$. Clearly, high realizations of $F(z_t)$ tend to be associated with NBER recessions. Importantly, our results are robust to the employment of alternative calibrations of the slope parameter γ that imply a number of recessions in our sample ranging from 10% to 20%, where the lower bound is determined by the minimum amount of observations each regime should contain according to Hansen (1999) (checks not shown here for the sake of brevity, but available upon request).

Identification of the anticipated fiscal shock. Following Fisher and Peters (2010), we order the news variable η^g_{13} last in our vector and orthogonalize the reduced-form residuals of the VAR via a Cholesky-decomposition of the variance-covariance matrix. We analyze the implications of this versus alternative strategies to identify fiscal news shocks in Section 5.

Statistical evidence in favor of nonlinearity. For our vector of endogenous variables $\mathbf{X}_t$, we test and clearly reject the null hypothesis of linearity in favor of the (Logistic) Smooth Transition Vector AutoRegression via the multivariate test

[18] The transition variable z_t is standardized to render our calibration of γ comparable to those employed in the literature. We employ a backward-looking moving average involving four realizations of the real GDP growth rate.

[19] The corresponding threshold value for the non-standardized moving average real GDP growth rate is equal to 0.34%. The sample mean of the non-standardized real GDP growth rate in moving average terms is equal to 0.71, while its standard deviation is 0.50. Then, its corresponding threshold value is obtained by "inverting" the formula we employed to obtain the standardized transition indicator z, i.e., $\overline{z}^{nonstd} = -0.75 \times 0.50 + 0.71 = 0.34$.

proposed by Teräsvirta and Yang (2013) in presence of a single transition variable. Details on this test and its implementation are presented in our Appendix.

Model estimation. Given the high nonlinearity of the model, we estimate it via the Monte-Carlo Markov-Chain algorithm developed by Chernozhukov and Hong (2003). The (linear/nonlinear) VARs include three lags. This choice is based on the Akaike criterion applied to a linear model estimated on the full-sample 1981Q3-2013Q1.

2.4 Generalized impulse responses and fiscal multipliers

This Section reports the estimated impulse responses to an anticipated fiscal spending shock. Following Koop, Pesaran, and Potter (1996), we compute generalized impulse responses to take into account the interaction between the evolution of the variables in the vector $\mathbf{X}_t$ and the transition variable, the latter being directly influenced by the evolution of output. In other words, we model the feedback from the evolution of output in the vector $\mathbf{X}_t$ to the transition indicator z_t and, consequently, the probability $F(z_{t-1})$. Hence, in computing our GIRFs, the probability $F(z)$ is endogenized.[20] Koop, Pesaran, and Potter (1996) and Ehrmann, Ellison, and Valla (2003) show that initial conditions affect the computation of the GIRFs. In our benchmark exercise, we randomize over all possible histories within each state, so to control for the role of initial conditions.[21] We compute the GIRFs by

[20]Recall that our transition indicator $z_t \equiv \frac{1}{4}(\Delta Y_t + \Delta Y_{t-1} + \Delta Y_{t-2} + \Delta Y_{t-3})$, i.e., the relationship between z_t and $\Delta Y_{t-i}, i = 0, 1, 2, 3$ features no stochastic elements. Hence, stochastic singularity prevents us from estimating our model jointly with the evolution of z_t. Following Koop, Pesaran, and Potter (1996), our GIRFs are based on simulations that take into account the link between $\mathbf{X}_t$ and z_t after the estimation of our econometric framework.

[21]Following Koop, Pesaran, and Potter (1996), our GIRFs are computed as follows. First, we draw an initial condition, i.e., starting values for the lags of our VARs as well as the transition indicator z, which - given the logistic function (2.8) - gives us the value for $F(z)$. Then, we simulate two scenarios, one with all the shocks identified with the Cholesky decomposition of the VCV matrix (2.7), and another one with the same shocks plus a $\delta > 0$ corresponding to the first realization of the news shock. The difference between these two scenarios (each of which accounts for the evolution of $F(z)$ by keeping track of the evolution of output and, therefore, z) gives us the GIRFs to a fiscal news shock δ. Per each given initial condition z, we compute

normalizing the news shocks to one.[22]

GIRFs. Figure 2.3 reports the impact of a government spending news shock computed with our linear and nonlinear VARs. The responses obtained with our linear model point to a delayed short-run increase in government expenditure and output, and a decrease in government receipts. Public spending reaches its peak value after about three years. Differently, output increases for the first three quarters after the shock, then gradually goes back to zero, and crosses the zero line about 10 quarters after the shock.

Next, we look at the evidence coming from the nonlinear VAR. Interestingly, the estimated response of output is persistently stronger under recessions. Output increases in expansions in the short-run, but the increase is much milder compared to recessions, and vanishes after about four quarters. Another difference between the two states is the reaction of government spending itself, which is always positive but stronger in recessions. Tax receipts react asymmetrically in the short run, then their patterns become more similar.

Are the reactions of output in recessions and expansions different from a statistical standpoint? Figure 2.4 plots the GIRFs and the associated 90% confidence intervals estimated for both states. Focusing on output, we see that the confidence bands overlap substantially. This result suggests that the reaction of output to a fiscal shock is not necessarily stronger if the economy is slack. This finding is in line with some recent results put forth by Valerie Ramey and coauthors (see Ramey, 2011b; Owyang, Ramey, and Zubairy, 2013; Ramey and Zubairy, 2014), which are obtained with a different identification strategy (fiscal spending news shocks constructed following Ramey, 2011b) and methodology (local projections

500 different stochastic realizations of our GIRFs, then store the median realization. We repeat these steps until 500 initial conditions (drawn by allowing for repetitions) associated to recessions (expansions) are considered. Then, we construct the distribution of our GIRFs by considering these 500 median realizations. Our Appendix provides details on the algorithm we employed to compute the GIRFs.

[22]The standard deviation of the news variable employed in the sample is 0.19 according to our linear model, 0.21 conditional on our framework under recessions, and 0.18 under expansions. While being theoretically size-dependent, we verified that the sensitivity of our impulse responses to reasonable changes in the size of the shock is negligible.

à la Jordá, 2005). At a first glance, the evidence seems to be at odds with the impulse response analysis proposed by Auerbach and Gorodnichenko (2012, 2013a), who find a statistically significant difference between the response of output conditional on different states. However, a subtle difference in the construction of the dynamic responses must be considered. Auerbach and Gorodnichenko (2012, 2013a) assume the economy hit by the fiscal shock to start and remain in a recession/expansion for twenty quarters. Differently, here we allow the economic system to switch from a state to another according to the endogenous evolution of the transition indicator. Moreover, the GIRFs plotted in Figure 2.4 are constructed by integrating over all histories belonging to a given state (recessions, expansions). We elaborate on the role played by initial conditions in Section 5.

Quantifying the multipliers. We now turn to the key issue of computing the multipliers and the associated 90% confidence intervals. We compute the "sum" (cumulative) multiplier as the integral of the response of output divided by the integral of the response of fiscal expenditure, i.e., $\sum_{h=1}^{H} Y_h / \sum_{h=1}^{H} G_h$, where H is a chosen horizon. Percent changes are then converted into dollars by rescaling such a ratio by the sample mean ratio of the levels of output over public spending.[23] This measure is designed to account for the persistence of fiscal shocks (Woodford, 2011).

Our results are reported in Table 2.3, where multipliers have been computed considering horizons from one to five years. The evidence clearly speaks in favor of larger (short-run) fiscal spending multipliers in recessions, with values between 3.05 after 8 quarters and 1.00 after 20 quarters. The point-estimates of our multipliers in expansions are substantially lower (from 0.33 to -2.27 after 8 and 20 quarters,

[23] Ramey and Zubairy (2014) warn against this practice by noticing that, in a long U.S. data sample spanning the 1889-2011 period, the output-over-public spending ratio varies from 2 to 24 with a mean of 8. Hence, the choice of a constant value for such ratio may importantly bias the estimation of the multipliers. In our sample, the mean value of such a ratio is 6, and it varies from 5.39 to 6.76. Hence, the commonly adopted *ex-post* conversion from the estimated elasticities to dollar increases does not appear to be an issue for our exercise. The average value of the output-public spending ratio in our sample in 5.81 in NBER recessions, and 6.02 in NBER expansions. Our results are robust to the employment of state-dependent output-public spending ratios.

respectively). The multipliers under recession are statistically larger than one in the short run (i.e., for the first four quarters)

Are multipliers statistically bigger in recessions? We answer this question by constructing a test based on the difference between the multiplier estimated under recessions and expansions. Such a test is constructed to account for the correlation between the estimated state-dependent multipliers.[24] Figure 2.5 plots the distribution of the difference for both measures of multipliers (peak, sum) and for a range of horizons of our impulse responses along with 90% confidence bands. Evidence in favor of state-dependent multipliers would be gained if zero were not included in the confidence bands. In all cases, although marginally, the difference turns out to be not different from a statistical standpoint.[25]

The stabilizing effects of anticipated fiscal shocks. Our STVAR allows also to estimate the impact of government spending shocks on the probability of being in a recession for each given horizon of interest after the shock. Figure 2.6 plots the estimated transition function implied by our model, $\widehat{F(z)}$, along with the 90% confidence bands. The Figure gives interesting information about the estimated impact of a positive government spending shock on the likelihood of remaining in the same phase of the business cycle. Looking at the behavior of the $\widehat{F(z)}$ under recession, we notice that the fiscal shock leads to a clear drop in the probability of remaining in recession. Given the large uncertainty surrounding the response of output to a fiscal shock, different paths of $\widehat{F(z)}$ are admittedly possible. However, the median indication clearly suggests a quick fall of such a probability under the threshold value $\overline{F} = 0.85$ just after five quarters, which is exactly the average duration of a NBER recession in the sample. In terms of

[24] In short, we compute differences of our multipliers in recessions vs. expansions conditional on the same set of draws of the stochastic elements of our model as well as the same realizations of the coefficients of the vector. The empirical density of the difference between our multipliers is based on 500 realizations of such differences for each horizon of interest.

[25] Importantly, our results are not driven by the systematic component of our STVAR *per se*. In other words, in absence of fiscal interventions, our model economy does not deliver large negative accumulated multipliers at longer forecast horizons when starting in expansions. This was verified by simulating a deterministic version of the STVAR, in which only initial conditions are responsible for the different evolution of the variables in recessions and expansions. Our simulations confirm that our cumulated multipliers are indeed driven by the interaction between fiscal shocks and the systematic component of our STVARs.

the econometric methodology employed to estimate the state-dependent effect of government spending shocks on output, this evidence shows the importance of allowing for the possibility of switching from one phase of the business cycle to another. Unsurprisingly, given its expansionary effect, the probability of falling into a recession after the news shock when starting from an expansions is basically zero, though such a probability is quite imprecisely estimated.

2.5 Fiscal multipliers in presence of "extreme" events

Extreme events analysis. So far, our analysis has focused on the possible state-dependence of output reactions to fiscal news shocks and fiscal multipliers, finding weak evidence in favor of countercyclical spending multipliers. The next question we address is whether evidence of nonlinearities might arise when recessions and expansions are "extreme" events. We then re-compute the GIRFs by randomizing over different subsets of histories associated to recessions and expansions. We label "deep" recessions/"strong" expansions the histories associated to realizations of the transition variable which are below/above two standard deviations. Given that our transition variable is standardized, this amounts to saying that all historical realizations of z above two are associated to a strong expansion, while all realizations below minus two are associated to a deep recession. This criterion leads us to isolate four realizations in deep recessions corresponding to the recent great recession (2008Q4-2009Q3) and three realizations which belong to the "strong" expansions category (1983Q4-1984Q2). In a complementary fashion, mild recessions/weak expansions are associated to histories consistent with realizations of the transition variable below/above the threshold value $\overline{z} = -0.75$ but within the range $[-2, 2]$. We then re-compute the GIRFs by randomizing over histories within each of these four sub-categories.

Figure 2.7 shows the GIRFs obtained by distinguishing between "deep" and "mild" recessions and "strong" and "weak" expansions. The estimated GIRFs show that the response of output is roughly proportional to the strength of the recession (expansion). Although in the short-run the response of output in the case of a "mild" recession is very similar to the response of output in a "deep" recession, the response of output is much more persistent at longer horizons when conditioning on the latter case. This, however, cannot be immediately turned into evidence about multipliers, since the persistence in output response might be driven by the persistence of government spending.

Table 2.4 reports the fiscal multipliers estimated in the four different cases under scrutiny. Interestingly, multipliers are still larger in recessions relative to expansions, regardless of the strength of the recession (expansion). When the economy is in a deep recession, we find 4-year horizon multipliers to be 1.6. A similar figures can be gauged for mild recessions, where government spending is found to be expansionary after up to four years. In strong expansions, short-run (one-year) multipliers are slightly above one, but they take negative value at longer horizons. Interestingly, while the difference between mild recessions and weak expansions might seem minimal, the impact of fiscal policy in these two states is much more dramatic. Such a difference may be interpreted in light of the different response of fiscal revenues in the two states (at least in the short-run). In good times, government receipts are found to increase after the shock, while in bad times they are found to decrease. In other words, our VAR suggests that recessions are associated to deficit-financed increases in public spending, while expansions are associated to increases in fiscal spending which are readily financed via an increase in revenues. Hence, recessions are associated with a higher net present value of the fiscal deficit relative to expansions. This can justify the large and positive real effects of fiscal news on the output multiplier if, during recessions, the Ricardian equivalence does not hold because of, say, binding liquidity constraints during recessions, of rule-of-thumb consumers. It can also offer a rationale for the negative multipliers in strong expansions, which is a state associated with a clearly positive response of revenues to fiscal spending shocks.[26]

[26]See Barro and Redlick (2011) for a discussion of deficit-financed versus balanced-budget fiscal multipliers.

Turning to multipliers in expansions, while our point estimates suggest values above one in the short-run, 90% confidence bands imply that we cannot reject values lower than unity. A possible interpretation of large short-run multipliers in expansions relates to the zero lower bound, which has been in place even after the end of the 2007-09 recession, hence in a period classified as ("weak") expansion in our sample. As shown by Leeper, Richter, and Walker (2012), multipliers may be larger than one when an active fiscal policy is accompanied by a passive monetary policy.[27]

When we turn to statistical difference, a comparison between the multipliers in the case of "deep" recessions and those conditional on "strong" expansions suggests that the confidence bands do not overlap, and point to a strong evidence in terms of nonlinear responses of the economy to an expansionary fiscal shock. Our results are confirmed also by looking at the distribution of the difference between the estimated state-dependent multipliers. As shown in Figure 2.8, the countercyclicality of fiscal multipliers conditional on extreme realizations of the business cycle is supported regardless of the horizon.

In our context, it might be more appropriate to test for the null hypothesis of equal multipliers versus the one-sided alternative of multipliers larger in recessions relative to expansions. Table 2.5 collects the fraction of multipliers that are larger in recessions for both "Normal" (recessions/expansions) and "Extreme" (deep recessions/strong expansions) phases of the business cycle. As before, these numbers are estimated by referring to different initial conditions, all else being equal. Hence, any entry greater than or equal to 90 might be interpreted as evidence in favor of larger multipliers in recessions at a 90% confidence level in the context of a one-sided test. The figures corresponding to the exercises conducted so far refer to the "Baseline" scenario. Under the "Normal" (i.e. all recessions vs. all expansions) case, evidence in favor of countercyclical multipliers is not present

[27] In our sample, the number of quarters associated to expansions by the NBER in which the zero lower bound is in place is 15, i.e., some 14% of all the quarters in expansions according to the NBER, which is a non-negligible share. For an analysis pointing to lower fiscal spending multipliers in a liquidity trap caused by a self-fulfilling state of low confidence in a model with nominal rigidies and a Taylor-type interest rate rule, see Mertens and Ravn (2014).

for all horizons. Differently, the analysis of extreme events robustly points towards larger multipliers during recessions. We postpone the analysis of the robustness of this result to a number of perturbations of the baseline framework to the next Section.

How does the economic system evolve after a fiscal shock hitting during an extreme phase of the business cycle? Figure 2.9 plots the estimated value of the $\widehat{F(z)}$ conditional on the four scenarios. For deep recessions, a sizeable decrease of the probability of remaining in such a state occurs as a consequence of the government spending shock: after about five quarters, the value of $\widehat{F(z)}$ decreases from 1 (the economy is in a recession with probability one) to about 0.5 (the economy is unlikely to be in a recession). This drop is quicker and more substantial than the one estimated in presence of mild recessions, and it is also more precisely estimated. Importantly, this suggests that government spending can be effective in lifting the U.S. economy from a deep recession to an expansionary path. The probability of moving away from a strong expansion is low, and more precisely estimated than the one of drifting away from a weak expansion. However, none of the two suggests a high likelihood of falling into a recession.

Estimated multipliers: Comparison with the literature. Our evidence points to larger multipliers in recessions (around 1.6 for the 4-year horizon), and smaller ones, but still somewhat high in the short-run (slightly larger than 1 after one year), in expansions. Are these multipliers in line with what suggested by the literature? A close look at some recent contributions suggests a positive answer. Auerbach and Gorodnichenko (2012, 2013a) deal with unexpected fiscal shocks in a nonlinear VAR framework and find multipliers in recessions of about 2.5. Bachmann and Sims (2012) control for the effects of business confidence and find the sum and peak multipliers in recessions to be 2.7 and 3.3, respectively. Corsetti, Meier, and Müller (2012) work with a flexible panel of OECD countries that allow them to study the effects of fiscal spending shocks under different scenarios. Conditional on periods of financial strains, they find fiscal spending multipliers to be

2.3 on impact, 2.9 at the peak, and larger than 2 in the medium run.[28] Christiano, Eichenbaum, and Rebelo (2011) work with a medium-scale DSGE model and find a multiplier of 2.3 conditional on the zero-lower bound being in place for one year. Evidence of large multipliers can be found also in linear frameworks which deal with the issue of fiscal foresight. Using Bayesian prior predictive analysis for a battery of closed- and open-economy DSGE models featuring different frictions and policy conducts, Leeper, Traum, and Walker (2011) rationalize fiscal spending multipliers of two or larger. Ben Zeev and Pappa (2014) find a peak multiplier larger than 4. Fisher and Peters (2010), using their measure of excess returns of large U.S. military contractors, find a multiplier of 1.5. The same figure is found by Ricco (2014), who employes a measure of news which accounts for the changes in the composition of the pool of forecasters compiling the SPF questionnaires. Depending on the set of restrictions imposed in their sign restriction-VAR analysis, Canova and Pappa (2011) find the U.S. fiscal multipliers to range between 2 and 4.

Our findings qualify those by Auerbach and Gorodnichenko (2012, 2013a), who suggest that recessions are associated with larger fiscal spending multipliers. As already pointed out, their general conclusion might be driven by the implicit assumption that all recessions are treated like "extreme events" when conducting their impulse response analysis. Our analysis suggests that this may very well be the case. This finding has important implications from a policy perspective too, given that a fiscal stimulus may be needed exactly in correspondence to deep recessions.

Overall, our analysis based on "disaggregated" recessions and expansions shows that nonlinearities are likely to arise when we look *within* each of the two states typically investigated in a business cycle context, i.e., recessions and expansions. In particular, we find support in favor of a larger fiscal multiplier when deep recessions are considered.

[28] As reported in the minutes of the *Economic Policy* Panel Discussion, Giancarlo Corsetti pointed out that financial crises, in their study, are not meant to represent recessions. However, he also added that the multipliers are even larger when one uses macro crisis episodes alone in their panel approach. See *Economic Policy*, 2012, 27(72), p. 562.

2.6 Further investigations

Our baseline analysis suggests that evidence in favor of countercyclical fiscal multipliers is borderline when we condition upon recessions vs. expansions, while it becomes much clearer and solid when conditioning upon extreme events. This Section discusses the solidity of our results to the employment of i) alternative identification strategies; ii) a longer sample; iii) debt; iv) several different VAR specifications.

2.6.1 Identification

Exogeneity of the change in government spending expectations. Our baseline analysis rests on revisions of government spending expectations. Such revisions may in principle be due to shocks other than merely fiscal ones. Suppose that $g_t = \boldsymbol{\delta} \boldsymbol{z}_t + \xi_t$, where $\boldsymbol{z}_t$ is a vector of m indicators of the business cycle (say, output, unemployment, inflation, interest rates), $\boldsymbol{\delta}$ is the vector of loadings relating $\boldsymbol{z}_t$ to g_t, and $\xi_t = \varepsilon_t + \phi_1 \varepsilon_{t-1} + \phi_2 \varepsilon_{t-2} + ... + \phi_n \varepsilon_{t-n}$ is a moving average process modeling the unexpected fiscal shock ε_t as well as the expected ones $\varepsilon_{t-j}, j = 1, ..., n$. Then, $\eta_{13}^g = \sum_{j=1}^{3} (E_t g_{t+j} - E_{t-1} g_{t+j}) = \boldsymbol{\delta} \sum_{j=1}^{3} (E_t \boldsymbol{z}_{t+j} - E_{t-1} \boldsymbol{z}_{t+j}) + \widetilde{\eta}_{13}^g$, where $\widetilde{\eta}_{13}^g = \sum_{j=1}^{3} \phi_j \varepsilon_{t-j}$. In words, systematic revisions of fiscal spending forecasts might be due not only to anticipated fiscal shocks, but also to revisions of other variables' forecasts possibly due to other shocks (technology, financial). We deal with this issue by regressing our measure of fiscal news η_{13}^g on a number of macroeconomic indicators available to professional forecasters when they are asked to form expectations about G: (the sums of forecasts revisions of) real GDP growth, unemployment, GDP deflator inflation, the 3-month Treasury bill rate, and the 10-year Treasury bond rate.[29] Figure 2.10 displays the raw and purged versions of

[29]Forecasts of the debt-to-GDP ratio are not included in the SPF survey. We run further regressions by adding lagged realizations of debt-to-GDP ratio to the regression described in the text. Such measures turn out to be insignificant. The choice of not including the contemporaneous realizations of the debt-to-GDP ratio on the right-hand side of the regression is due to the timing of the Survey of Professional Forecasters (SPF). The questionnaire of such survey is sent to the pool of respondents after the advance report of the national income and product accounts by the Bureau of Economic Analysis (BEA) is released to the public. Hence, the questionnaire

the news variable, denoted by η^g_{13} and $\widetilde{\eta}^g_{13}$ respectively. Two considerations are in order. First, the correlation between these two variables is quite high (0.95). Second, the most extreme realizations, documented in Figure 2.1 and reproposed here, are clearly captured by both variables. Hence, most of the information content of the (unpurged version of the) η^g_{13} variable is likely to come from its genuinely exogenous component. To corroborate this statement, we replace the η^g_{13} variable with its purged version $\widetilde{\eta}^g_{13}$ in our VAR, and re-run our estimations and simulations. Table 2.6 ("$\widetilde{\eta}^g_{13}$ last") collects the results of this exercise for our extreme events analysis.[30] These results, as well as those in Table 2.5 on the difference of the multipliers in extreme business cycle phases, confirm our baseline findings.

Contemporaneous effects of fiscal spending shocks. Another issue affecting our baseline analysis regards the timing of the impact of the news shocks. The baseline vector features a recursive identification scheme in which the news variable is ordered last. This choice aims at purging the movements of the η^g_{13} fiscal variable by accounting for its systematic response to government spending, tax revenues, and output. However, such a choice has an obvious limitation, i.e., output is not allowed to move immediately after the realization of the news shock. We then perform a robustness check by focusing on the four-variate VAR $\boldsymbol{X}_t^{\widetilde{\eta}^g} = [\widetilde{\eta}^g_{13,t}, G_t, T_t, Y_t]'$, which enables fiscal news shocks to affect output on impact.[31] We run this exercise with our purged measure of anticipated fiscal shocks to control for the systematic movements of fiscal news due to news hitting other macroeconomic indicators, as explained above. Table 2.6 ("$\widetilde{\eta}^g_{13}$ last") documents slightly different, but statistically equivalent, multipliers relative to the baseline. Most importantly, as also documented by Table 2.5, we find again larger multipliers in deep recessions than in strong expansions.

contains the first estimate of GDP and its components for the *previous* quarter. Thus, in formulating and submitting their projections, the information sets of the SPF panelists include the data reported in the advance report and related to quarter $t-1$ but not data regarding quarter t. For information on the variables included in the survey and the information set possessed by respondents, see http://www.philadelphiafed.org/research-and-data/real-time-center/survey .

[30] Multipliers computed by considering a four-year time span. Similar results are obtained when considering a two-year time span.

[31] An alternative, not pursued here, would be to work with sign restrictions. For an analysis of sign restrictions in fiscal VARs and their implications for the implied fiscal elasticities, see Caldara and Kamps (2012).

2.6.2 Longer sample

The nonlinear estimator we employ is data intensive. Because of limited data availability for the SPF forecast revisions, our baseline analysis rests on a relatively short sample, i.e., 1981Q3-2013Q1. Hence, small-sample issues may lead to distortions of our estimated coefficients, which could then lead us to obtain biased multipliers. We then conduct a robustness check by employing a much longer sample, i.e., 1947Q1-2013Q1. To do so, we use an updated version of Ramey (2011b) widely known fiscal news variable (available at Valerie Ramey's website), and put it first in a VAR including fiscal spending, fiscal revenues, and output. Following Ramey (2011b), we estimate a VAR with four lags and a quadratic trend. Table 2.6 ("Long sample, Ramey's news") collects the outcome of our estimations. Reassuringly, this exercise produces multipliers very much in line with our baseline ones, and it offers support to the importance of looking at extreme events to find nonlinearities in the fiscal multipliers even in long samples.

2.6.3 The role of debt

Our baseline VAR does not feature debt. However, controlling for debt fluctuations in our regressions is important to better understand the drivers of our countercyclical multipliers. The reason is simple. Recent panel-data studies have shown that countries with "high" levels of debt have smaller multipliers than countries with lower levels of debt (see, e.g., Corsetti, Meier, and Müller, 2012; Ilzetzki, Mendoza, and Végh, 2013). Hence, it could in principle be possible that the nonlinearities we have found are driven by different levels of debt rather than different phases of the business cycle. It is then of interest to check if the relevant initial conditions could be related to different degrees of fiscal distress. To this aim, we modify our baseline vector along two dimensions. First, we include the debt/GDP ratio in our VAR. Following a common modeling choice in the literature (see, among others, Leeper, Traum, and Walker, 2011; Corsetti, Meier, and Müller, 2012; Leeper, Walker, and Yang, 2013), we assume the debt/GDP ratio to affect the fiscal instruments with a lag, and put it last in the vector. Second, we employ our debt/GDP ratio as the variable which dictates the switch from a regime to another. This second modification is exactly aimed at capturing the idea of different "debt-contingent" regimes. To discriminate between "high" vs. "low" realizations of debt, we focus on the cyclical component of the debt/GDP ratio, which is extracted from the raw series (in log) by applying a standard Hodrick-Prescott filter with smoothing weight equal to 1,600. Realizations of the debt/GDP ratio one standard deviation above (below) the HP-trend are interpreted as phases of "high" ("low") debt. Positive (negative) realizations within one standard deviation are classified as "moderately high" ("moderately low"). A possible interpretation of this series is that of a "debt/GDP gap" computed by considering a time-varying debt/GDP target, which may be consistent with the clear upward-trending behavior displayed by this ratio in our sample.

Table 2.6 ("Debt/GDP ratio") collects the multipliers produced by this exercise. We distinguish between extreme phases of "high" and "low" fiscal distress, as well as intermediate ones, i.e. "moderately high" and "moderately low", which we indicate with "$Mod.^{+}$ *debt*" and "$Mod.^{-}$ *debt*", respectively. Our results point to fairly similar fiscal multipliers when computed conditional on "high" vs. "low" debt levels. Hence, countercyclical fiscal multipliers do not seem to be guided by the "fiscal cycle".[32] Our results echo those by Favero and Giavazzi (2012), who also find no major empirical differences in a fiscal model for the U.S. when adding debt. It is important to stress, however, that this conclusion is not inconsistent with cross-country studies which point to relevant nonlinearities of fiscal policy effects due to different levels of debt, in particular for developing countries.

2.6.4 Further robustness checks

Our results are robust to a variety of further perturbations of our baseline model, which include: i) a "FAST-VAR" (Factor Augmented Smooth Transition-VAR)

[32] An analysis conducted by adding the debt-to-GDP ratio to our otherwise baseline framework while keeping the moving average of real GDP as our transition indicator returned multipliers very similar to our baseline ones.

version of our VAR model, which we estimate to further control for nonfundamentalness as suggested by Forni and Gambetti (2014a); ii) the estimation of a five-variate VAR featuring the sum of forecast revisions regarding future real GDP as first variable in the vector, again to control for revisions of real GDP forecasts; iii) the employment of revisions over total spending forecasts (as opposed to Federal spending only); iv) a measure of news which accounts for the changes in the composition of the pool of forecasters compiling the SPF questionnaires as in Ricco (2014).[33] The solidity of our baseline results is confirmed also by this battery of robustness checks, which is available upon request.

2.7 Conclusions

This paper quantifies the fiscal spending multiplier in the U.S. and tests the theoretical prediction of a larger reaction of output to fiscal shocks in economic downturns. Following Gambetti (2012a) (2012; 2012) and Forni and Gambetti (2014a), we tackle the issue of non-fundamentalness due to fiscal foresight by identifying anticipated government spending shocks via sums of forecasts revisions collected by the Survey of Professional Forecasters. We show that such a measure of fiscal spending news carries relevant information to predict the future evolution of fiscal expenditures and Granger-causes other measures of fiscal news recently proposed in the literature. Then, we augment a macro-fiscal nonlinear VAR with this measure of fiscal news and estimate the size of fiscal spending multipliers across different phases of the business cycle.

Our empirical investigation points to fiscal multipliers larger than one in recessionary periods. However, conditional on a standard "recession vs. expansion" classification of the phases of the U.S. business cycle, our results do not support the idea of a countercylical fiscal multiplier. Differently, when we condition the estimates of the fiscal multipliers on the *strength* of the business cycle (namely,

[33]We thank Giovanni Ricco for providing us with his measure of fiscal news.

when we distinguish between deep and mild recessions, and weak and strong expansions), we find that fiscal multipliers are statistically larger in deep recessions relative to strong expansionary periods.

The results of our paper highlight the relevance of the different initial economic conditions *within* each of the two states typically considered for classifying the U.S. business cycle. Fiscal multipliers may very well be larger when a fiscal shock occurs in presence of a deep recession like that of 2007-09 than when it occurs in presence of milder economic downturns. Our results imply that a correct measurement of the fiscal multipliers can be performed just if flexible-enough econometric models are put at work.

TABLE 2.1: Anticipated fiscal spending shocks: Statistical relevance

News	(1,3)	(1,1)	(2,2)	(3,3)	(0,0)
$p - value$	**0.00**	**0.00**	**0.00**	**0.00**	0.11

Notes: P-values related to the exclusion Wald-test of one period-lagged News variables entering (one at a time) a regression involving government spending (dependent variable), a constant, three lags of government spending, three lags of fiscal receipts, and three lags of real GDP. Figures in bold are associated to a predictive power of news found to be significant at a 10 percent confidence level. News are expressed in cumulated terms to have an order of integration comparable to that of the other variables. Estimation conducted by considering Newey-West standard errors robust to heteroskedasticity and serial correlation.

TABLE 2.2: News á la Ramey vs. forecast revisions: Granger-causality tests

Sample	*Ramey*	η^g_{13}	*ORZ*	η^g_{13}
1981:III-2008:IV	0.44	**0.06**		
1986:IV-2008:IV	0.28	**0.02**		
1981:III-2010:IV			0.71	**0.06**
1986:IV-2010:IV			0.59	**0.02**

Notes: 'Ramey' stands for the news variable employed by Ramey (2011), 'ORZ' stands for its updated version employed by Owyang, Ramey, and Zubairy (2013). P-values related to the exclusion Wald-test of one period-lagged covariate of interest. Figures in bold are associated to a predictive power of news found to be significant at a 10 percent confidence level. Results based on a bivariate VAR with one lag. Null hypothesis: Column variable does not Granger cause the alternative news measure.

TABLE 2.3: Fiscal spending multipliers

	Sum	
Horizon/State	*Expansion*	*Recession*
4	1.73 [0.52,3.50]	3.15 [1.71,4.27]
8	0.33 [−1.05,2.77]	3.05 [0.68,4.70]
12	−0.57 [−2.24,1.54]	2.13 [0.13,3.82]
16	−1.41 [−3.96,0.74]	1.54 [−0.42,2.95]
20	−2.27 [−6.23,−0.01]	1.00 [−0.94,2.47]

Notes: Figures conditional on our baseline VAR analysis. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

Table 2.4: Fiscal spending multipliers: Extreme events

	Sum			
Hor./State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
4	1.03 [−0.51,2.03]	3.42 [2.05,4.35]	1.69 [0.64,3.40]	3.09 [1.71,4.14]
8	−0.26 [−2.01,0.84]	3.42 [1.22,5.14]	0.30 [−0.87,2.83]	2.94 [0.56,4.46]
12	−1.32 [−3.68,−0.03]	2.21 [0.61,3.54]	−0.62 [−2.15,1.48]	2.06 [0.03,3.78]
16	−2.26 [−5.63,−0.78]	1.60 [0.18,2.63]	−1.40 [−3.91,0.65]	1.38 [−0.48,3.02]
20	−3.28 [−7.00,−1.56]	1.09 [−0.31,2.07]	−2.37 [−6.08,0.01]	0.83 [−0.97,2.54]

Notes: Figures conditional on our VAR analysis with GIRFs conditional on four different sets of initial conditions. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

Table 2.5: Fiscal spending multipliers: Shares of multipliers larger in recessions

	Sum					
Scenario/Horizon	*Cycle*	h=4	h=8	h=12	h=16	h=20
Baseline	*Normal*	84.8	91.6	93.6	95.4	96.6
	Extreme	100	100	100	100	100
$\tilde{\eta}^g_{13}$ *last*	*Normal*	78.2	86.4	89.4	90.6	92.6
	Extreme	100	100	100	100	100
$\tilde{\eta}^g_{13}$ *first*	*Normal*	58.2	76.2	82.2	89.8	92.0
	Extreme	71.6	93.0	97.8	98.8	99.2
Long sample (Ramey's news)	*Normal*	82.8	89.6	87.6	86.4	86.6
	Extreme	90.2	92.8	92.8	93.0	93.6

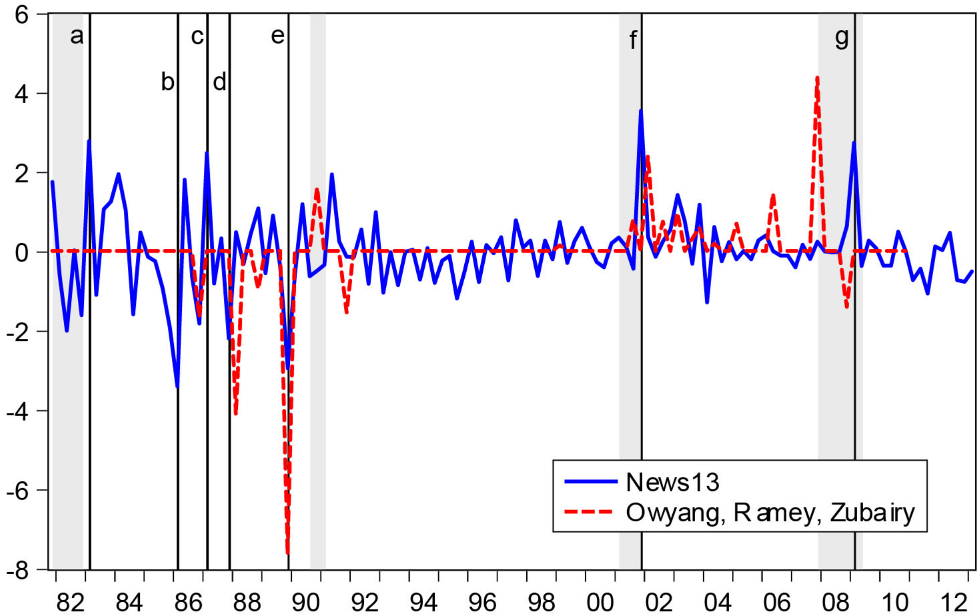
Notes: Normal scenarios- Fraction of multipliers which are larger in recessions than expansions out of 500 draws from their empirical distributions. Extreme scenarios- Fraction of multipliers which are larger in deep recessions than strong expansions out of 500 draws from their empirical distributions. 'h' identifies the number of quarters after the shock.

TABLE 2.6: Fiscal spending multipliers: Extreme events. Different Scenarios

	Sum			
Scenario/State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
Baseline	−2.26 [−5.63,−0.78]	1.60 [0.18,2.63]	−1.40 [−3.91,0.65]	1.38 [−0.48,3.02]
$\widetilde{\eta}^g_{13}$ *last*	−1.57 [−2.92,−0.91]	2.28 [1.23,3.10]	−0.44 [−1.97,2.29]	2.16 [0.22,3.00]
$\widetilde{\eta}^g_{13}$ *first*	−0.70 [−2.50,0.43]	2.36 [0.99,4.29]	0.66 [−1.04,2.90]	2.50 [0.59,4.39]
Long sample (Ramey's news)	0.15 [−0.24,0.53]	1.74 [0.08,3.92]	0.07 [−1.23,0.96]	1.52 [0.60,4.62]
	High debt	*Mod.$^+$ debt*	*Mod.$^-$ debt*	*Low debt*
Debt/GDP ratio	0.68 [0.15,1.37]	0.74 [−1.02,1.15]	1.33 [0.95,1.66]	1.33 [0.81,1.97]

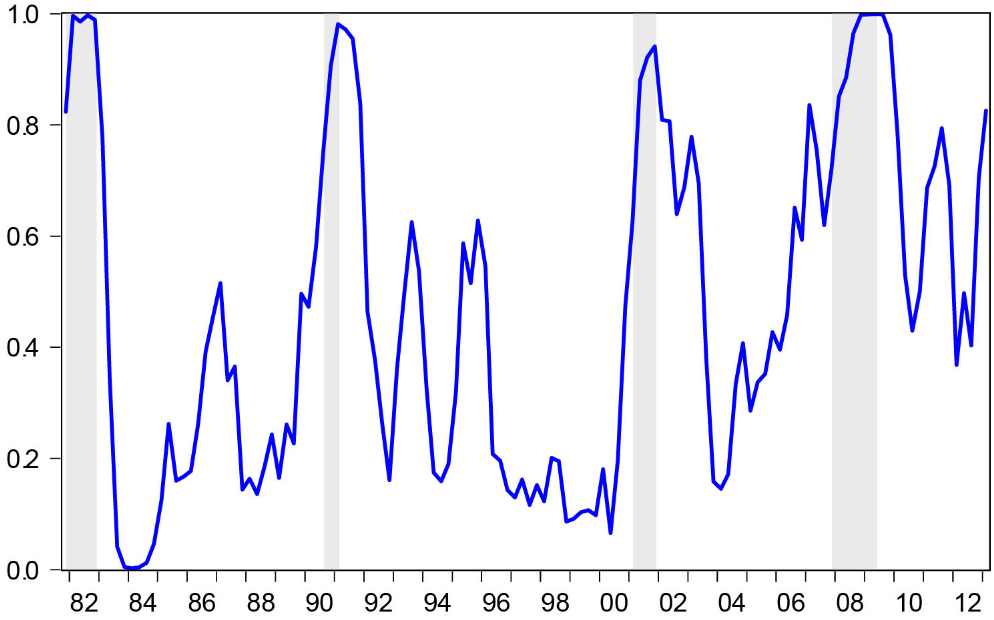
Notes: Four-year integral multipliers. Figures conditional on our VAR analysis with GIRFs conditional on four different sets of initial conditions. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

FIGURE 2.1: News13 (this paper) vs. Owyang, Ramey, and Zubairy's (2013) news variable



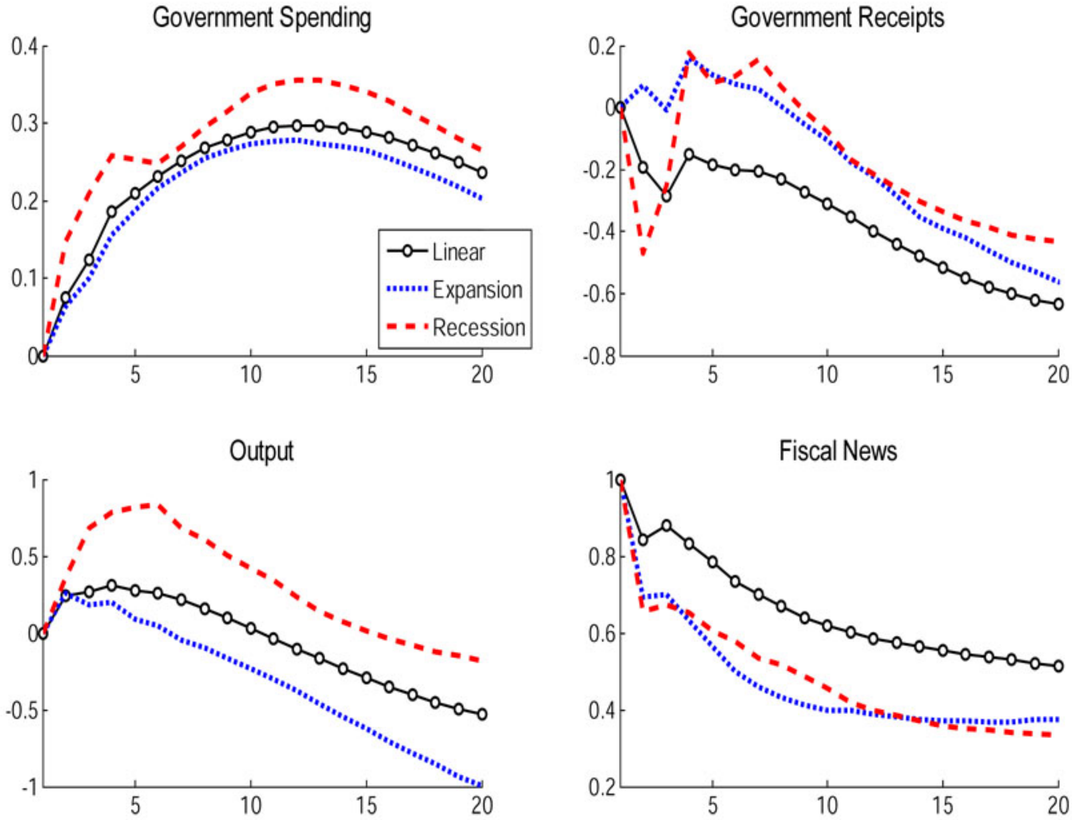
Notes: Blue, solid line: News variable constructed by considering the sum of Survey of Professional Forecasters' forecast revisions regarding future public spending from one-to-three quarter-ahead. Extreme values, interpretation: (a) 1983Q1: Reagan's "Evil Empire" and "Star Wars" speeches; (b) 1986Q1: Perestrojka; (c) 1987Q1: Senate elections won by the Democrats a quarter before; (d) 1987Q4: Spending cuts as for the Pentagon; (e) 1989Q4: Berlin wall; (f) 2001Q4: War in Afghanistan; (g) 2009Q1: Obama's stimulus package. Red, dashed line: News variable constructed by Owyang, Ramey, and Zubairy (2013), who extended Ramey's (2011) news variable up to 2010Q4. Ramey's (2011) variable is constructed by considering the present discounted value of expected changes in defense spending (nominal spending divided by nominal GDP one period before). Both news measures in this Figure are standardized.

FIGURE 2.2: Probability of being in a recessionary phase



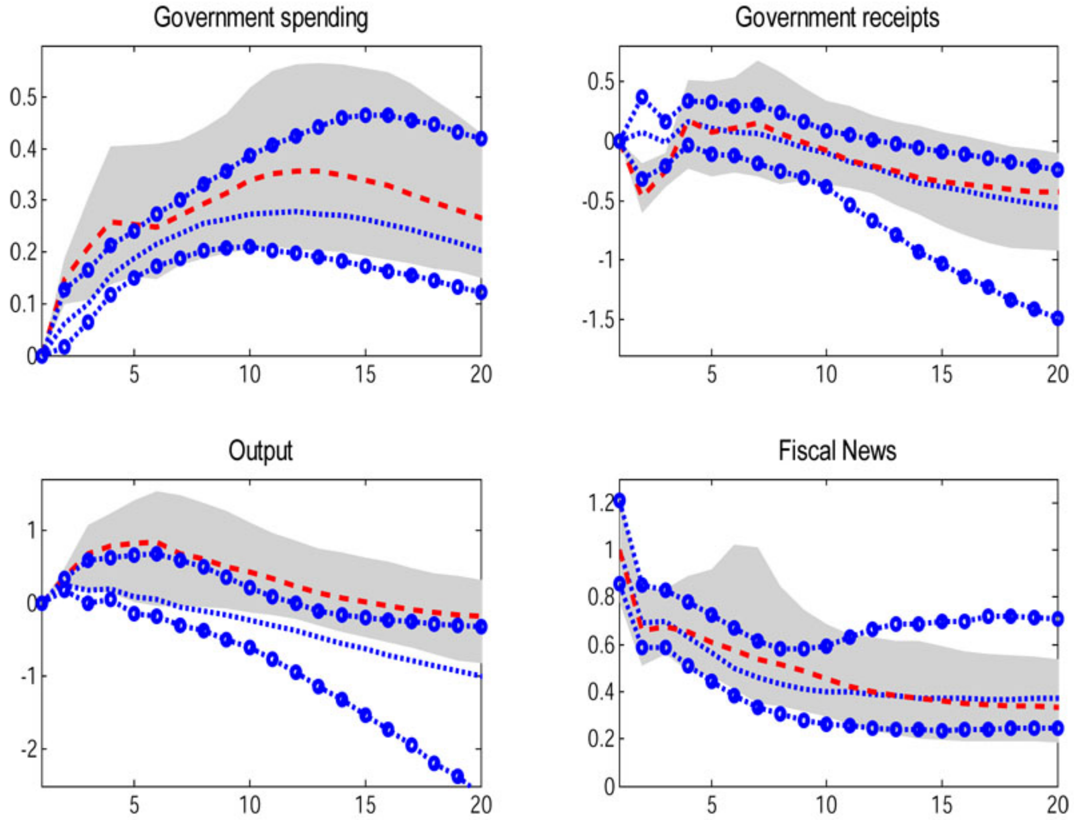
Notes: $F(z)$ computed according to the logistic function presented in the text. Transition variable: Standardized backward-looking moving average constructed with four realizations of the quarter-on-quarter real GDP growth rate. Value of the slope parameter: 2.3.

FIGURE 2.3: Generalized impulse responses to a fiscal news (anticipated) spending shock: Linear model, recessions, expansions.



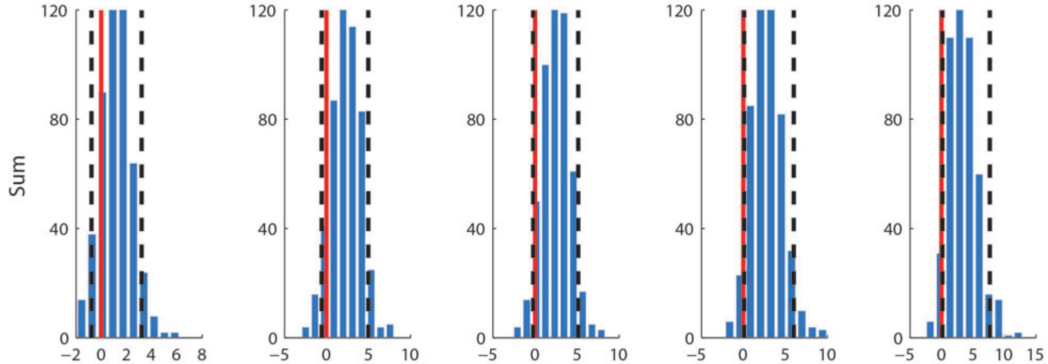
Notes: Median responses to a fiscal news shock normalized to one. News variable constructed as the sum of the revisions of the one, two, and three step-ahead expectation values over future fiscal spending growth. News variable expressed in cumulated terms to have the same order of integration as the one of the log-real variables in the vector. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

FIGURE 2.4: Generalized impulse responses to a fiscal news (anticipated) spending shock: Recessions vs expansions.



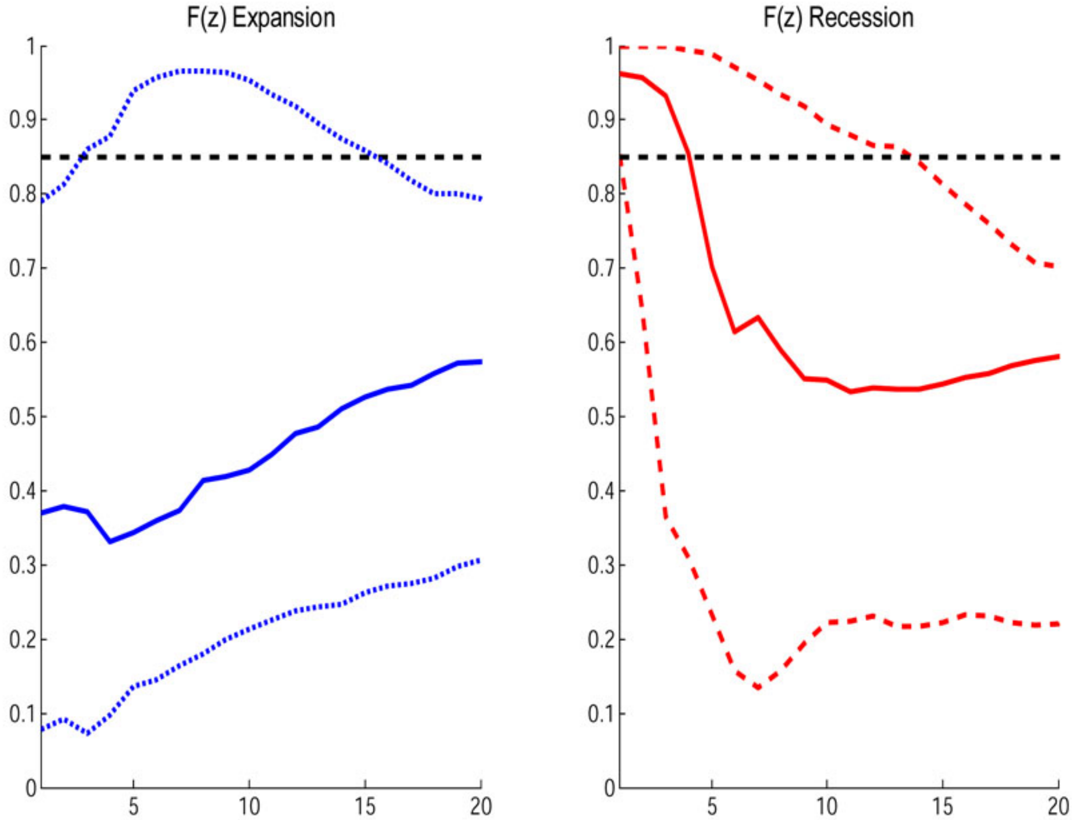
Notes: Median responses to a fiscal news shock normalized to one. 90 percent confidence intervals identified with gray areas (recessions) and circled lines (expansions). Red dashed lines: Recessions. Dotted blue lines: Expansions. News variable constructed as the sum of the revisions of the one, two, and three step-ahead expectation values over future fiscal spending growth. News variable expressed in cumulated terms to have the same order of integration as the one of the log-real variables in the vector. Sample 1981Q3-2013Q1. VAR models estimated with a constant and three lags. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

FIGURE 2.5: Difference in multipliers between recessions and expansions: All histories



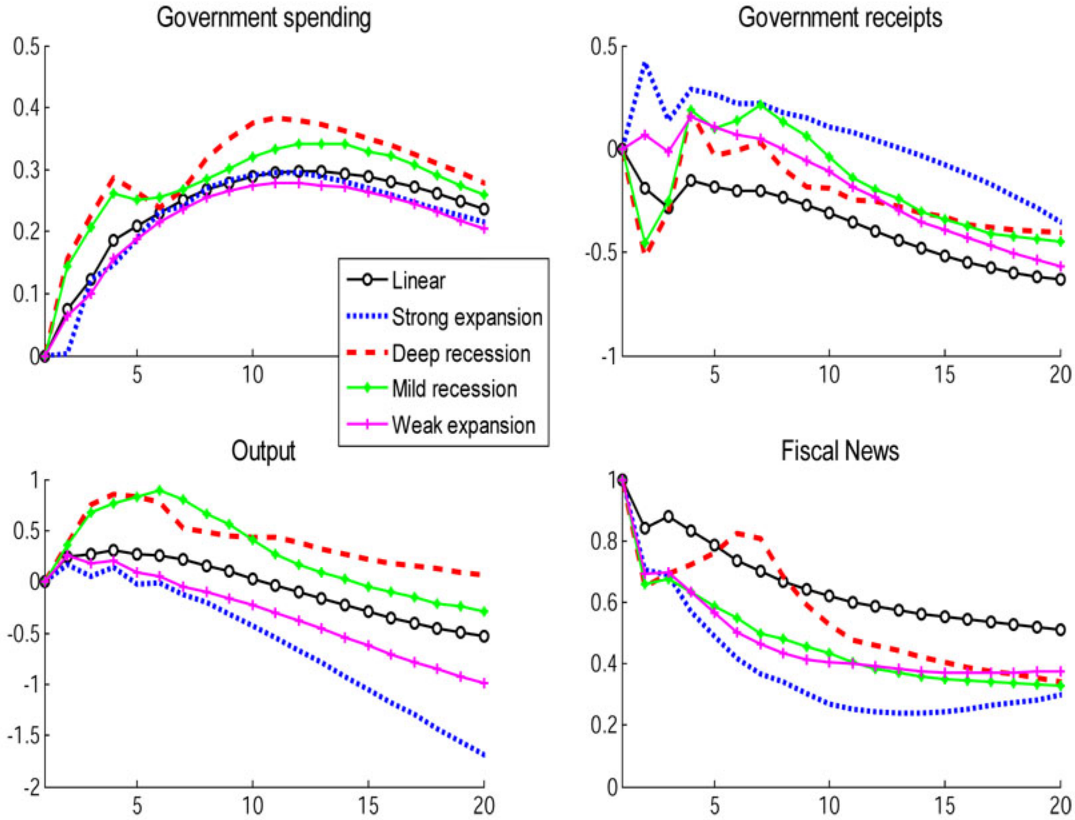
Notes:Empirical densities of the differences computed as multipliers in recessions minus multipliers in expansions. Densities constructed by considering all recessions and expansions (initial conditions) present in the sample. Multipliers conditional on the same set of draws of the stochastic elements of our STVAR model as well as the same realizations of the coefficients of the vector. Densities based on 500 realizations of such differences per each horizon of interest. 'h' identifies the number of quarters after the shock.

FIGURE 2.6: Evolution of the probability of being in a recessionary phase F(z) consistent with our GIRFs



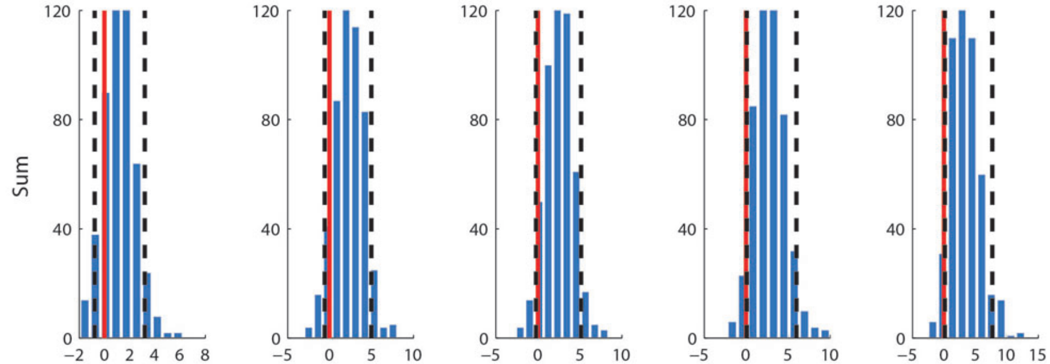
Notes: Solid lines: Median reactions. Blue dotted/red dashed lines: 90 percent confidence intervals. Black dashed horizontal line: Threshold value to switch from a regime to another. Probability computed according to the logistic function presented in the text and the evolution of output conditional on a fiscal news shock. Transition variable: Standardized backward-looking moving average constructed with four realizations of the quarter-on-quarter real GDP growth rate. Value of the slope parameter: 2.3.

FIGURE 2.7: Generalized impulse responses to a fiscal news (anticipated) spending shock: Linear model, deep vs. mild recessions, strong vs. weak expansions



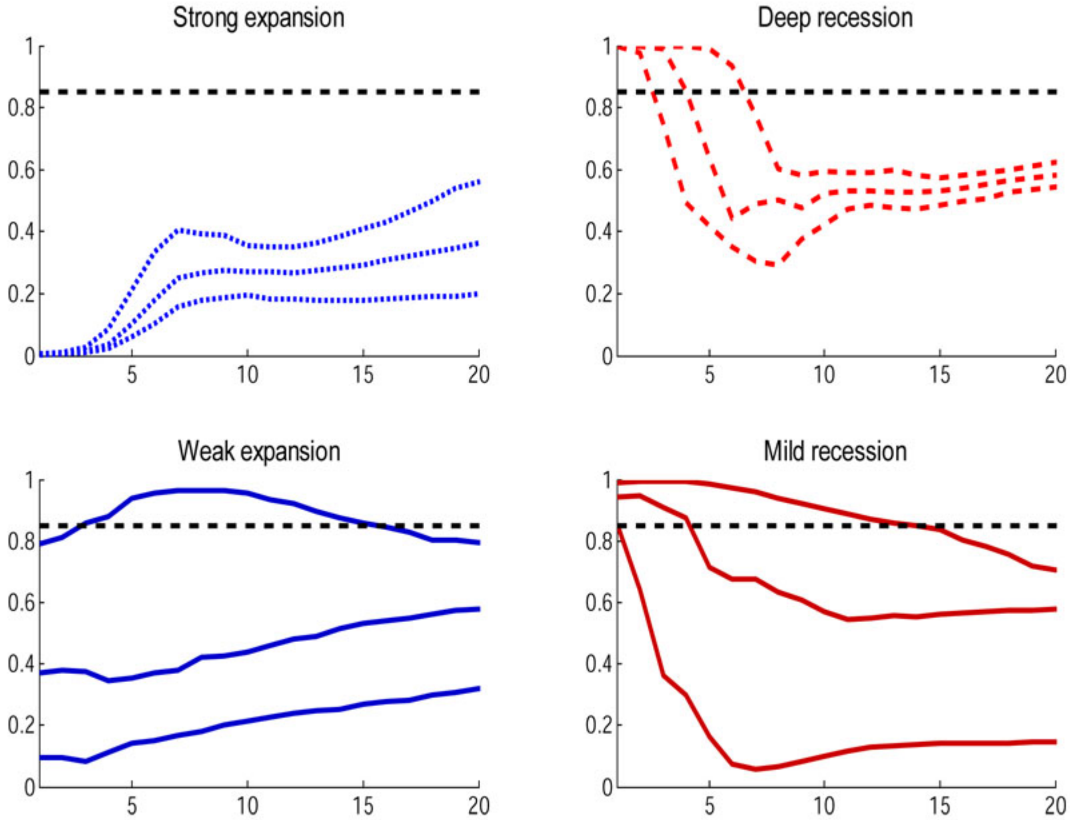
Notes: Deep recessions/strong expansions associated to histories consistent with realizations of our transition variable which are below/above two standard deviations. Mild recessions/weak expansions associated to histories consistent with realizations of our transition variable below/above -0.75 but within the range [-2,2]. Median responses to a fiscal news shock normalized to one. News variable constructed as the sum of the revisions of the one, two, and three step-ahead expectation values over future fiscal spending growth. News variable expressed in cumulated terms to have the same order of integration as the one of the log-real variables in the vector. VAR models estimated with a constant and three lags. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

FIGURE 2.8: Difference in multipliers between recessions and expansions: Extreme events



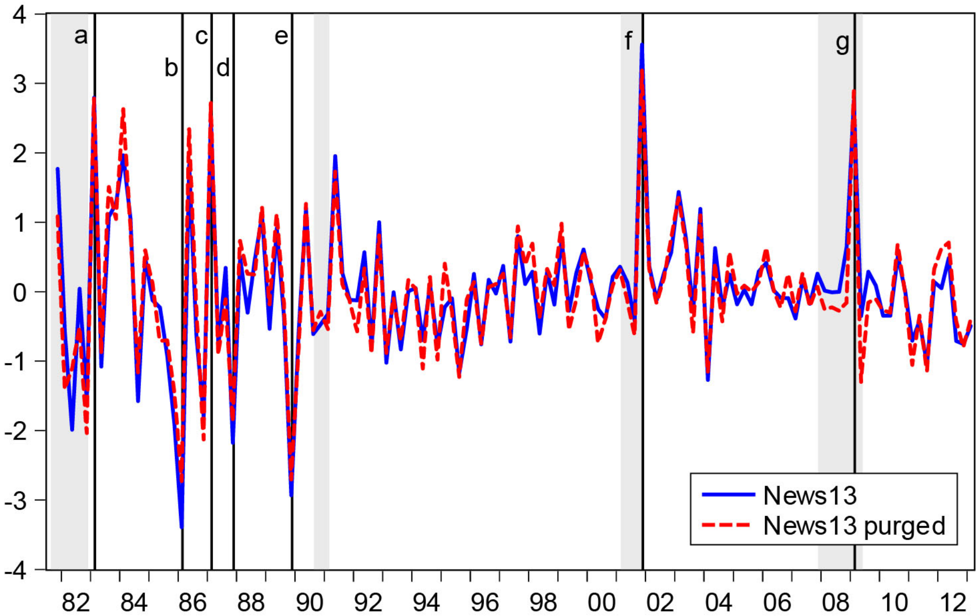
Notes: Empirical densities of the differences computed as multipliers in recessions minus multipliers in expansions. Densities constructed by considering just extreme realizations of recessions and expansions (initial conditions) present in the sample. Multipliers conditional on the same set of draws of the stochastic elements of our STVAR model as well as the same realizations of the coefficients of the vector. Densities based on 500 realizations of such differences per each horizon of interest. 'h' identifies the number of quarters after the shock.

FIGURE 2.9: Evolution of the probability of being in a recessionary phase F(z) consistent with our GIRFs: Extreme events



Notes: Median reactions and 90 percent confidence intervals. Black dashed horizontal line: Threshold value to switch from a regime to another. Deep recessions/strong expansions associated to histories consistent with realizations of our transition variable which are below/above two standard deviations. Mild recessions/weak expansions associated to histories consistent with realizations of our transition variable below/above -0.75 but within the range [-2,2]. Probability computed according to the logistic function presented in the text and the evolution of output conditional on a fiscal news shock. Transition variable: Standardized backward-looking moving average constructed with four realizations of the quarter-on-quarter real GDP growth rate. Value of the slope parameter: 2.3.

FIGURE 2.10: News13 vs. News13 purged



Notes: Blue, solid line: News variable constructed by considering the sum of Survey of Professional Forecasters' forecast revisions regarding future public spending from one to three period-ahead. Red, dashed line: News variable constructed by regressing News13 over a constant and the sums of the forecasts revisions of real GDP growth, unemployment, GDP deflator inflation, the three-month Treasury bill rate, and the 10-year Treasury bond rate. Extreme values, interpretation: (a) 1983Q1: Reagan's "Evil Empire" and "Star Wars" speeches; (b) 1986Q1: Perestrojka; (c) 1987Q1: Senate elections won by the Democrats a quarter before; (d) 1987Q4: Spending cuts as for the Pentagon; (e) 1989Q4: Berlin wall; (f) 2001Q4: War in Afghanistan; (g) 2009Q1: Obama's stimulus package. Both news measures in this Figure are standardized

Appendix 2A

This Appendix reports further details on non-fundamentalness in fiscal SVARs and the role of expectations revisions, the estimation of our nonlinear VARs, the computation of the Generalized Impulse Responses, a number of robustness checks not included in the paper and the computation of the factors employed in one of our robustness checks.

Non-fundamentalness and the role of expectations revisions

Structural VARs have been extensively employed to recover the impulse responses of key macroeconomic variables to fiscal shocks. The implicit assumption when working with SVARs is that their VMA representations are invertible in the past, or in other words that they are fundamental Wold representations of the vector of interest. When such conditions are met, the econometrician has the same information set as the economic agents and can recover the structural shocks by conditioning the VAR estimates on past and current observables.

Fiscal foresight and non-fundamentalness. It is well known, however, that in presence of fiscal foresight (and news shocks in general), this assumption may not hold and fundamental shocks to fiscal policy cannot be recovered from past and current observations. The non-fundamentalness is due to the different discount patterns employed by agents and the econometrician: while the agents attach a larger weight to realizations of the shock occurring in the past, the econometrician discounts in the usual way, and attach lower weights to past observations compared to more recent ones, the reason being that the econometrician's information set lags that of the agents (Leeper, Walker, and Yang, 2013). Hence, in presence of a non-fundamental process, an econometrician not endowed with a large enough information set will not be able to correctly recover the impulse response function of a variable of interest to the structural shock.

How severe is the non-fundamentalness problem? As pointed out by Sims (2012) and Beaudry and Portier (2014), the answer to this question depends on the very same process(es) one wants to model. In terms of fiscal shocks, Leeper, Walker, and Yang (2013) convincingly show that when non-fundamentalness holds the magnitude of the error is quite severe. They employ two DSGE models of the business cycle - a calibrated RBC model and an estimated DSGE model with a number of nominal and real frictions á la Smets and Wouters (2007) - to quantify the mistake an econometrician makes when failing to model fiscal foresight. They show that fiscal multipliers may turn out to be off by hundreds of percent, and can even get the wrong sign.[34] Moreover, Forni and Gambetti (2010b) and Ramey (2011b) show that government spending shocks estimated with standard fiscal VARs can be predicted, evidence supporting the case for non-fundamentalness.

VAR analysis in presence of anticipated shocks. In this section, we propose a framework to fix ideas about the relationship between fiscal foresight and non-fundamentalness and to discuss how the problem can be tackled. To this aim, consider the model

$$y_t = \delta E_t y_{t+1} + g_t + \omega_t \tag{2A.1}$$

$$g_t = \varepsilon_{t-h} + \phi_1 \varepsilon_{t-h-1} + \ldots + \phi_{q-h} \varepsilon_{t-q} = \Phi(L)\varepsilon_t \tag{2A.2}$$

where $|\delta| < 1, \phi_i > 0 \; \forall i, h \geq 0, q \geq h$. The forward-looking process y_t - say, output measured as log-deviations from its trend - is affected by the exogenous stationary process g_t - say, a fiscal shock - plus a random shock ω_t, which is assumed to capture non-fiscal spending shocks affecting output and which is assumed to be *i.i.d.* with zero mean and unit variance. The process (2A.2) features an unanticipated contemporaneous shock ε_t as well as anticipated shocks ε_{t-h} for $h > 0$, where h is the number of foresight periods. The latter are known in advance by rational agents, i.e., agents foresee fiscal moves occurring h-periods ahead. The process g_t

[34] Leeper, Walker, and Yang (2013) model fiscal foresight associated to tax policies. Schmitt-Grohe and Uribe (2012) find government spending shocks anticipated up to eight quarters to be responsible of about 60% of the overall variability of government spending.

is a news-rich process if $|\phi_i| > 1$ for at least one $i > 0$ (Beaudry and Portier, 2014). In all cases, $\{\varepsilon_{t-j}\}_{j=h}^{q}$ is said to be fundamental for g_t if the roots of the polynomial $\Phi(L)$ lie outside the unit circle (Hansen and Sargent, 1991). Importantly, if the g_t process is non-fundamental, its structural shock is not recoverable by employing current and past realizations of g_t only. Consequently, its impulse response to an anticipated shock as well as the dynamic responses of other variables – in this example, y_t – will not be correctly recovered by estimating a VAR in y_t and g_t. For simplicity, and without loss of generality, consider the case in which the unanticipated component is zero, i.e., $h > 0$. We assume that agents have rational expectations and observe news shocks without noise.[35] To begin with, consider the case $h = q = 1$, so that[36]

$$g_t = \varepsilon_{t-1}.$$

Under rational expectations, the solution for the process y_t reads

$$y_t = \delta\varepsilon_t + \varepsilon_{t-1} + \omega_t. \tag{2A.3}$$

The VMA representation of the vector (y_t, g_t) is:

$$\begin{bmatrix} y_t \\ g_t \end{bmatrix} = \underbrace{\begin{bmatrix} \delta & 1 \\ 0 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} 1 & 0 \\ 1 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix}. \tag{2A.4}$$

The VMA representation (2A.4) is fundamental if all the roots of $|\sum_{i=0}^{q} A_i z^i|$ in absolute value lie outside the unit circle. It is easy to verify that in this case the condition is not met, since one gets $|z| = 0$. Hence, in this economic system, inference based on an estimated VAR which includes y_t and g_t only would be incorrect.

[35] Forni, Gambetti, Lippi, and Sala (2013) investigate the case in which economic agents deal with noisy news. Agents are assumed to receive signals regarding the future realization of TFP shocks. Since such signals are noisy, agents react not only to genuinely informative news, but also to noise shocks that are unrelated to economic fundamentals. They find that such noise shocks explain about a third of the variance of output, consumption, and investment. We leave the quantification of the role of noise shocks in the fiscal context to future research.

[36] This process is termed "degenerated news-rich process" by Beaudry and Portier (2014). For an application, see Féve, Matheron, and Sahuc (2009).

Importantly, if a variable η_t added to the econometrician's information set contains "enough" information about the structural shock ε_t, then the VMA representation becomes invertible and the non-fundamentalness issue is circumvented (Giannone and Reichlin, 2006; Sims, 2012; Beaudry and Portier, 2014; Forni and Gambetti, 2014b). Based on this argument, a way to tackle the issue of non-fundamentalness is to include in the VAR a variable which is informative about the effects that news shocks exert on the endogenous variables of interest.[37] In the case of fiscal foresight, then, one has to find a measure of anticipated fiscal spending shocks to correctly gauge the reaction of output to such shocks. It is easy to show that, in the context of model (2A.4), replacing g_t with its one-step-ahead forecast, i.e. $E_t g_{t+1}$, leads to a fundamental VMA representation for the vector $(y_t, E_t g_{t+1})$:

$$\begin{bmatrix} y_t \\ E_t g_{t+1} \end{bmatrix} = \underbrace{\begin{bmatrix} \delta & 1 \\ 1 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} 1 & 0 \\ 0 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix}.$$

This can be seen by verifying that $|A_0 + A_1 z| \neq 0$, $\forall z$.

It is important to notice that expectations *per se* do not necessarily provide a correct measure of fiscal shocks. Consider the case $h = 1$ and $q = 2$, so that

$$g_t = \varepsilon_{t-1} + \phi_2 \varepsilon_{t-2}. \tag{2A.5}$$

The VMA representation for (y_t, g_t) is:

$$\begin{bmatrix} y_t \\ g_t \end{bmatrix} = \underbrace{\begin{bmatrix} \delta\,(1+\delta\phi_2) & 1 \\ 0 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} 1+\delta\phi_2 & 0 \\ 1 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix} + \underbrace{\begin{bmatrix} \phi_2 & 0 \\ \phi_2 & 0 \end{bmatrix}}_{A_2} \begin{bmatrix} \varepsilon_{t-2} \\ \omega_{t-2} \end{bmatrix}, \tag{2A.6}$$

[37] Alternative ways of dealing with this issue have been proposed in the literature. Lippi and Reichlin (1993) propose to use Blaschke matrices to "flip" the roots that are outside the unit circle in order to recover the fundamental representation of the process of interest. Alessi, Barigozzi, and Capasso (2011) and Forni and Gambetti (2014b) propose to augment the VAR with information coming from factors extracted from large datasets. However, in the context of fiscal foresight, non-fundamentalness has a clearly detectable cause, i.e., omitted information due to the absence in the VAR of an informative measure regarding (variations concerning) future fiscal spending moves (Lippi and Reichlin, 1993), (Beaudry and Portier, 2014). Hence, a direct, fiscal-related way of tackling the presence of foresight appears to be desirable.

which is non-fundamental since the roots of $|A_0 + A_1 z + A_2 z^2|$ are $z_1 = 0$ and $|z_2| = \phi_2^{-1}$. In this case, adding the one-step-ahead forecast of g_t does not solve the problem. The VMA representation for the vector $(y_t, E_t g_{t+1})$ is given by:

$$\begin{bmatrix} y_t \\ E_t g_{t+1} \end{bmatrix} = \underbrace{\begin{bmatrix} \delta(1+\delta\phi_2) & 1 \\ 1 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} 1+\delta\phi_2 & 0 \\ \phi_2 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix} + \underbrace{\begin{bmatrix} \phi_2 & 0 \\ 0 & 0 \end{bmatrix}}_{A_2} \begin{bmatrix} \varepsilon_{t-2} \\ \omega_{t-2} \end{bmatrix},$$

which is non-fundamental if $|\phi_2| > 1$.

The role of forecast revisions. Expectation *revisions* help solving the problem. Consider the variable $\eta_t = E_t g_{t+1} - E_{t-1} g_{t+1}$. The VMA representation for the vector (y_t, η_t) is given by:

$$\begin{bmatrix} y_t \\ \eta_t \end{bmatrix} = \underbrace{\begin{bmatrix} \delta(1+\delta\phi_2) & 1 \\ 1 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} 1+\delta\phi_2 & 0 \\ 0 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix} + \underbrace{\begin{bmatrix} \phi_2 & 0 \\ 0 & 0 \end{bmatrix}}_{A_2} \begin{bmatrix} \varepsilon_{t-2} \\ \omega_{t-2} \end{bmatrix},$$

which is fundamental, since $|A_0 + A_1 z + A_2 z^2| \neq 0, \forall z$. It can recursively be shown that expectations revisions of the form $E_t g_{t+1} - E_{t-1} g_{t+1}$ help tackling the issue of non-fundamentalness for any $q > h = 1$.

However, when $h > 1$ is unknown, even expectation revisions are not of help. Consider for example the process:

$$g_t = \varepsilon_{t-2} + \phi_3 \varepsilon_{t-3}.$$

This is not an unlikely case, given that typically the implementation lag for fiscal policy decisions is longer than one quarter. The VMA representation for the vector (y_t, g_t) is:

$$\begin{bmatrix} y_t \\ g_t \end{bmatrix} = \underbrace{\begin{bmatrix} \delta^2(1+\delta\phi_3) & 1 \\ 0 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} \delta(1+\delta\phi_3) & 0 \\ 0 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix}$$
$$+ \underbrace{\begin{bmatrix} 1+\delta\phi_3 & 0 \\ 1 & 0 \end{bmatrix}}_{A_2} \begin{bmatrix} \varepsilon_{t-2} \\ \omega_{t-2} \end{bmatrix} + \underbrace{\begin{bmatrix} \phi_3 & 0 \\ \phi_3 & 0 \end{bmatrix}}_{A_3} \begin{bmatrix} \varepsilon_{t-3} \\ \omega_{t-3} \end{bmatrix},$$

and the roots of $|A_0 + A_1 z + A_2 z^2 + A_3 z^3|$ are $z_{1,2} = 0, |z_3| = \phi_3^{-1}$. Using expectations revisions as before is in this case uninformative, since $E_t g_{t+1} - E_{t-1} g_{t+1} = 0$.

Knowing exactly the number of anticipation periods h would solve the problem, since $E_t g_{t+2} - E_{t-1} g_{t+2} = \varepsilon_t$. However, h is typically unknown. To solve this issue, Gambetti (2012a) proposes to use an alternative, more general measure of expectations revisions, i.e., the news variable defined as:

$$\eta_{1J}^g = \sum_{j=1}^{J} \left(E_t g_{t+j} - E_{t-1} g_{t+j} \right),$$

with J large enough to ensure that $J \geq h$. It can be shown that setting $J \geq 2$ leads to a fundamental representation associated with the vector (y_t, η_{1J}^g), since $\eta_{12}^g = \varepsilon_t$, $\eta_{13}^g = (1+\phi_3)\,\varepsilon_t$ and so on. In our example, if $J = 2$, the VMA representation for (y_t, η_{12}^g) is:

$$\begin{bmatrix} y_t \\ \eta_{12}^g \end{bmatrix} = \underbrace{\begin{bmatrix} \delta^2(1+\delta\phi_3) & 1 \\ 1 & 0 \end{bmatrix}}_{A_0} \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} + \underbrace{\begin{bmatrix} \delta(1+\delta\phi_3) & 0 \\ 0 & 0 \end{bmatrix}}_{A_1} \begin{bmatrix} \varepsilon_{t-1} \\ \omega_{t-1} \end{bmatrix}$$
$$+ \underbrace{\begin{bmatrix} 1+\delta\phi_3 & 0 \\ 0 & 0 \end{bmatrix}}_{A_2} \begin{bmatrix} \varepsilon_{t-2} \\ \omega_{t-2} \end{bmatrix} + \underbrace{\begin{bmatrix} \phi_3 & 0 \\ 0 & 0 \end{bmatrix}}_{A_3} \begin{bmatrix} \varepsilon_{t-3} \\ \omega_{t-3} \end{bmatrix},$$

where the determinant of $|A_0 + A_1 z + A_2 z^2 + A_3 z^3| \neq 0, \forall z.$[38]

[38] It is important to notice that, though related in spirit, Perotti (2011) variable $(E_t g_t - E_{t-1} g_t) + (E_t g_{t+1} - E_{t-1} g_{t+1})$ is uninformative in a case like this, because it does not contain any valuable information about ε_t, i.e., it is equal to zero. The reason is that the forecast horizon covered by such a variable is too short.

In general, when the period of foresight h is unknown or uncertain, the solution would be to include in the VAR a measure of expectations revisions taken over a long enough horizon:

$$\sum_{j=1}^{J}(E_t g_{t+j} - E_{t-1}g_{t+j}) = \eta_{1J}^{g} = \sum_{j=1}^{J}(E_t g_{t+j} - E_{t-1}g_{t+j}) = \begin{cases} (1+\phi_1+...+\phi_{J-h})\,\varepsilon_t \text{ if } J<q \\ (1+\phi_1+...+\phi_{q-h})\,\varepsilon_t \text{ if } J\geqslant q \end{cases} \tag{2A.7}$$

(where $\phi_0 = 0$), which correctly identifies the news shock if $J \geqslant h$.

Estimation of the nonlinear VARs

Consider the model (9)-(12). Its log-likelihood reads as follows:[39]

$$\log L = const + \frac{1}{2}\sum_{t=1}^{T}\log|\mathbf{\Omega}_t| - \frac{1}{2}\sum_{t=1}^{T}\boldsymbol{u}_t'\mathbf{\Omega}_t^{-1}\boldsymbol{u}_t \tag{A1}$$

where the vector of residuals $\boldsymbol{u}_t = \boldsymbol{X}_t -$ $(1-F(z_{t-1})\mathbf{\Pi}_E\boldsymbol{X}_{t-1} - F(z_{t-1})\mathbf{\Pi}_R\boldsymbol{X}_{t-1}$. Our goal is to estimate the following parameters $\mathbf{\Psi} = \{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E, \mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$, where $\mathbf{\Pi}_j(L) = \begin{bmatrix} \mathbf{\Pi}_{j,1} & ... & \mathbf{\Pi}_{j,p} \end{bmatrix}$, $j \in \{R, E\}$. The high-non linearity of the model and its many parameters render its estimation with standard optimization routines problematic. Following Auerbach and Gorodnichenko (2012), we employ the procedure described below. Conditional on $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$, the model is linear in $\{\mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$. Then, for a given guess on $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$, the coefficients $\{\mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$ can be estimated by minimizing $\frac{1}{2}\sum_{t=1}^{T}\boldsymbol{u}_t'\mathbf{\Omega}_t^{-1}\boldsymbol{u}_t$. This can be seen by re-writing the regressors as follows.

Let $\boldsymbol{W}_t = \begin{bmatrix} F(z_{t-1})\boldsymbol{X}_{t-1} & (1-F(z_{t-1})\boldsymbol{X}_{t-1} & ... & F(z_{t-1})\boldsymbol{X}_{t-p} & 1-F(z_{t-1})\boldsymbol{X}_{t-p} \end{bmatrix}$ be the extended vector of regressors, and $\mathbf{\Pi} = \begin{bmatrix} \mathbf{\Pi}_R(L) & \mathbf{\Pi}_E(L) \end{bmatrix}$. Then, we can write $\boldsymbol{u}_t = \boldsymbol{X}_t - \mathbf{\Pi}\boldsymbol{W}_t'$. Consequently, the objective function becomes

$$\frac{1}{2}\sum_{t=1}^{T}(\boldsymbol{X}_t - \mathbf{\Pi}\boldsymbol{W}_t')'\mathbf{\Omega}_t^{-1}(\boldsymbol{X}_t - \mathbf{\Pi}\boldsymbol{W}_t').$$

[39] This Section heavily draws on Auerbach and Gorodnichenko (2012) "Appendix: Estimation Procedure".

It can be shown that the first order condition with respect to $\mathbf{\Pi}$ is

$$vec\mathbf{\Pi}' = \left(\sum_{t=1}^{T}\left[\mathbf{\Omega}_t^{-1}\otimes \boldsymbol{W}_t'\boldsymbol{W}_t\right]\right)^{-1} vec\left(\sum_{t=1}^{T}\boldsymbol{W}_t'\boldsymbol{X}_t\mathbf{\Omega}_t^{-1}\right). \tag{A2}$$

This procedure iterates over different sets of values for $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$. For each set of values, $\mathbf{\Pi}$ is obtained and the $logL$ (A1) computed.

Given that the model is highly nonlinear in its parameters, several local optima might be present. Hence, it is recommended to try different starting values for $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$. To ensure positive definiteness of the matrices $\mathbf{\Omega}_R$ and $\mathbf{\Omega}_E$, we focus on the alternative vector of parameters $\mathbf{\Psi} = \{\gamma, chol(\mathbf{\Omega}_R), chol(\mathbf{\Omega}_E), \mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$, where *chol* implements a Cholesky decomposition.

We estimate our nonlinear model by employing the Monte-Carlo Markov-Chain Metropolis-Hastings algorithm proposed by Chernozhukov and Hong (2003). Given a starting value $\mathbf{\Psi}^{(0)}$, the procedure constructs chains of length N of the parameters of our model following these steps:

Step 1. Draw a candidate vector of parameter values $\mathbf{\Theta}^{(n)} = \mathbf{\Psi}^{(n)} + \boldsymbol{\psi}^{(n)}$ for the chain's $n+1$ state, where $\mathbf{\Psi}^{(n)}$ is the current state and $\boldsymbol{\psi}^{(n)}$ is a vector of i.i.d. shocks drawn from $N(0, \mathbf{\Omega}_\Psi)$, and $\mathbf{\Omega}_\Psi$ is a diagonal matrix.

Step 2. Set the $n+1$ state of the chain $\mathbf{\Psi}^{(n+1)} = \mathbf{\Theta}^{(n)}$ with probability $min\left\{1, L(\mathbf{\Theta}^{(n)})/L(\mathbf{\Psi}^{(n)})\right\}$, where $L(\mathbf{\Theta}^{(n)})$ is the value of the likelihood function conditional on the candidate vector of parameter values, and $L(\mathbf{\Psi}^{(n)})$ the value of the likelihood function conditional on the current state of the chain. Otherwise, set $\mathbf{\Psi}^{(n+1)} = \mathbf{\Psi}^{(n)}$.

The starting value $\mathbf{\Theta}^{(0)}$ is computed by working with a second-order Taylor approximation of the model (8)-(11), so that the model can be written as regressing $\boldsymbol{X}_t$ on lags of $\boldsymbol{X}_t$, $\boldsymbol{X}_t z_t$, and $\boldsymbol{X}_t z_t^2$. The residuals from this regression are employed to fit the expression for the reduced-form time-varying variance-covariance matrix of the VAR (see our paper) using maximum likelihood to estimate $\mathbf{\Omega}_R$ and $\mathbf{\Omega}_E$. Conditional on these estimates and given a calibration for γ, we can construct $\mathbf{\Omega}_t$.

Conditional on $\boldsymbol{\Omega}_t$, we can get starting values for $\boldsymbol{\Pi}_R(L)$ and $\boldsymbol{\Pi}_E(L)$ via equation (A2).

The initial (diagonal matrix) $\boldsymbol{\Omega}_\Psi$ is calibrated to one percent of the parameter values. It is then adjusted "on the fly" for the first 20,000 draws to generate an acceptance rate close to 0.3, a typical choice for this kind of simulations (Canova, 2007). We employ $N = 50,000$ draws for our estimates, and retain the last 20% for inference.

As shown by CH, $\overline{\boldsymbol{\Psi}} = \frac{1}{N}\sum_{n=1}^{N}\boldsymbol{\Psi}^{(n)}$ is a consistent estimate of $\boldsymbol{\Psi}$ under standard regularity assumptions on maximum likelihood estimators. Moreover, the covariance matrix of $\boldsymbol{\Psi}$ is given by $\boldsymbol{V} = \frac{1}{N}\sum_{n=1}^{N}(\boldsymbol{\Psi}^{(n)} - \overline{\boldsymbol{\Psi}})^2 = var(\boldsymbol{\Psi}^{(n)})$, that is the variance of the estimates in the generated chain.

Generalized Impulse Response Functions

Once calibrated our VAR with the point estimates obtained via the procedure presented in the previous sub-Section, we compute the Generalized Impulse Response Functions from our STVAR model by following the approach proposed by Koop, Pesaran, and Potter (1996). The algorithm features the following steps.

1. Consider the entire available observations, with sample size $t = 1981Q3,\ldots,2013Q1$, with $T = 123$, and construct the set of all possible histories $\boldsymbol{\Lambda}$ of length $p = 6$:[40] $\{\lambda_i \in \boldsymbol{\Lambda}\}$. $\boldsymbol{\Lambda}$ will contain $T - p + 1$ histories λ_i.

2. Separate the set of all recessionary histories from that of all expansionary histories. For each λ_i calculate the transition variable z_{λ_i}. If $z_{\lambda_i} \leq \overline{z} = -0.75\%$, then $\lambda_i \in \boldsymbol{\Lambda}^R$, where $\boldsymbol{\Lambda}^R$ is the set of all recessionary histories; if $z_{\lambda_i} > -\overline{z} = -0.75\%$, then $\lambda_i \in \boldsymbol{\Lambda}^E$, where $\boldsymbol{\Lambda}^E$ is the set of all expansionary histories.

[40]The choice $p = 6$ is due to the number of moving average terms (four) of our transition variable z_t, which is constructed by considering five realization of the levels of the (log-)real GDP, i.e., four realizations of the growth rates. Moreover, such transition variable enters our STVAR model via the transition probability $F(z_{t-1})$ with one lag.

3. Select at random one history λ_i from the set $\mathbf{\Lambda}^R$. For the selected history λ_i, take $\widehat{\mathbf{\Omega}}_{\lambda_i}$ obtained as:

$$\widehat{\mathbf{\Omega}}_{\lambda_i} = F(z_{\lambda i})\,\widehat{\mathbf{\Omega}}_R + (1 - F(z_{\lambda_i}))\,\widehat{\mathbf{\Omega}}_E, \tag{A3}$$

where $\widehat{\mathbf{\Omega}}_R$ and $\widehat{\mathbf{\Omega}}_E$ are derived from model (8)-(11) estimated over the entire sample. z_{λ_i} is the transition variable calculated for the selected history λ_i.

4. Cholesky-decompose the estimated variance-covariance matrix $\widehat{\mathbf{\Omega}}_{\lambda_i}$:

$$\widehat{\mathbf{\Omega}}_{\lambda_i} = \widehat{\mathbf{C}}_{\lambda_i}\widehat{\mathbf{C}}'_{\lambda_i} \tag{A4}$$

and orthogonalize the residuals to get the structural shocks:

$$\mathbf{e}^{(j)}_{\lambda_i} = \widehat{\mathbf{C}}^{-1}_{\lambda_i}\widehat{\varepsilon}. \tag{A5}$$

5. From $\mathbf{e}_{\lambda_i}$ draw with replacement h four-dimensional shocks and get the vector of bootstrapped shocks

$$\mathbf{e}^{(j)*}_{\lambda_i} = \left\{\mathbf{e}^*_{\lambda_i,t}, \mathbf{e}^*_{\lambda_i,t+1}, \ldots, \mathbf{e}^*_{\lambda_i,t+h}\right\}, \tag{A6}$$

where h is the horizon for the IRFs we are interested in.

6. Form another set of bootstrapped shocks which will be equal to (A6) except for the k_{th} shock in $\mathbf{e}^{(j)*}_{\lambda_i,t}$ which is the shock we want to perturbate (news in our model) by an amount equal to δ. Denote the vector of bootstrapped perturbated shocks by $\mathbf{e}^{(j)\delta}_{\lambda_i}$.

7. Transform back $\mathbf{e}^{(j)*}_{\lambda_i}$ and $\mathbf{e}^{(j)\delta}_{\lambda_i}$ as follows:

$$\widehat{\varepsilon}^{(j)*}_{\lambda_i} = \widehat{\mathbf{C}}_{\lambda_i}\mathbf{e}^{(j)*}_{\lambda_i} \tag{A7}$$

and

$$\widehat{\varepsilon}^{(j)\delta}_{\lambda_i} = \widehat{\mathbf{C}}_{\lambda_i}\mathbf{e}^{(j)\delta}_{\lambda_i}. \tag{A8}$$

8. Use (A7) and (A8) to generate two sequences $\mathbf{X}_{\lambda_i}^{(j)*}$ and $\mathbf{X}_{\lambda_i}^{(j)\delta}$ and get the $GIRF^{(j)}(h, \delta, \lambda_i)$.

9. Conditional on history λ_i, repeat for $j = 1, \ldots, B$ vectors of bootstrapped residuals and get $GIRF^{(1)}(h, \delta, \lambda_i), GIRF^{(2)}(h, \delta, \lambda_i), \ldots, GIRF^{(B)}(h, \delta, \lambda_i)$. Set $B = 500$.

10. Calculate the GIRF conditional on history λ_i as

$$\widehat{GIRF}^{(i)}(h, \delta, \lambda_i) = B^{-1} \sum_{j=1}^{B} GIRF^{(i,j)}(h, \delta, \lambda_i). \tag{A9}$$

11. Repeat all previous steps for $i = 1, \ldots, 500$ randomly drawn histories belonging to the set of recessionary histories, $\lambda_i \in \mathbf{\Lambda}^R$.
Get $\widehat{GIRF}^{(1,R)}(h, \delta, \lambda_{1,R}), \widehat{GIRF}^{(2,R)}(h, \delta, \lambda_{2,R}), \ldots, \widehat{GIRF}^{(500,R)}(h, \delta, \lambda_{500,R})$, where now the subscript R denotes explicitly that we are *conditioning upon recessionary histories.*

12. Take the average and get $\widehat{GIRF}^{(R)}\left(h, \delta, \mathbf{\Lambda}^R\right)$, which is the average GIRF under recessions.

13. Repeat all previous steps - 3 to 12 - for 500 histories belonging to the set of all expansions and get $\widehat{GIRF}^{(E)}\left(h, \delta, \mathbf{\Lambda}^E\right)$.

14. The computation of the 90% confidence bands for our impulse responses is undertaken by picking up, per each horizon of each state, the 5th and 95th percentile of the densities $\widehat{GIRF}^{([1:500],R)}$ and $\widehat{GIRF}^{([1:500],E)}$.

Further robustness checks

Our baseline analysis suggests that evidence in favor of countercyclical fiscal multipliers is borderline when we condition upon recessions vs. expansions, while it becomes much clearer and solid when conditioning upon extreme events. The paper presents the robustness checks conducted by considering a different measure of fiscal spending news (obtained by regressing the baseline fiscal news variable on a constant and a number of controls), a different ordering of the variables in our VAR, the debt/GDP ratio as an extra-variable in our VAR as well as the transition indicator, and a longer sample (an analysis that we conducted by working with Ramey (2011b) indicator of fiscal spending news). Table 2.6 in the paper documents the robustness of our results by collecting multipliers computed over a 4-year horizon. Table 2A.1 in this Appendix confirms the solidity of our results conditional on a 2-year horizon.

We then conduct a variety of robustness checks to verify the solidity of our results. We present the robustness checks below and discuss our results by referring to Table 2A.2, which summarizes the outcome.

FAVAR. Our baseline VAR is meant to parsimoniously model a set of key macroeconomic indicators crucial to quantify fiscal spending multipliers. A further reason to prefer a parsimonious VAR is the somewhat limited number of observations available to construct the measures of forecast revisions we deal with, as well as the nonlinearity of our framework, in which a large number of VAR coefficients is estimated. Despite its advantages, a parsimonious model might suffer from an omitted-variable problem, which may bias the results of our baseline scenario. In particular, reactions of variables like the real interest rate and the real exchange rate may be important for the computation of the fiscal spending multipliers. Interactions between financial variables and real aggregates may also be at work conditional on our fiscal news shock. We tackle this informational insufficiency issue by adding to our VAR a factor extracted from a large dataset, so to purge the (possibly bias-contaminated) estimated shocks. This strategy leads us to deal with a nonlinear version of the Factor-Augmented VAR (FAVAR) model popularized, in the monetary policy context, by Bernanke, Boivin, and Eliasz (2005). In particular, we consider a large dataset composed of 150 time-series, and extract the common factors which maximize the explained variance of such series (a description of the series included in our dataset, their transformations, and the computation of the factors is provided in the Appendix 2A). Following Stock and Watson (2012) in their recent analysis on the drivers of the post-WWII U.S. economy, we extract six common factors and then focus on the fiscal FAVAR

$\boldsymbol{X}_t^{favar} = [f_t^1, G_t, T_t, Y_t, \eta_{13,t}^g]'$, where "$f_t^1$" is the factor explaining the largest share of variance of the series in our enlarged database. Due to the limited number of degrees of freedom, we focus on a VAR model with two lags, a choice that we will keep for all the five-variate VAR we estimate to check the robustness of our baseline results.[41] Results on the difference of the fiscal multiplier in different states of the economy are collected in Table 2A.2 under the label "FAVAR".

Expectation revisions of output. Our baseline results rests on the identifying assumption that our fiscal news variable carries valuable information regarding fiscal shocks which may have led economic agents to revise their expectations of future public spending. However, such revisions may have been undertaken because of "news" about some other shocks. Suppose news about the future evolution of technology become part of agents' information sets between time $t-1$ and t. This might induce agents to revise their expectations regarding future realizations of output. Given the link between output and public spending (due to, e.g., automatic stabilizers), such revisions may induce agents to further revise their expectations of future fiscal spending as well. Hence, revisions of future fiscal spending may be triggered not only by anticipated fiscal shocks, but also by anticipated shocks of a different nature (say, news concerning technology).

We tackle this issue by modeling the five-variate VAR $\boldsymbol{X}_t^Y = [\eta_{13,t}^Y, G_t, T_t, Y_t, \eta_{13,t}^g]'$, where η_{13}^Y stands for the sum of forecast revisions regarding future real GDP. The construction of this variable replicates the construction of η_{13}^g explained in Section 2. We put η_{13}^Y before η_{13}^g in the vector to control for the effects exerted by contemporaneous movements in η_{13}^Y on η_{13}^g.[42] Notice that one can interpret this robustness check as pointing to the role of an identified factor omitted in the baseline analysis, i.e., the role of expectation revisions on output. Table 2A.2 collects our results under the label "η_{13}^Y".

Contemporaneous effects of η_{13}^g shocks. Our approach features a recursive identification scheme. Our choice aims at purging the movements of the

[41] The entire set of results regarding our robustness checks is not documented in this paper to save space, but it is available upon request.

[42] Given the choice of a Cholesky-identification scheme, the ordering of the variables before η_{13}^g is irrelevant for the computation of our impulse responses to a fiscal news shock.

η^g_{13} fiscal variable by accounting for its systematic response to government spending, tax revenues, and output. However, such a choice has an obvious limitation, i.e., output is not allowed to move immediately after the realization of the news shock. We then perform a robustness check by focusing on the five-variate VAR $\boldsymbol{X}_t^{\eta^g} = [\eta^g_{13,t}, \eta^Y_{13,t}, G_t, T_t, Y_t]'$, which enables fiscal news shocks to move output immediately. We keep the measure of news on output to control for the systematic movements of fiscal news due to output news. Notice that this VAR allows for (without forcing) an immediate response of fiscal spending G, which would however be inconsistent with the idea of a news shock. Interestingly, a look at our GIRFs (available upon request) suggest that public spending moves in neither of the two states. This result confirms the potential of the measure of fiscal news shocks employed in this paper to capture anticipated fiscal shocks, i.e., shocks which do not exert an immediate impact on public spending but, possibly, trigger an immediate reaction of output.[43] As for the difference in fiscal multipliers, the results are presented in Table 2A.2 under "η^g_{13} first".

Expectation revisions of total government spending. Our baseline analysis hinges upon a η^g_{13}, which is based on revisions of forecasts over the growth rates of federal spending only. However, expectations concerning levels of future fiscal spending regarding state and local expenditures are also available. We then construct levels of expected total spending and compute the growth rates of such expected realizations. We use this variable as a proxy of the expected growth rates of total fiscal spending that are not readily available in the SPF dataset. We then use this proxy as an alternative to our η^g_{13} variable in our vector. Our results are collected in Table 2A.2 under the label "η^g_{13} total".

Ricco's news indicator. In a recent paper, Ricco (2014) shows that the news variable we employ in our study to account for fiscal foresight may be affected by aggregation bias. Our measure is based on forecast revisions constructed by

[43] Interestingly, our impulse responses suggest that output moves immediately in recessions, while its contemporaneous response is not significant when expansions are considered (IRFs not shown for the sake of brevity, but available upon request). The contemporaneous zero reaction of public spending to changes in output is consistent with the evidence on the zero contemporaneous output elasticity of government spending in the U.S. surveyed by Caldara and Kamps (2012).

appealing to location measures (e.g., mean, median) of the distribution of the forecasts (across forecasters). However, since the composition of the pool of respondents to the SPF changes over time, one problem related with our measure is that use of measures of central tendency might induce a non negligible bias if the distribution of forecast revisions is skewed. The resulting aggregation bias may in principle imply important quantitative effects for the computation of fiscal multipliers. Ricco (2014) circumvents this problem by constructing a measure of news based on the revisions of expectations of each individual forecaster in the pool, whose forecast is available for at least two consecutive quarters. Ex-post aggregation of such revisions gives rise to a "microfounded" measure of aggregate news. Even though the correlation between the two measures of fiscal anticipation in our sample is quite high (it reads 0.84), it is of interest to repeat our exercise by employing Ricco's news measure as an alternative to our η^g_{13}.[44] Results are documented in Table 2A.2 under "η^g_{13} à la Ricco".

Table 2A.2 collects the figures related to the robustness checks discussed above. Two main messages arise. First, the "Normal" scenarios generally points to a rather fragile evidence of countercyclical fiscal multipliers. The most evident exception is the case of the news variable *à la* Ricco, which leads to larger multipliers in recessions. This is in line with the fact that, in presence of a skewed distribution of forecast revisions, our measure of news would downward-bias the estimated fiscal multipliers (see Ricco (2014) for a detailed explanation of the sources of this bias). Second, our extreme events analysis robustly supports larger multipliers in recessions. Hence, our results corroborate a recent statement by Blanchard and Leigh (2013) on the magnitude of fiscal multipliers and the effectiveness of fiscal stabilization policies in periods of substantial economic slack. These results lend support also to Parker's (2011) call for empirical models able to capture the possible countercyclicality of fiscal multipliers.

[44] We thank Giovanni Ricco for providing us with his measure of fiscal news.

Computation of the factors for the FAVAR approach

We follow Stock and Watson (2012) to estimate the factors from a large unbalanced data set of US variables. Let $\boldsymbol{X}_t = (X_{1t}, \ldots, X_{nt})'$ denote a vector of n macroeconomic time series, with $t = 1, \ldots, T$. X_{it} is a single time series transformed to be stationary and to have mean zero. The dynamic factor model expresses each of the n time series as the sum of a common component driven by r unobserved factors $\boldsymbol{F}_t$ plus an idiosyncratic disturbance term e_{it}:

$$\mathbf{X}_t = \mathbf{\Lambda F}_t + \mathbf{e}_t \tag{A10}$$

where $\boldsymbol{e}_t = (e_{1t}, \ldots, e_{nt})'$ and $\mathbf{\Lambda}$ is the $n \times r$ matrix of factor loadings.

The factors are assumed to follow a linear and stationary vector autoregression:

$$\mathbf{\Phi}\left(L\right)\boldsymbol{F}_t = \boldsymbol{\eta}_t \tag{A11}$$

where $\mathbf{\Phi}\left(L\right)$ is a $r \times r$ matrix of lag polynomials with the vector of r innovations $\boldsymbol{\eta}_t$. Stationarity implies that $\mathbf{\Phi}\left(L\right)$ can be inverted and $\boldsymbol{F}_t$ has the moving average representation:

$$\boldsymbol{F}_t = \mathbf{\Phi}\left(L\right)^{-1}\boldsymbol{\eta}_t. \tag{A12}$$

With n large, under the assumption that there is a single-factor structure, simple cross-sectional averaging provides an estimate of $\boldsymbol{F}_t$ good enough to treat $\widehat{\boldsymbol{F}}_t$ as data in a regression without a generated regressor problem. With multiple factors, Stock and Watson (2012) show that a consistent estimate of $\boldsymbol{F}_t$ is obtained using principal components.

Our data set is standard in the recent literature on factor models (see Stock and Watson, 2012; Forni and Gambetti, 2014a). It contains an unbalanced panel of 150 quarterly series, with starting date 1947Q1 and end date 2012Q3. The data are grouped into 12 categories: NIPA variables (31); industrial production (16); employment and unemployment (14); housing starts (6); inventories, orders and sales (12); prices (15); earnings and productivity (13); interest rates (10); money and credit (12); stock prices (5); exchange rates (7); and other (9). Earnings and productivity data include TFP-adjusted measures of capacity utilization introduced by Basu, Fernald, and Kimball (2006). The category labeled "other" includes expectations variables.

The transformation implemented for the series to be stationary with zero mean are reported in Table 2A.3. The factors were estimated using principal components as in Stock and Watson (2012). The assumption that the factors can be estimated with no breaks over the period 1947Q2-2012Q3 is motivated by the findings of Stock and Watson (2012), who show that the space spanned by the factors can be estimated consistently even if there is instability in $\mathbf{\Lambda}$.

Multipliers: "Sum" vs. "Peak" measures.

The multipliers documented in the paper are "sum" multipliers. They are computed as the integral of the response of output divided by the integral of the response of fiscal expenditure, i.e., $\sum_{h=1}^{H} Y_h / \sum_{h=1}^{H} G_h$, where Y_h and G_h represent the impulse responses of output and public spending respectively h-horizon after the shock, and the ratio is then rescaled for the sample mean ratio of the levels of Y over G. This measure is designed to account for the persistence of fiscal shocks (Woodford, 2011). Another measure often employed by the literature (see Stock and Watson, 2012; Forni and Gambetti, 2014a) is the "peak" one, which is calculated as the peak response of output divided by the peak response of fiscal expenditure over the first H horizons, i.e., it is equal to $\frac{max_{h=1,\dots,H}\{Y_h\}}{max_{h=1,\dots,H}\{G_h\}}$. Again, percent changes are then converted into dollars by rescaling such a ratio by the sample mean ratio of the levels of output over public spending.[45] Tables 2A.4-2A.7

[45]Ramey and Zubairy (2014) warn against this practice by noticing that, in a long U.S. data sample spanning the 1889-2011 period, the output-over-public spending ratio varies from 2 to 24 with a mean of 8. Hence, the choice of a constant value for such ratio may importantly bias the estimation of the multipliers. In our sample, the mean value of such a ratio is 6, and it varies from 5.39 to 6.76. Hence, the commonly adopted *ex-post* conversion from the estimated elasticities to dollar increases does not appear to be an issue for our exercise. The average value of the output-public spending ratio in our sample in 5.81 in NBER recessions, and 6.02 in NBER expansions. Our results are robust to the employment of state-dependent output-public spending ratios.

extend the information contained in Tables 2.3-2.6 in the main text, and Figures 2A.1 and 2A.2 extend the one in Figures 2.5 and 2.8.

TABLE 2A.1: Fiscal spending multipliers: Extreme events. Different Scenarios.

	Peak			
Scenario/State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
Baseline	0.79 [0.45,1.09]	2.27 [1.45,2.93]	1.09 [0.72,2.31]	2.72 [1.32,3.96]
$\widetilde{\eta}^g_{13}$ *last*	0.45 [0.20,0.63]	3.37 [2.03,4.34]	1.05 [0.48,3.77]	3.15 [1.50,4.21]
$\widetilde{\eta}^g_{13}$ *first*	1.21 [0.25,1.94]	3.05 [1.84,6.72]	2.17 [0.93,4.97]	3.64 [1.58,6.80]
Long sample (Ramey's news)	0.47 [0.19,0.80]	2.83 [1.56,5.92]	0.68 [0.23,1.56]	2.59 [1.22,6.60]
	High debt	*Mod.$^+$ debt*	*Mod.$^-$ debt*	*Low debt*
Debt/GDP ratio	1.79 [1.62,2.00]	1.35 [0.68,2.15]	1.95 [1.68,2.44]	2.08 [1.54,2.78]

	Sum			
Scenario/State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
Baseline	−2.26 [−5.63,−0.78]	1.60 [0.18,2.63]	−1.40 [−3.91,0.65]	1.38 [−0.48,3.02]
$\widetilde{\eta}^g_{13}$ *last*	−0.42 [−1.56,0.13]	3.65 [2.09,4.99]	0.76 [−0.62,3.86]	3.17 [0.99,4.43]
$\widetilde{\eta}^g_{13}$ *first*	0.76 [−1.02,2.20]	3.95 [1.59,8.72]	2.35 [0.38,5.43]	3.95 [1.27,8.17]
Long sample (Ramey's news)	0.43 [0.06,0.85]	2.49 [0.19,8.66]	0.02 [−1.77,1.08]	2.21 [−0.68,9.72]
	High debt	*Mod.$^+$ debt*	*Mod.$^-$ debt*	*Low debt*
Debt/GDP ratio	2.43 [2.13,2.72]	0.99 [0.36,1.77]	2.29 [1.93,2.59]	2.07 [1.43,2.54]

Notes: Two-year integral multipliers. Figures conditional on our VAR analysis with GIRFs conditional on four different sets of initial conditions. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

TABLE 2A.2: Fiscal spending multipliers: Shares of multipliers larger in recessions

		Peak				
Scenario/Horizon	*Cycle*	$h=4$	$h=8$	$h=12$	$h=16$	$h=20$
Baseline	*Normal*	87.80	90.80	90.00	90.60	90.20
	Extreme	99.60	100.00	100.00	100.00	100.00
FAVAR	*Normal*	87.40	91.00	93.20	93.40	93.40
	Extreme	100.00	99.80	99.60	99.60	99.60
η_{13}^{Y}	*Normal*	62.60	80.60	82.20	84.00	84.80
	Extreme	93.00	99.20	99.40	99.20	99.20
η_{13}^{g} *first*	*Normal*	81.00	86.80	88.60	90.00	90.00
	Extreme	97.60	99.20	99.40	99.60	99.60
η_{13}^{g} *total*	*Normal*	94.60	92.60	92.60	93.20	93.40
	Extreme	100.00	100.00	100.00	100.00	100.00
η_{13}^{g} *à la Ricco*	*Normal*	95.00	94.00	94.00	94.20	94.40
	Extreme	100.00	100.00	100.0	100.00	100.00
		Sum				
Scenario/Horizon	*Cycle*	$h=4$	$h=8$	$h=12$	$h=16$	$h=20$
Baseline	*Normal*	84.80	91.60	93.60	95.40	96.60
	Extreme	100.00	100.00	100.00	100.00	100.00
FAVAR	*Normal*	89.80	85.20	85.60	88.20	89.80
	Extreme	100.00	100.00	100.00	100.00	100.00
η_{13}^{Y}	*Normal*	36.80	73.00	79.80	83.00	86.40
	Extreme	86.20	100.00	100.00	100.00	100.00
η_{13}^{g} *first*	*Normal*	74.20	84.60	88.20	90.40	91.40
	Extreme	96.20	99.80	100.00	100.00	100.0
η_{13}^{g} *total*	*Normal*	89.80	86.60	85.40	85.80	87.00
	Extreme	98.60	95.20	99.00	100.00	100.00
η_{13}^{g} *à la Ricco*	*Normal*	93.00	90.80	90.60	90.20	90.40
	Extreme	99.80	99.80	99.80	99.80	99.80

Notes: Figures conditional on our VAR analysis with GIRFs conditional on four different sets of initial conditions. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

TABLE 2A.3: Time series employed for the computation of the factors

N	Series	Mnemonic	Tr.	Start	End
1	Real Gross Domestic Product, 1 Decimal	GDPC1	5	1947Q1	2012Q3
2	Real Gross National Product	GNPC96	5	1947Q1	2012Q3
3	Real National Income	NICUR/GDPDEF	5	1947Q1	2012Q3
4	Real Disposable Income	DPIC96	5	1947Q1	2012Q3
5	Real Personal Income	RPI	6	1959Q1	2012Q3
6	Nonfarm Business Sector: Output	OUTNFB	5	1947Q1	2012Q3
7	Real Final Sales of Domestic Product, 1 Decimal	FINSLC1	5	1947Q1	2012Q3
8	Real Private Fixed Investment, 1 Decimal	FPIC1	5	1995Q1	2012Q3
9	Real Private Residential Fixed Investment, 1 Decimal	PRFIC1	5	1995Q1	2012Q3
10	Real Private Nonresidential Fixed Investment, 1 Decimal	PNFIC1	5	1995Q1	2012Q3
11	Real Gross Private Domestic Investment, 1 Decimal	GPDIC1	5	1947Q1	2012Q3
12	Real Personal Consumption Expenditure	PCECC96	5	1947Q1	2012Q3
13	Real Personal Consumption Expenditure: Nondurable Goods	PCNDGC96	5	1995Q1	2012Q3
14	Real Personal Consumption Expenditure: Durable Goods	PCDGCC96	5	1995Q1	2012Q3
15	Real Personal Consumption Expenditure: Services	PCESVC96	5	1995Q1	2012Q3
16	Real Gross Private Saving	GPSAVE/GDPDEF	5	1947Q1	2012Q3
17	Real Federal Consumption Expenditures, Gross Investment, 1 Decimal	FGCEC1	5	1995Q1	2012Q3
18	Federal Goverment: Current Expenditures, Real	FGEXPND/GDPDEF	5	1947Q1	2012Q3
19	Federal Goverment: Current Receipts, Real	FGRECPT/GDPDEF	5	1947Q1	2012Q3
20	Net Federal Government Saving	FGDEF	2	1947Q1	2012Q3
21	Government Current Expenditures/GDP Deflator	GEXPND/GDPDEF	5	1947Q1	2012Q3
22	Government Current Receipts/GDP Deflator	GRECPT/GDPDEF	5	1947Q1	2012Q3
23	Government Real Expenditures minus Real Receipts	GDEF	2	1947Q1	2012Q3
24	Real Government Consumption Expenditures, Gross Investment, 1 Decimal	GCEC1	5	1947Q1	2012Q3
25	Real Change in Private Inventories, 1 Decimal	CBIC1	1	1947Q1	2012Q3
26	Real Exports of Goods and Services, 1 Decimal	EXPGSC1	5	1947Q1	2012Q3
27	Real Imports of Goods and Services, 1 Decimal	IMPGSC1	5	1947Q1	2012Q3
28	Corporate Profits After Tax, Real	CP/GDPDEF	5	1947Q1	2012Q3
29	Nonfinancial Corporate Business: Profits After Tax, Real	NFCPATAX/GDPDEF	5	1947Q1	2012Q3
30	Corporate Net Cash Flow, Real	CNCF/GDPDEF	5	1947Q1	2012Q3
31	Net Corporate Dividends, Real	DIVIDEND/GDPDEF	5	1947Q1	2012Q3
32	Industrial Production Index	INDPRO	5	1947Q1	2012Q3
33	Industrial Production: Business Equipment	IPBUSEQ	5	1947Q1	2012Q3
34	Industrial Production: Consumer Goods	IPCONGD	5	1947Q1	2012Q3
35	Industrial Production: Durable Consumer Goods	IPDCONGD	5	1947Q1	2012Q3
36	Industrial Production: Final Products (Market Group)	IPFINAL	5	1947Q1	2012Q3
37	Industrial Production: Materials	IPMAT	5	1947Q1	2012Q3
38	Industrial Production: Nondurable Consumer Goods	IPNCONGD	5	1947Q1	2012Q3
39	Capacity Utilization: Manufacturing	MCUMFN	4	1972Q1	2012Q3
40	Industrial Production: Manufacturing	IPMAN	5	1972Q1	2012Q3
41	Industrial Production: Durable Manufacturing	IPDMAN	5	1972Q1	2012Q3
42	Industrial Production: Mining	IPMINE	5	1972Q1	2012Q3
43	Industrial Production: Nondurable Manufacturing	IPNMAN	5	1972Q1	2012Q3
44	Industrial Production: Durable Materials	IPDMAT	5	1947Q1	2012Q3
45	Industrial Production: Electric and Gas Utilities	IPUTIL	5	1972Q1	2012Q3
46	ISM Manufacturing: PMI Composite Index	NAPM	1	1948Q1	2012Q3
47	ISM Manufacturing: Production Index	NAPMPI	1	1948Q1	2012Q3
48	Average Weekly Hours of Production and Nonsupervisory Employees: Manuf.	AWHMAN	1	1948Q1	2012Q3
49	Average Weekly Overtime Hours of Prod. and Nonsupervisory Employees: Manuf.	AWOTMAN	2	1948Q1	2012Q3
50	Civilian Labor Force Participation Rate	CIVPART	2	1948Q1	2012Q3

Notes: Description of the Table in two pages.

N	Series	Mnemonic	Tr.	Start	End
51	Civilian Labor Force	CLF160V	5	1948Q1	2012Q3
52	Civilian Employment	CE160V	5	1948Q1	2012Q3
53	All Employees: Total Private Industries	USPRIV	5	1947Q1	2012Q3
54	All Employees: Goods-Producing Industries	USGOOD	5	1947Q1	2012Q3
55	All Employees: Service-Providing Industries	SRVPRD	5	1947Q1	2012Q3
56	Unemployed	UNEMPLOY	5	1948Q1	2012Q3
57	Average (Mean) Duration of Unemployment	UEMPMEAN	2	1948Q1	2012Q3
58	Civilian Unemployment Rate	UNRATE	2	1948Q1	2012Q3
59	Index of Help-Wanted Advertising in Newspapers	A0M046	1	1959Q1	2012Q3
60	HOANBS/CNP160V	HOANBS/CNP160V	4	1948Q1	2012Q3
61	Initial Claims	ICSA	5	1967Q3	2012Q3
62	Housing Starts: Total: New Privately Owned Units Started	HOUST	5	1959Q1	2012Q3
63	Housing Starts in Northeast Census Region	HOUSTNE	5	1959Q1	2012Q3
64	Housing Starts in Midwest Census Region	HOUSTMW	5	1959Q1	2012Q3
65	Housing Starts in South Census Region	HOUSTS	5	1959Q1	2012Q3
66	Housing Starts in West Census Region	HOUSTW	5	1959Q1	2012Q3
67	New Private Housing Units Authorized by Building Permits	PERMIT	5	1960Q1	2012Q3
68	US Manufacturers New Orders for Non Defense Capital Goods	USNOIDN.D	5	1959Q2	2012Q3
69	US New Orders of Consumer Goods and Materials	USCNORCGD	5	1959Q2	2012Q3
70	US ISM Manufacturers Survey: New Orders Index SADJ	USNAPMNO	1	1950Q2	2012Q3
71	Retail Sales: Total (Excluding Food Services)	RSXFS	5	1992Q1	2012Q3
72	Value of Manufacturers' Total Inventories for All Manufacturing Industries	UMTMTI	5	1992Q1	2012Q3
73	Value of Manufacturers' Total Inventories for Durable Goods	AMDMTI	5	1992Q1	2012Q3
74	Value of Manufacturers' Total Inventories for Nondurable Goods Industries	AMNMTI	5	1992Q1	2012Q3
75	ISM Manufacturing: Inventories Index	NAPMII	1	1948Q1	2012Q3
76	ISM Manufacturing: New Orders Index	NAPMNOI	1	1948Q1	2012Q3
77	Value of Manufacturers' New Orders for Cons. Goods: Cons. Dur. Goods Ind.s	ACDGNO	5	1992Q1	2012Q3
78	Manuf.s' New Orders: Durable Goods	DGORDER	5	1992Q1	2012Q3
79	Value of Manuf.s' New Orders for Dur. Goods Ind.: Transp. Equipment	ANAPNO	5	1992Q1	2012Q3
80	Gross Domestic Product: Chain-type Price Index	GDPCTPI	5	1947Q1	2012Q3
81	Gross National Product: Chain-type Price Index	GNPCTPI	5	1947Q1	2012Q3
82	Gross Domestic Product: Implicit Price Deflator	GDPDEF	5	1947Q1	2012Q3
83	Gross National Product: Implicit Price Deflator	GNPDEF	5	1947Q1	2012Q3
84	Consumer Price Index for All Urban Consumers: All Items	CPIAUCSL	6	1947Q1	2012Q3
85	Consumer Price Index for All Urban Consumers: All Items Less Food	CPIULFSL	6	1947Q1	2012Q3
86	Consumer Price Index for All Urban Consumers: All Items Less Energy	CPILEGSL	6	1957Q1	2012Q3
87	Consumer Price Index for All Urban Consumers: All Items Less Food & Energy	CPILFESL	6	1957Q1	2012Q3
88	Consumer Price Index for All Urban Consumers: Energy	CPIENGSL	6	1947Q1	2012Q3
89	Consumer Price Index for All Urban Consumers: Food	CPIUFDSL	6	1947Q1	2012Q3
90	Producer Price Index: Finished Goods: Capital Equipment	PPICPE	6	1947Q1	2012Q3
91	Producer Price Index: Crude Materials for Further Processing	PPICRM	6	1947Q1	2012Q3
92	Producer Price Index: Finished Consumer Goods	PPIFCG	6	1947Q1	2012Q3
93	Producer Price Index: Finished Goods	PPIFGS	6	1947Q1	2012Q3
94	Spot Oil Price: West Texas Intermediate	OILPRICE	6	1947Q1	2012Q3
95	Nonfarm Business Sector: Hours of All Persons	HOANBS	5	1947Q1	2012Q3
96	Nonfarm Business Secotr: Output Per Hour of All Persons	OPHNFB	5	1947Q1	2012Q3
97	Nonfarm Business Sector: Unit Nonlabor Payments	UNLPNBS	5	1947Q1	2012Q3
98	Nonfarm Business Sector: Unit Labor Cost	ULCNFB	5	1947Q1	2012Q3
99	Compensation of Employees: Wages and Salary Accruals, Real	WASCUR/CPI	5	1947Q1	2012Q3
100	Nonfarm Business Sector: Compensation Per Hour	COMPNFB	5	1947Q1	2012Q3

Notes: Table 2A.3 (continued). Time series employed for the computation of the factors. Description of the Table in the following page.

N	Series	Mnemonic	Tr.	Start	End
101	Nonfarm Business Sector: Real Compensation Per Hour	COMPRNFB	5	1947Q1	2012Q3
102	Growth in utilization-adjusted TFP	dtfp_util	1	1947Q2	2012Q3
103	Growth in business sector TFP	dtfp	1	1947Q2	2012Q3
104	Utilization in producing investment	du_invest	1	1947Q2	2012Q3
105	Utilization in producing non-investment business output	du_consumption	1	1947Q2	2012Q3
106	Utilization-adjusted TFP in producing equipment and consumer durables	dtfp_I_util	1	1947Q2	2012Q3
107	Utilization-adjusted TFP in producing non-equipment output	dtfp_C_util	1	1947Q2	2012Q3
108	Effective Federal Funds Rate	FEDFUNDS	2	1954Q3	2012Q3
109	3-Month Treasury Bill: Secondary Market Rate	TB3MS	2	1947Q1	2012Q3
110	1-Year Treasury Constant Maturity Rate	GS1	2	1953Q2	2012Q3
111	10-Year Treasury Constant Maturity Rate	GS10	2	1953Q2	2012Q3
112	Moody's Seasoned Aaa Corporate Bond Yield	AAA	2	1947Q1	2012Q3
113	Moody's Seasoned Baa Corporate Bond Yield	BAA	2	1947Q1	2012Q3
114	Bank Prime Loan Rate	MPRIME	2	1949Q1	2012Q3
115	GS10-FEDFUNDS Spread	GS10-FEDFUNDS	1	1954Q3	2012Q3
116	GS1-FEDFUNDS Spread	GS1-FEDFUNDS	1	1954Q3	2012Q3
117	BAA-FEDFUNDS Spread	BAA-FEDFUNDS	1	1954Q3	2012Q3
118	Non-Borrowed Reserves of Depository Institutions	BOGNONBR	5	1959Q1	2012Q3
119	Board of Gov. Total Reserves, Adjusted for Changes in Reserve Requirements	TRARR	5	1959Q1	2012Q3
120	Board of Gov. Monetary Base, Adjusted for Changes in Reserve Requirements	BOGAMBSL	5	1959Q1	2012Q3
121	M1 Money Stock	M1SL	5	1959Q1	2012Q3
122	M2 Less Small Time Deposits	M2MSL	5	1959Q1	2012Q3
123	M2 Money Stock	M2SL	5	1959Q1	2012Q3
124	Commercial and Industrial Loans at All Commercial Banks	BUSLOANS	5	1947Q1	2012Q3
125	Consumer Loans at All Commercial Banks	CONSUMER	5	1947Q1	2012Q3
126	Bank Credit at All Commercial Banks	LOANINV	5	1947Q1	2012Q3
127	Real Estate Loans at All Commercial Banks	REALLN	5	1947Q1	2012Q3
128	Total Consumer Credit Owned and Securitized, Outstanding	TOTALSL	5	1947Q1	2012Q3
129	St. Louis Adjusted Monetary Base	AMBSL (CHNG)	5	1947Q1	2012Q3
130	US Dow Jones Industrials Share Price Index (EP)	USSHRPRCF	5	1950Q2	2012Q3
131	US Standard & Poor's Index of 500 Common Stocks	US500STK	5	1950Q2	2012Q3
132	US Share Price Index NADJ	USI62...F	5	1957Q2	2012Q3
133	Dow Jones/GDP Deflator	DOW Jones/GDPDEF	5	1950Q2	2012Q3
134	S&P/GDP Deflator	S&P/GDPDEF	5	1950Q2	2012Q3
135	Trade Weighted U.S. Dollar Index: Major Currencies	TWEXMMTH	2	1973Q1	2012Q3
136	Euro/U.S. Foreign Exchange Rate	EXUSEU(-1)	5	1999Q1	2012Q3
137	Germany/U.S. Foreign Exchange Rate	EXGEUS	5	1971Q1	2001Q4
138	Switzerland/U.S. Foreign Exchange Rate	EXSZUS	5	1971Q1	2012Q3
139	Japan/U.S. Foreign Exchange Rate	EXJPUS	5	1971Q1	2012Q3
140	U.K./U.S. Foreign Exchange Rate	EXUSUK(-1)	5	1971Q1	2012Q3
141	Canada/U.S. Foreign Exchange Rate	EXCAUS	5	1971Q1	2012Q3
142	US The Conference Board Leading Economic Indicators Index SADJ	USCYLEADQ	5	1959Q1	2012Q3
143	US Economic Cycle Research Institute Weekly Leading Index	USECRIWLH	5	1950Q2	2012Q3
144	University of Michigan Consumer Sentiment: Personal Finances, Current	USUMPFNCH	2	1978Q1	2012Q3
145	University of Michigan Consumer Sentiment: Personal Finances, Expected	USUMPFNEH	2	1978Q1	2012Q3
146	University of Michigan Consumer Sentiment: Economic Outlook, 12 Months	USUMECO1H	2	1978Q1	2012Q3
147	University of Michigan Consumer Sentiment: Economic Outlook, 5 Years	USUMECO5H	2	1978Q1	2012Q3
148	University of Michigan Consumer Sentiment: Buying Conditions, Durables	USUMBUYDH	2	1978Q1	2012Q3
149	University of Michigan Consumer Sentiment Index	USUMCONSH	2	1991Q1	2012Q3
150	University of Michigan Consumer Sentiment - Current Conditions	USUMCNSUR	2	1991Q1	2012Q3

Notes: Table 2A.3 (continued). Time series employed for the computation of the factors. Classification of the series: 1-31: "NIPA"; 32-47: "Industrial Production"; 48-61: "Employment and Unemployment"; 62-67: "Housing Starts"; 68-79: "Inventories", "Orders and Sales"; 80-94: "Prices"; 95-107: "Earnings and Productivity"; 108-117: "Interest Rates"; 118-129: "Money and Credit"; 130-134: "Stock Prices"; 135-141: "Exchange Rates"; 142-150: "Others". The column labeled "Tr." indicates the transformation applied to the series (1 = level, 2 = first difference, 3 = logarithm, 4 = second difference, 5 = first difference of logarithm, 6 = second difference of logarithm). Data source: Federal Reserve Bank of St. Louis' website.

TABLE 2A.4: Fiscal spending multipliers

	Peak		*Sum*	
Horizon/State	*Expansion*	*Recession*	*Expansion*	*Recession*
4	1.68 [1.12,3.49]	3.38 [1.77,4.70]	1.73 [0.52,3.50]	3.15 [1.71,4.27]
8	1.24 [0.80,3.19]	3.32 [1.55,4.91]	0.33 [−1.05,2.77]	3.05 [0.68,4.70]
12	1.11 [0.74,2.69]	2.77 [1.40,4.28]	−0.57 [−2.24,1.54]	2.13 [0.13,3.82]
16	1.09 [0.71,2.43]	2.60 [1.38,3.96]	−1.41 [−3.96,0.74]	1.54 [−0.42,2.95]
20	1.09 [0.71,2.41]	2.58 [1.38,3.90]	−2.27 [−6.23,−0.01]	1.00 [−0.94,2.47]

Notes: Figures conditional on our baseline VAR analysis. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

TABLE 2A.5: Fiscal spending multipliers: Extreme events

	Peak			
Hor./State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
4	1.24 [0.78,1.88]	3.57 [2.14,4.73]	1.68 [1.15,3.44]	3.23 [1.74,4.69]
8	0.86 [0.53,1.25]	3.58 [1.94,4.75]	1.24 [0.82,3.16]	3.24 [1.56,4.72]
12	0.79 [0.48,1.10]	2.39 [1.48,3.30]	1.11 [0.75,2.56]	2.88 [1.32,4.20]
16	0.79 [0.45,1.09]	2.27 [1.45,2.93]	1.09 [0.72,2.31]	2.72 [1.32,3.96]
20	0.79 [0.43,1.08]	2.24 [1.44,2.90]	1.09 [0.72,2.29]	2.71 [1.31,3.94]

	Sum			
Hor./State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
4	1.03 [−0.51,2.03]	3.42 [2.05,4.35]	1.69 [0.64,3.40]	3.09 [1.71,4.14]
8	−0.26 [−2.01,0.84]	3.42 [1.22,5.14]	0.30 [−0.87,2.83]	2.94 [0.56,4.46]
12	−1.32 [−3.68,−0.03]	2.21 [0.61,3.54]	−0.62 [−2.15,1.48]	2.06 [0.03,3.78]
16	−2.26 [−5.63,−0.78]	1.60 [0.18,2.63]	−1.40 [−3.91,0.65]	1.38 [−0.48,3.02]
20	−3.28 [−7.00,−1.56]	1.09 [−0.31,2.07]	−2.37 [−6.08,0.01]	0.83 [−0.97,2.54]

Notes: Figures conditional on our VAR analysis with GIRFs conditional on four different sets of initial conditions. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

TABLE 2A.6: Fiscal spending multipliers: Shares of multipliers larger in recessions

	Peak					
Scenario/Horizon	*Cycle*	h=4	h=8	h=12	h=16	h=20
Baseline	*Normal*	87.8	90.8	90.0	90.6	90.2
	Extreme	100	100	100	100	100
$\tilde{\eta}^g_{13}$ *last*	*Normal*	84.0	87.0	87.8	88.8	89.2
	Extreme	100	100	100	100	100
$\tilde{\eta}^g_{13}$ *first*	*Normal*	69.0	76.2	76.8	79.8	80.6
	Extreme	86.4	96.4	96.2	96.0	96.0
Long sample (Ramey's news)	*Normal*	96.8	98.2	98.0	98.0	98.0
	Extreme	99.0	100	100	100	100
	Sum					
Scenario/Horizon	*Cycle*	h=4	h=8	h=12	h=16	h=20
Baseline	*Normal*	84.8	91.6	93.6	95.4	96.6
	Extreme	100	100	100	100	100
$\tilde{\eta}^g_{13}$ *last*	*Normal*	78.2	86.4	89.4	90.6	92.6
	Extreme	100	100	100	100	100
$\tilde{\eta}^g_{13}$ *first*	*Normal*	58.2	76.2	82.2	89.8	92.0
	Extreme	71.6	93.0	97.8	98.8	99.2
Long sample (Ramey's news)	*Normal*	82.8	89.6	87.6	86.4	86.6
	Extreme	90.2	92.8	92.8	93.0	93.6

Notes: Normal scenarios- Fraction of multipliers which are larger in recessions than expansions out of 500 draws from their empirical distributions. Extreme scenarios- Fraction of multipliers which are larger in deep recessions than strong expansions out of 500 draws from their empirical distributions. 'h' identifies the number of quarters after the shock.

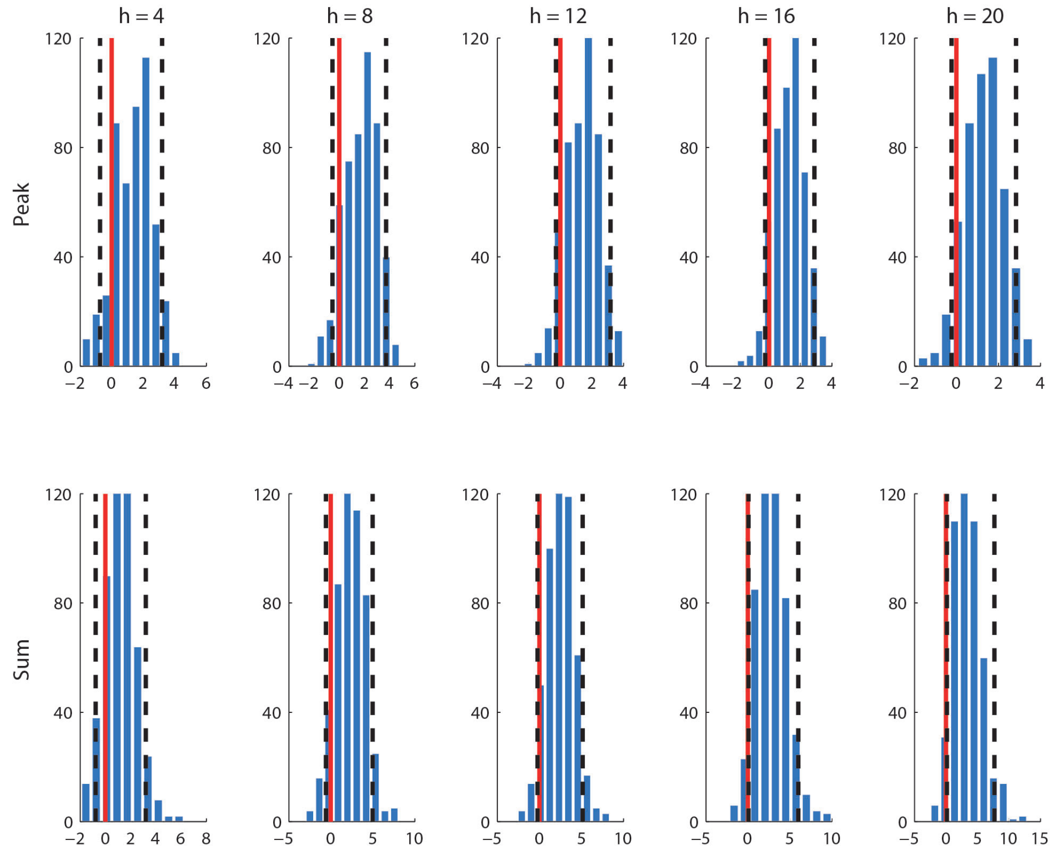
TABLE 2A.7: Fiscal spending multipliers: Extreme events. Different Scenarios

	Peak			
Scenario/State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
Baseline	0.79 [0.45,1.09]	2.27 [1.45,2.93]	1.09 [0.72,2.31]	2.72 [1.32,3.96]
$\widetilde{\eta}_{13}^{g}$ *last*	0.43 [0.19,0.61]	2.55 [1.66,3.34]	0.97 [0.45,3.01]	2.88 [1.44,3.72]
$\widetilde{\eta}_{13}^{g}$ *first*	1.14 [0.24,1.82]	2.74 [1.65,4.48]	1.91 [0.85,3.72]	3.23 [1.51,5.14]
Long sample (Ramey's news)	0.49 [0.20,0.81]	2.61 [1.55,4.62]	0.77 [0.28,1.50]	2.51 [1.21,5.31]
	High debt	*Mod.$^{+}$ debt*	*Mod.$^{-}$ debt*	*Low debt*
Debt/GDP ratio	1.35 [1.15,1.54]	1.22 [0.58,1.81]	1.56 [1.31,2.00]	1.66 [1.24,2.55]

	Sum			
Scenario/State	*Strong exp.*	*Deep rec.*	*Weak exp.*	*Mild rec.*
Baseline	−2.26 [−5.63,−0.78]	1.60 [0.18,2.63]	−1.40 [−3.91,0.65]	1.38 [−0.48,3.02]
$\widetilde{\eta}_{13}^{g}$ *last*	−1.57 [−2.92,−0.91]	2.28 [1.23,3.10]	−0.44 [−1.97,2.29]	2.16 [0.22,3.00]
$\widetilde{\eta}_{13}^{g}$ *first*	−0.70 [−2.50,0.43]	2.36 [0.99,4.29]	0.66 [−1.04,2.90]	2.50 [0.59,4.39]
Long sample (Ramey's news)	0.15 [−0.24,0.53]	1.74 [0.08,3.92]	0.07 [−1.23,0.96]	1.52 [0.60,4.62]
	High debt	*Mod.$^{+}$ debt*	*Mod.$^{-}$ debt*	*Low debt*
Debt/GDP ratio	0.68 [0.15,1.37]	0.74 [−1.02,1.15]	1.33 [0.95,1.66]	1.33 [0.81,1.97]

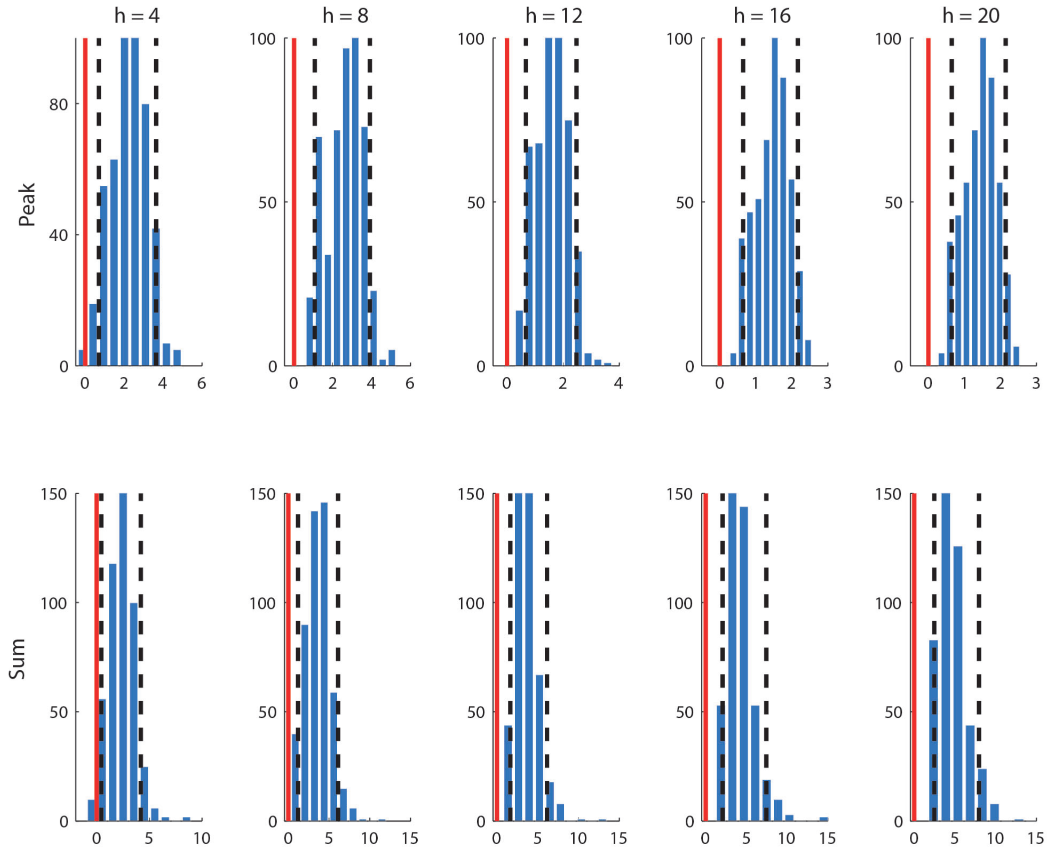
Notes: Four-year integral multipliers. Figures conditional on our VAR analysis with GIRFs conditional on four different sets of initial conditions. Log-values of the impulse response of output rescaled by the sample mean of output over public spending (both taken in levels) to convert percent changes in dollars.

FIGURE 2A.1: Difference in multipliers between recessions and expansions: All histories



Notes: Empirical densities of the differences computed as multipliers in recessions minus multipliers in expansions. Densities constructed by considering all recessions and expansions (initial conditions) present in the sample. Multipliers conditional on the same set of draws of the stochastic elements of our STVAR model as well as the same realizations of the coefficients of the vector. Densities based on 500 realizations of such differences per each horizon of interest. 'h' identifies the number of quarters after the shock.

FIGURE 2A.2: Difference in multipliers between recessions and expansions: Extreme events



Notes: Empirical densities of the differences computed as multipliers in recessions minus multipliers in expansions. Densities constructed by considering just extreme realizations of recessions and expansions (initial conditions) present in the sample. Multipliers conditional on the same set of draws of the stochastic elements of our STVAR model as well as the same realizations of the coefficients of the vector. Densities based on 500 realizations of such differences per each horizon of interest. 'h' identifies the number of quarters after the shock.

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Chapter 3

Opening the Red Budget Box: Real Effects of a Tax Shock in the UK

Abstract

This paper studies the real effects of an exogenous UK tax change in recessions and expansions. The tax shock is identified via the measure recently proposed by Cloyne (2013). Combining local projection techniques (Jordá, 2005) with smooth transition regressions (Granger and Teräsvirta, 1994), tax policy shock is found to affect UK macroeconomic variables depending on the phase of business cycle the economy is when tax shock occurs. A positive tax shock in recessions triggers a large, persistent, negative, and statistically significant reaction in output, consumption, investment, imports and government consumption. The results suggest that output tax multipliers are negative and above one (in absolute value) in recessions but not in expansions. The size and the sign of responses of a number of macroeconomic variables are also found to be state-contingent.

3.1 Introduction

Each year in Spring, the UK Chancellor of Exchequer puts the Budget Statement in a red bag, the red budget box, and carries it from 11 Downing Street to the House of Commons to read the Financial Statement and Government's proposals for change in taxation. Suppose the Chancellor proposes an unexpected increase in taxes, then the following questions arise: *What are the effects of tax changes on output? Are tax shock effects different in recessions and in expansions? How large are tax multipliers in the UK?*

This paper studies the heterogeneous impact of tax shocks in the UK in recessions and expansions. We show that unexpected tax changes exert different effects on a number of macroeconomic variables depending on the phase of business cycle the economy is in when the tax change occurs. Disentangling the (non-linear) effects of taxes on the GDP components (consumption, investment, imports, exports, and government consumption), we highlight which variable drives the GDP reaction in recessions and expansions. We quantify tax multipliers on output and its components in "good" and "bad" times.

We find that the reaction of several variables is asymmetric along the business cycle. In particular, the position in the business cycle when the shock occurs statistically affects the sign and the size of tax multipliers on real variables. In recessions, tax multipliers are negative and above one (in absolute value) on consumption, investments, exports, imports, and output but below one on government consumption. In expansions, tax multipliers are estimated to be negative and above one (in absolute value) on investments but positive and above one on exports, imports, and government consumption. Multipliers are found statistically different across regimes. We show that the effects of tax shocks are quantitatively larger in recessions and smaller in expansions than those predicted by a linear framework.

These results support the empirical evidences that tax policy changes may generate different outcomes according to the macroeconomic conditions. Tagkalakis (2008), analysing a panel of nineteen OECD countries, finds that the effects of tax shocks on private consumption are different in recessions and expansions. This asymmetry can be explained by liquidity constraints of households that can be more severe in recessions than in expansions. Blanchard and Leigh (2013) emphasise that during the "Great Recession" the size of fiscal multipliers have been underestimated. This suggests that fiscal multipliers may vary over time.

Despite the importance to evaluate whether the effects of a tax shock is asymmetric across the business cycle (i.e., in recessions and in expansions), the literature focusing on the non-linear effects of tax change is scant. Indeed, it focuses on the linear effects of a tax shock on output on a single country (i.e., Pereira and Wemans, 2013; Hayo and Uhl, 2014; Cloyne, 2013), and particularly on the US economy (i.e., Blanchard and Perotti, 2002; Mountford and Uhlig, 2009; Romer and Romer, 2010; Favero and Giavazzi, 2012; Perotti, 2012; Mertens and Ravn, 2014), whereas a few studies have focused on a cross-country panel datasets (see e.g., Guajardo, Leigh, and Pescatori, 2011) or on multi-country analysis (i.e., Bénassy-Quéré and Cimadomo, 2012). As for the literature dealing with non-linear effects of UK tax shocks, one exception is Baum, Poplawski-Ribeiro, and Weber (2012) which estimate the effects of a tax shock on output relying on a Threshold VAR. They find that the output tax multiplier is very small and not statistically significant. It means that whether the Chancellor of the Exchequer proposes a cut or an increase in taxes, his proposal is unlikely to stimulate or dampen the economic activity, whatever are the economic conditions.[1]

Two issues make our aim challenging. Firstly, the identification of tax changes because of the endogeneity problem between tax revenues and GDP. For instance, tax revenues shocks might trigger output fluctuations, while shocks affecting output might cause revenue fluctuations. To overcome the endogeneity problem, two main approaches have been proposed in the empirical literature. The first one, pioneered by Blanchard and Perotti (2002), relies on structural vector autoregressive (SVAR) analysis in which cyclically adjusted tax revenues are used to proxy

[1] Afonso, Baxa, and Slavik (2011) study the nonlinear effects of a fiscal policy in Germany, Italy, the UK, and the US. However, using the debt ratio as a proxy for fiscal policy shock, they do not distinguish between revenues and government spending shock.

tax shocks, and it is based on some assumptions about the implementation lags in fiscal policymaking and on the calibration of the fiscal elasticity.[2] The second one, the narrative approach proposed by Romer and Romer (2010), identifies an unexpected tax change analysing written official records and distinguishing tax shocks due to reasons not related to countercyclical concerns (exogenous) from those related to them (endogenous).[3] Several concerns arise from the identification of tax shocks á la Blanchard and Perotti (2002), because it may fail to capture tax shifts that are exogenous. For instance, Romer and Romer (2010) argue that other non-policy movement (i.e., asset and commodity price fluctuation) may affect the cyclically-adjusted revenues and a SVAR may not to address the correlations between these factors. Auerbach and Gorodnichenko (2012) claim that the effects of tax shocks on output may be very sensitive to the calibrated elasticity. Furthermore, Caldara and Kamps (2012) find that the calibrated elasticity may bias downward the effects of such shock on output. Secondly, Ramey and Zubairy (2014) raise another issue that may be behind biased results for the fiscal multipliers in SVAR analysis. The estimated size of fiscal multipliers may be very sensitive to the value of *ex post* conversion factor, i.e. the ratio of the GDP/fiscal variables, used to convert elasticity into multiplier when the model is estimated including logarithm transformed variables.[4]

Our analysis jointly tackles these two issues. We estimate a (linear) Structural VARs identifying the structural tax shock á la Blanchard and Perotti (2002). We

[2]In the Blanchard and Perotti (2002) approach a change in tax revenues depend on the automatic response of taxes to output and on exogenous tax changes. To purge the tax revenues from automatic stabilizers, they calibrate the elasticity of taxes to output via the OECD method and assumptions proposed by Giorno, Richardson, Roseveare, and van den Noord (1995) and van den Noord (2002). The elasticity of taxes to output is calibrated combining the estimation of elasticity of tax revenues to their tax base with the elasticity of tax base to output. The tax revenues purged by its automatic response to output are the cyclically-adjusted measure of tax revenues. Then, the calibrated elasticity is used to pin down the relations linking the reduced form residual to the structural shock in a SVAR framework. The identification of structural shocks is recovered relying on some assumptions about the implementation lags.

[3]This method has been advocated to identify government spending shocks (see e.g., Ramey and Shapiro, 1998; Ramey, 2011), fiscal consolidations (see e.g., Devries, Guajardo, Leigh, and Pescatori, 2011; Guajardo, Leigh, and Pescatori, 2011), tax shocks in the US (Romer and Romer, 2010), in Portugal (Pereira and Wemans, 2013), in Germany (Hayo and Uhl, 2014), and in the UK (Cloyne, 2013).

[4]The transformation of variables in logarithm form is a common practice in the VAR literature, but not only. Indeed, Auerbach and Gorodnichenko (2013a) relying on Local Projection regressions use log-transformed variables.

set the elasticity of taxes to output borrowing two coefficient restriction's values proposed by Perotti (2005) and Cloyne (2013), 0.76 and 1.61, respectively. Then, we convert elasticities into multipliers using different *ex post* conversion factors. The results suggest that tax multipliers increase in the value of coefficient restrictions (i.e., lower when the coefficient restriction is set to 0.76 and higher when it is equal to 1.61). This result for the UK is in line with the one found by Caldara and Kamps (2008) for the US. They highlight that the effects of tax shock will be biased downward whether the calibrated elasticity is too small. Moreover, estimating two different sample sizes (1963:I-2001:II and 1955:I-2009:IV), we find that increasing the value of the coefficient restriction affects the persistence of tax shocks. Furthermore, the combination of identifying tax shock via coefficient restrictions with *ex post* conversion factors may lead to another bias on tax multiplier estimates (see Appendix 3A for details).

To overcome the tax shock identification problem discussed above, we identify the UK tax shocks using the measure constructed by narrative-approach and proposed by Cloyne (2013), whereas to avoid the *ex post* conversion factor one we define the variables as in Hall (2009) and in Barro and Redlick (2011). To estimate the effects of tax shocks conditionally on the state of economy and to avoid dealing with some implicit assumptions of the regime-switching model, we combine the Local Projection (Jordá, 2005) estimations with smooth transition regressions (Granger and Teräsvirta, 1994).[5]

Our main results show that the impact of tax shocks on the macroeconomic variables is asymmetric over the business cycle. Researchers disagree over the (linear) effects of a tax shock in the UK. For instance, Perotti (2005), relying on the Blanchard and Perotti (2002) approach, finds that a positive tax shock has expansionary effects on output, opposite to the conventional wisdom. Cloyne (2013), identifying the tax shock à la Romer and Romer (2010), finds opposite results: an unexpected increase in taxes has negative and statistically significant effects on

[5] The use of single-equation technique in a non-linear framework has been also advocated by Auerbach and Gorodnichenko (2013a; 2013b; 2014), Owyang, Ramey, and Zubairy (2013), Ramey and Zubairy (2014), Ben Zeev and Pappa (2014), Leduc and Wilson (2003) and others as an simple alternative to the VARs.

output (-0.5 and -2.15 over three years). We reconcile these differences considering the phase of the economy in which tax shock occurs. The difference of the results across regimes (recessions and expansion) lies on the relative position of the AD-AS curves. To rationalise these results we consider an AS curve which is relatively flat before the point of full employment level of national income, and then it becomes almost vertical afterwards. In expansions, the aggregate demand curve is in the steeper part of the aggregate supply curve and the effects of tax shocks on output are small. Conversely, in recessions the aggregate demand curve is in the flatter part of the aggregate supply curve, and therefore the variation of output to taxes is greater in recessions than in expansions. We show that the effects of such shocks are quantitatively different than those predicted by a linear framework. A linear estimation overshadows the effects of tax shocks across regimes because it works as an average of the two different effects. Our results are important for a policy standpoint, calling for a tailored use of fiscal policy instruments across the business cycle.

A battery of robustness checks, dealing with alternative specification of tax shocks and different specifications of baseline regressions, confirms the asymmetric effects of a tax shock on GDP and its components.

Our paper is close to Cloyne (2013) and Baum, Poplawski-Ribeiro, and Weber (2012). There are differences between theirs contributions and ours. First, Cloyne (2013) studies the effects of tax shocks on key macroeconomic variables via a linear VAR, and identifying such shocks through the narrative approach. Conversely, we investigate the impact of tax shocks conditionally on the phase of the business cycle the economy is when tax shocks occurs. Second, Baum, Poplawski-Ribeiro, and Weber (2012) study the effects of tax shocks in a non-linear VAR and identifying the structural shock á la Blanchard and Perotti (2002). Differently, we identify the tax shocks via the narrative measure proposed by Cloyne (2013), and to estimate the effects of tax shocks on output, but also on its components, we rely on a non-linear version of the Local Projection (Jordá, 2005) technique.

The remainder of this paper is organized as follows. Section 2 presents the linear Local Projection exercise and relative results. Section 3 extends the previous section focusing on a nonlinear framework. Section 4 provides some robustness checks, whereas section 5 concludes.

3.2 Data definition and Local Projection specification

We estimate the effects of a tax shock on UK macroeconomic aggregates relying on the Local Projection (LP) technique introduced by Jordá (2005). LP allows to project the value of the dependent variable shifted h periods ahead on the information set available at time t. Thus, those projections are local to each horizon.

Consider a h set of regressions for $h = 0,1,2\dots H$ for each variable of interest, $\tilde{X}_{t+h}$, such as output, consumption, investment, imports, exports and government consumption:

$$\tilde{X}_{t+h} = \alpha_h + \zeta_h + B_{Lh}(L)y_{t-i} + \theta_{Lh}\epsilon_t^{Cloyne} + u_{t+h} \tag{3.1}$$

where α and ζ are the constant and the linear trend, B_{Lh} is the coefficient matrix at each horizon h and y_{t-i} is the vector of control variables which include i lags of variables that usually enter in a "fiscal" VAR, such as the log real per-capita terms of the government spending, GDP and tax revenues. To avoid degree of freedom constraints due to lag length and dimension of covariate vector on the maximum horizon h (Jordá, 2005), we opt for a parsimonious specification of y_{t-i} which includes four lags for each variables, as in (Auerbach and Gorodnichenko, 2013a).[6] The tax shock variable (ϵ_t^{Cloyne}) in equation (3.1) is the new tax change measure proposed by Cloyne (2013). It is constructed via the narrative approach proposed

[6]This lag specification is quite standard in the VARs estimated on quarterly data.

by Romer and Romer (2010) and allows to separate exogenous components of tax changes from the endogenous ones (i.e., tax policy change not due to countercyclical concern versus these due as response to the macroeconomic fluctuations). In particular, Cloyne's tax shock measure includes four categories of exogenous tax changes.[7] Firstly, it includes "long-run" economic reforms not aimed at offsetting macroeconomic fluctuations. The second component is the "ideological" tax changes adopted for political reasons, whereas the third one refers to the "external change" (for example, imposed from a court judgments or European directives). The fourth component is the "deficit consolidation" not driven by current movement in deficit or as consequence of other macroeconomic shock but, for example, to anchor Government's credibility. The series is aggregated according the implementation date to avoid contemporaneous endogeneity of tax revenue to GDP. The changes in revenues are normalized by the GDP and expressed as percentage. Then, a change in Cloyne's measure will reflect the forecast "full year" change in revenues in each quarter. The fact of having an estimate of the unanticipated fiscal shock enables us to employ a uniequational approach to compute dynamic responses of a given macroeconomic variable of interest. In other words, we need not appeal to a VAR framework to identify the effects of an exogenous variations in taxes. The advantage of the uniequational approach is that it is less prone to model misspecification, hence – all else being equal - it reduces the risk of producing biased impulse responses. For further discussions on this approach, see Alesina, Favero, and Giavazzi (2014).[8] The effects of a tax change (ϵ_t^{Cloyne}) on each variables of interest ($\tilde{X}_{t+h}$) are captured by parameter θ_{Lh} in equation (3.1) which, also, depicts the contemporaneous effect of an exogenous tax shock on the

[7]The source for revenue estimates are the Financial Statement and Budget report and the official parliamentary record

[8]Different model specifications have been proposed in the tax literature. For instance, Romer and Romer (2010) regress the dependent variable (GDP) on the contemporaneous value and 12 lags of their tax measure. Cloyne (2013) includes 12 lags of his tax measure, as in Romer and Romer (2010), but in an "augmented" VAR which includes the consumption, investment and GDP equations, as in Mertens and Ravn (2014). Favero and Giavazzi (2012) include only the contemporaneous value of the Romer and Romer (2010) tax shock in a VAR which models, among variables, also the revenues one. Our specification is very close to the one in Favero and Giavazzi (2012). However, we address the issue of different lag length of tax shocks to be included in our specification in the robustness check section.

UK macroeconomic aggregates. Thus, the IRFs are constructed as a sequence of estimated $\{\theta_{Lh}\}_{h=0}^{20}$.

The main advantage of this methodology for the tax multiplier estimations is that it does not require that the left-hand side variables in equation (3.1) should be specified in the same form as the right-hand side variables. This property allows to define each dependent variable of interest $\tilde{X}_{t+h}$ as in Hall (2009), Barro and Redlick (2011) and Owyang, Ramey, and Zubairy (2013). In particular, $\tilde{X}_{t+h}$ is defined as following:

$$\tilde{X}_{t+h} \approx (lnX_{t+h} - lnX_{t-1})\frac{X_{t-1}}{GDP_{t-1}} \tag{3.2}$$

where $(lnX_{t+h} - lnX_{t-1})$ refers to the accumulated change from time *t-1* to *t+h*, whereas the ratio X_{t-1}/GDP_{t-1} converts *ex ante* the percent change to pound change at each point on time, as in Owyang, Ramey, and Zubairy (2013). Thus, this specification overcomes the problem of *ex post* conversion factors, and avoid bias in the estimation of tax multipliers. According to Hall (2009), if the dependent variable is divided by the same denominator as the independent one, the definition of multipliers is preserved. Notice that the Cloyne tax shock measure is normalized by the GDP. The dependent variables are transformed according to equation (3.2). Thus, the dependent variables $(\tilde{X}_{t+h})$ and the tax measure (ϵ_t^{Cloyne}) are divided by the same denominator. Hence, the coefficient θ_{Lh} of equation (3.1) captures the contemporaneous tax multiplier of each variable of interest $\tilde{X}_{t+h}$ at each horizon h, useful to evaluate strictly temporary tax changes.

Jordà's method implies the serial correlation in the error terms. To account for it, we computed confidence intervals relying on the (circular) block bootstrap (Politis and Romano, 1992). We estimate equation (3.2) using quarterly data from 1955Q1-2009Q4. The beginning of the period is motivated by the availability of the quarterly data, whereas the end by the exogenous tax change measure. Table 3.1 summarizes the variables used and their sources.

3.2.1 (Linear) Results

We look at the effects of tax shock on GDP, before turning to the analysis of tax multipliers on GDP components. Figure 3.1 displays the impulse response of output following a 1% percent of GDP increase in taxes. The blue line denotes the sequence of the estimated θ_{Lh} coefficients,[9] whereas the dark and light shaded bands refer to the 68% and 90% confidence intervals, respectively. However, throughout the paper, we define as statistically significant those estimates for which 90% confidence intervals do not include the zero line.

An exogenous tax increase has a negative, statistically significant, and persistent effect on output. The GDP decreases on impact by -0.5% and troughs at -1.8%, 17 quarters after the initial shock, then slowly goes back to its steady state.

Which components of GDP drive the response of output? To answer this question, we transform each component of the GDP, such as private consumption, investment, exports, imports and government consumption according to equation (3.2), and we run h equations for each dependent variables of interest. Figure 3.2 depicts the reaction of each GDP component to a tax shock. The impulse responses can be interpreted as deviations from the baseline and expressed as percent change of GDP. The private consumption decreases on impact by -0.62% hitting a trough of -2%, one quarter before the GDP. This reaction is statistically significant, negative and persistent for all the h horizons considered. Afterwards, consumption gradually returns to its steady-state. The investment (gross fixed capital formation) reaction is statistically significant for the first 8 quarters after the shock occurs. It decreases on impact by a small amount (-0.1%) and hits the trough (-0.5%) in the 4th quarter. Tax shock does not affect exports in the short-run, whereas it has a negative and persistent effects on imports. The reaction of imports mimics the shape of the consumption and output one, albeit quantitatively smaller than those. Indeed, the imports decrease on impact by -0.14% and reach a trough of -0.64% (the 8th quarter), then gradually and go back to

[9] The transformation of the dependent variables as in Hall (2009) and Barro and Redlick (2011), and the construction of IRFs are a sequence of $\{\theta_{Lh}\}_{h=0}^{20}$ imply that the contemporaneous tax multipliers on each variable of interest and at each horizon h can be read directly from IRFs.

the steady-state. Also, our estimations predict a positive reaction of government consumption, albeit small.

How large are the cumulative tax multipliers in the UK? The cumulative multipliers are useful not only to capture short-run effects but also permanent tax shock effects. The cumulative tax multipliers from one to five years are computed as the ratio of the cumulative change in the variable $\tilde{X}_{t+h}$ to the the cumulative change in tax revenues, in response of an unanticipated tax shocks identified by ϵ_t^{Cloyne}. [10] In order to obtain the estimated θ_{Lh} coefficients of the variable in denominator (tax revenues), we transform the revenues series as in (3.2), plug it in equation 3.2, and we run h equations. After having obtained the estimated θ_{Lh} coefficients for the tax revenues equation, we divide the sum of bootstrapped θ_{Lh} coefficients until horizon H of each variable of interest to the counterpart bootstrapped θ_{Lh} coefficients of the tax revenues' equations. Table 3.2 shows the median cumulative tax multipliers. Bold numbers refer to tax multipliers statistically significant at the 90% level. Results show that the cumulative output tax multipliers range between -0.46 (1-year tax multiplier) and -2 (5-year tax multiplier).

Turning to the tax multipliers on the GDP components, the tax shock has a bigger effect on the private consumption with multipliers estimated between -1.56 and -2.24, whereas these on investment are lower and more stable than the ones on consumption (from -0.5 to -1). The tax multipliers on imports are negative and statistically significant overall the periods, ranging from -1 (1 year cumulative multiplier) to -0.75 (5 year cumulative multiplier). Turning to the government consumption multipliers, a tax shock increases the government consumption around 0.5, albeit very small.

Comparing our result to the literature, Perotti (2005) finds that a tax cut has a contractionary effect on output over the 20 quarters. Focusing on the early part of his sample he finds that the 1-year and 3-year cumulative output tax multipliers are equal to 0.2, whereas those become negative (-0.4 and -0.7, respectively)

[10]The cumulative multipliers are computed as $\frac{\sum_{h=1}^{H}\Delta\tilde{X}_{t+h}}{\sum_{h=1}^{H}\Delta \tilde{Revenues}_{t+h}}$ for H=*4,8,12,16,20*, where $\Delta\tilde{X}_{t+h} \equiv \tilde{X}_{t+h} - \tilde{X}_{t-1}$.

when he focuses on the late part of the sample.[11] Bénassy-Quéré and Cimadomo (2012), relying on the coefficient restriction method á la Blanchard and Perotti (2002), estimate that a tax cut generates on impact a positive GDP multiplier (0.12) and negative (-0.12) at the 8th quarter, but not statistically significant. Conversely, the HR Treasury (2003) reports that a tax cut generate an 1-year cumulative output tax multiplier of 0.4, positive and statistically significant albeit very small. According to Cloyne (2013), a tax increase has a recessionary effect on output that decreases on impact by -0.6 up to -2.15, 10-12 quarters after the initial shock. Our response of GDP to a tax shock is quantitatively in line with the Cloyne's one (2013).

Overall, our results are in line with the Keynesian theory. A tighter fiscal policy negatively affects aggregate demand as disposable income and private consumption decrease. Households decrease their demand for both domestic and foreign goods, firms decrease investment, whereas government consumption increases.

3.3 The Non linear Model

Do tax multipliers change according to the state of the economy? Following Auerbach and Gorodnichenko (2013a; 2013b; 2014) we study the non-linear effect of a fiscal shock on variables of interest combining the LPs (Jordá, 2005) with smooth transition regressions (Granger and Teräsvirta, 1994). The response of dependent variables $\tilde{X}_{t+1}$ to a tax shock is estimated by the following regression:

$$\tilde{X}_{t+h} = F(z_{t\_i})(B_{R,h}(L)y_{t-i} + \theta_{R,h}\epsilon_t^{Cloyne}) + (1 - F(z_{t\_i}))(B_{E,h}(L)y_{t-i} + \theta_{E,h}\epsilon_t^{Cloyne}) + u_{t+h} \tag{3.3}$$

where R stands for Recession and E for Expansion. We estimate the equation (3.3) including a constant and a linear trend. Each variable of interest $\tilde{X}_{t+1}$ is

[11] As Perotti (2005) noted, he finds such contractionary effects on GDP for these countries (UK, Australia, Germany) for which he estimated the smallest output elasticity of net taxes. This highlights that the latter may have been underestimated.

projected on the same vector of covariates y_{t-i} of the linear specification, and $B_{R,h}$ and $B_{E,h}$ refer to coefficient matrices of the recessionary and expansionary phase, respectively. The lagged variables in y_{t-i} are used to control for the history of the shock, as in Auerbach and Gorodnichenko (2013a). The effect of a tax shock on $\tilde{X}_{t+h}$ at horizon h is captured in recessions by $\theta_{R,h}$, whereas in expansions by $\theta_{E,h}$.

As in Auerbach and Gorodnichenko (2013a; 2013b; 2014), the transition of $\tilde{X}_{t+1}$ from one regime to another is governed by a logistic function that depends on z_t:

$$F(z_t) = \frac{exp(-\gamma z_t)}{1 + exp(-\gamma z_t)}, \qquad \gamma > 0, \qquad z_t \sim N(0,1) \tag{3.4}$$

The transition function in (3.4) is a monotonically increasing function of z_t, where F is a continuous transition function bounded between zero and one and z_t is the transition variable. The slope parameter γ determines the smoothness of the change between zero (strong expansions) to one (strong recessions), and the identification restriction is that $\gamma > 0$. If $\gamma \to \infty$ in (3.4), then equation (3.3) becomes a two-regime switching regression model.

Before estimating equation (3.3) for each variable of interest $\tilde{X}_{t+h}$, we formally test for the presence of nonlinearities. Linearity is tested replacing the transition variable $F(z_{t-i})$ by the third order Taylor series approximation around $\gamma = 0$, as suggested by Lukkonen, Saikkonen, and Teräsvirta (1988).

We test linearity as following:

$$X_t = w_t\beta_0^{'} + (\tilde{w}_t z_{t-i})'\beta_1 + (\tilde{w}_t z_{t-i}^2)'\beta_2 + (\tilde{w}_t z_{t-i}^3)'\beta_3 + u_t \tag{3.5}$$

where vector w_t contains four lags of covariates (log-real GDP, government spending, revenues) and the contemporaneous value of tax shock (ϵ_t^{Cloyne}). Testing the null hypothesis of linearity versus nonlinearity is equivalent to perform an LM (χ^2) test of $H_0^{'} : \beta_\iota = 0$, ι= 1, 2, 3, against $H_1^{'}$: at least one $\beta_\iota \neq 0$. We perform the linearity test plugging in X_t each variable of interest and in z_{t-i} each potential transition variable, such as the lagged *(t-i)* standardized backward-looking moving average (MA) over *(j)* quarter(s) of the output growth rate with $i \in I = 1, .., 5$ and $j \in J = 2, ...8$. The choice of i is justified to avoid that tax shocks may have some contemporaneous feedback on the state of economy. Notice that all the transition variable candidates have been standardized in order to be comparable. Table 3.3 reports the p-value (multiplied by 100) of linearity tests. The tests suggest a clear rejection of the null hypothesis of linearity. We choose for each variable of interest the transition variables $MA(j)$ lagged at time *t-i* corresponding to the smallest p-value (Teräsvirta, 1988). That because whether there is a correct transition variable among the different alternatives, the power of the test is maximized against it. Table 3.3 highlights that the nonlinearity of the GDP, consumption, investment, exports is governed by a $MA(2)$ lagged at *t-1*, whereas that one of import and government consumption by a $MA(2)$ lagged at time *t-5*.

Then, we calibrate the smoothing parameters γ in order to match the probability of being in recession obtained applying the BBQ algorithm on the logarithm of the real GDP (more details in the Appendix 3B). We define a recessionary regime a period for which $F(z_t) \geq 0.85 \approx 0.15$. It means that the economy spends about 15% of the time in the recessionary state and 85% of the time in the expansionary one.[12] This implies setting $\gamma = 1.7$. Figure 3.3 plots Cloyne's tax change measure versus the recessionary (shaded area) and expansionary phases,[13] whereas figures 3.4 and 3.5 refer to the transition variable z_t and transition function $F(z_t)$, respectively.[14]

Notice that one important advantage of the LPs is that the impulse responses incorporate the average transitions of the economy from one regime to another.

[12]The values of γ are in line with estimates obtained regressing in a logit model the dummy variables (R=1 and E=0) obtained by the BBQ algorithm on transition variables (results available upon request).

[13]The correlation between the tax shock measure and $F(z_t)$ or $(1 - F(z_t))$ is equal to zero.

[14]In the transition function $F(z_t)$ (figure 3.5), the frequency of non-alterning points is high. It depends on the characteristic of the transition variable z_t. This is in line with Harding (2008) which show that in the UK "the frequency of non-alterning points is four times higher than the US". Of course, if z_t has some non-alterning points, this characteristic will be amplified in the transition function $F(z_t)$ in which those points are bounded between zero and one.

According to Ramey and Zubairy (2014), the estimated coefficients in equation (3.3) depend on the characteristic of the economy from time t to $t+h$, given the initial conditions (the tax shock, the initial state of the economy, and the control variables). Since the control variables in equation (3.3) do not change at each horizon h, then the estimated coefficients on the covariates capture the average transition of the economy from one state to another occurring in the sample. Also, the estimated coefficients ($\theta_{R,h}$ and $\theta_{E,h}$) on the ϵ_t^{Cloyne} will reflect the effects of the tax shock on the future state of economy. For example, suppose that a tax shock has a negative effect on output in recessions and positive in expansions, and a tax shock occurs in an expansionary period bringing the economy in a recessionary one. Then, the estimated parameters $\theta_{E,h}$ will incorporate the transition of the economy from the expansionary to the recessionary regime changing its values from positive to negative.[15]

3.3.1 (Non Linear) Results

Figure 3.6 depicts the IRFs of macroeconomic indicators to an increase in the exogenous tax shock series of 1% of GDP, according to the two different regimes. Notice that in expansions IRFs are constructed as a sequence of estimated $\{\theta_{Eh}\}_{h=0}^{20}$, whereas in recessions as $\{\theta_{Rh}\}_{h=0}^{20}$ one.[16] The blue dotted lines denote the IRFs in expansions, whereas the pink lines the ones in recessions.[17] The dark and light shaded area represent the 68% and 90% confidence intervals.[18]

[15]Using a SVAR we can account for this feedback only through Generalised IRFs, as in Caggiano, Castelnuovo, Colombo, and Nodari (2015). As noted by Owyang, Ramey, and Zubairy (2013), the difference between GIRFs and LPIRFs is based on how the two IRFs account for this feedback: in the GIRFs using the response at time *t-1* to estimate the response at time t, whereas in LPs computing the average h-period-ahead value forecast given the information set at time t. See Owyang, Ramey, and Zubairy (2013) for a careful discussion and comparison between the GIRFs and the LPIRFs

[16]Notice that since dependent variables are transformed as in Hall (2009)) and Barro and Redlick (2011) each point estimated of the IRF at time h correspond to the tax multiplier at time h

[17]See Auerbach and Gorodnichenko (2013a), note 6, for analytic comparison between the LPIRFs and the conventional IRFs.

[18]A confidence level of 68% is reported only for comparison reason since this level is quite common in the relative literature.

Focusing on the recessionary regime, the impact response of private consumption is negative (-1%), hitting its lowest value (-4.5%) 8 quarters after the shock occurs, then goes back to its steady-state. The investment decreases and reaches its trough at -1%. Tax shock is found to affect statistically and negatively exports and imports. On impact, imports decrease by -0.5% and, after it troughs at around -3%, then gradually goes back to zero. The positive response of government consumption in the first four quarters is not statistically significant. However, when it troughs at -1.2%, the response becomes statistically significant. Overall, the above effects translate in the output's response. Indeed, following a tax shock the output decreases on impact by -1.2% and hits its lowest value at -4.9% eight quarters after the shock, then gradually returns to the steady-state.

Next, we look at the response of macroeconomic aggregates to a tax shock in expansions. On impact, a tax shock affects negatively consumption (-0.33%) but positively investments (0.12%). However, both reactions are not statistically significant at 90% confidence intervals. Also, tax shock is found to not affect statistically the impact reaction of imports and exports. Conversely, government consumption reacts statistically significant and positively to a tax shock increasing on impact by 0.2% and, after reaching its peak at 0.97, then gradually goes back to zero. The reaction of output to tax shock it is not statistically significant for the first sixteen quarters.

Are tax multipliers state-dependent? Table 3.4 reports cumulative tax multipliers from one to five years for all the variables of interest computed, as in the linear specification, à la Owyang, Ramey, and Zubairy (2013).[19] From columns (1) to (3) the table reports tax multipliers for the recessionary, expansionary and linear case, respectively. Bold numbers indicate the multipliers statistically significant at the 90% level, whereas in brackets the 90% confidence intervals is reported.

[19] As in the linear specification, to obtain the estimated θ_{Rh} and θ_{Eh} coefficients of the variable in denominator, we transform revenues as in (3.2), plug it in equation 3.3, and we run h equations. For instance, to compute the cumulative multipliers in recessions, $\frac{\sum_{h=1}^{H}\Delta\tilde{X}_{R,t+h}}{\sum_{h=1}^{H}\Delta Rev\tilde{e}nues_{R,t+h}}$, we divide the sum of bootstrapped θ_{Rh} of each variable of interest to the counterpart bootstrapped θ_{Rh} coefficients of revenues.

In recessions, tax multipliers are statistically significant and negative for all the variables considered. The output and private consumption cumulative multipliers are always larger than one (in absolute value), even on impact. The multipliers of imports quantitatively mimic the multipliers for consumption. Also for exports, tax multipliers are larger (in absolute value) than one. The multipliers of investment are larger than one (in absolute value) only for the two-year cumulative multiplier, while for government consumption the multipliers are negative but smaller than one.

Turning on the expansionary cumulative tax multipliers, for output and private consumption the multipliers are not statistically significant for the first four-years, whereas for investment tax multipliers are negative and larger (in absolute value) than in recessions for the four and five-year cumulative multipliers. For exports, imports, and government consumption tax multipliers are positive and larger than one.

Figure 3.6 and table 3.4 show that the effect of tax shocks varies across the business cycle. In recessions, the increase in taxes reduces consumptions. This results accords with the analysis of Tagkalakis (2008): since the fraction of liquidity constrained households is likely to increase, the increase in taxes decreases their disposable income. A tighter fiscal policy decreasing the disposable income of households has negative wealth effects, and therefore households consume less. Conversely, in expansions the negative effect of an increase in taxes is counteracted by an increase in government consumption. It turns out that the reaction of private consumption is not statistically significant. With regard to output, the results have some similarities to consumption. In recessions, the shock decreases output, but this effect disappears in expansions. Since in our sample consumption represents 57% of GDP, the reaction of output is likely to be driven by the reaction of consumption. This tendency also holds in expansions. Turning to investment, our results show that an increase in taxes has negative effects, both in recessions and in expansions. The increase in taxes reduces the business profits and the investment financed by those profits. The response of imports and exports is asymmetric across the business cycle. In recessions, imports strongly decrease likely because of the fall in income, and therefore in domestic demand for foreign goods. The reaction of exports may depends on the exchange rate[20] and on external factors. Interestingly, imports and exports increase in expansions.

Overall, our results predict asymmetric effects of tax shocks across the business cycle. The results can be read through the lenses of the AD-AS model. Suppose that the economy is producing at its full employment level of natural income and the aggregate supply curve is flat, and becomes steeper and steeper and vertical at this point. If the economy starts from its equilibrium level and there is an expansionary phase, the aggregate demand shifts rightwards and the impact on output will be small, since the aggregate supply is almost vertical. The only effect is on prices. If the economy is in expansions and the government increases taxes, the new aggregate demand curve will shift towards to the original curve leading to a small loss in output. If the economy is in recession, the original aggregate demand shifts on the left. In this case, the movement of the aggregate demand is happening in a point of the aggregate demand that lies flatter than the one considered before. A backwards shift of the AD curve will affect output negatively stronger than before. Suppose the economy is in a recessionary phase and government increase taxes. This causes a further movement toward left yet in even flatter part of the aggregate supply curve. The negative effects on output of this policy intervention will be in absolute value stronger than the same policy intervention that happens during expansions. The difference between the two phases of the economy (recessions and expansion) lies on the relative position of the AD-AS curves. In expansions, the AD curve is in the steeper part of the AS curve and the effects on output are small. Conversely, in recessions the AD curve is in the flatter part of the AS curve, and therefore the $\left|\frac{\Delta Y}{\Delta T}\right|^{Rec} > \left|\frac{\Delta Y}{\Delta T}\right|^{Exp}$.

[20] Notice that the UK has experimented different exchange rate regimes in the postwar period. Indeed, until 1972 it was part of the Bretton Woods system. From 1972 to 1990 it adopted a semi-managed floating regime. From 1990 to 1992 the UK was part of the European Exchange Rate Mechanism. From 1993 the UK has adopted a floating exchange rate regime. Studying the reaction of the exchange rate to a tax shock is already in the agenda. However, it suffices here to say that focusing on the subsample 1972-2009 -according to Ilzetzki, Mendoza, and Végh (2013) can be considered a flexible exchange rate regime- and adding among the covariates the real exchange rate does not affect the results.

Statistically evidences in favor of state-dependent tax multipliers. According to our estimation the tax multipliers are state-dependent. To best of our knowledge, only Baum, Poplawski-Ribeiro, and Weber (2012) have studied the non-linear effect of tax shock in the UK (via a Threshold VAR), albeit focusing on the reaction of output. They apply the Blanchard and Perotti (2002) identification strategy and find that a tax increase has negative (-0.4) or positive (0.2) effect on the 1-year output tax multiplier depending on whether the economy is in the "positive output" regime or in the "negative" one, respectively. However, the estimated multipliers are not statistically significant.[21]

To statistically support whether the multiplier are different according the regimes, we test the difference between the multipliers estimated under recessionary and expansionary regimes.[22] We run this test for all our variables of interest. Table 3.5 reports the results. If the value of zero is not included in the confidence bands, then there will be evidence of state-dependent tax multipliers. Indeed, we find statistically significat evidences of non-linearity in the tax multipliers for the GDP and private consumption during the second and third year. Turning to investments, table 3.5 shows that at longer horizon the zero value is not included in the confidence bands. We find statistically evidences of non-linearity in the tax multipliers for the exports, imports and government consumption tax multipliers. Overall, whether we jointly read table 3.4 and table 3.5, there are statistically evidences that tax multipliers in UK are state-dependent. Thus, evaluating tax multipliers in a linear framework may lead to underestimate the effects of increasing taxes in recessions and overestimated the ones in expansions. Blanchard and Leigh (2013) highlight that for the recent recession the size of fiscal multipliers

[21]Baum, Poplawski-Ribeiro, and Weber (2012) identify the tax shock in two steps. First of all, they eliminate from the tax revenue series cases of revenues change not related to fiscal policy decisions (i.e., movement in commodity price and asset). To this aim, they compare the IMF (2010) action-based measure with the cyclical adjusted primary balance (CAPB) and whether the divergence between the two measures was large, then revenue changes unrelated to fiscal policy decision are removed from the revenue series. Doing that, the revenues series reflects change in output and fiscal policy decisions. Secondly, to identify a structural tax shock unrelated to movement in output they apply the Blanchard and Perotti (2002) procedure.

[22]The empirical density of the difference between multipliers is based on 10,000 realizations of such differences for each horizon h.

has been underestimated. We highlight that macroeconomic conditions can affect fiscal multiplier estimates.

3.4 Robustness checks

Results highlight that tax multipliers in UK are state-dependent. In this section, we check the robustness of our findings.

Alternative measure of tax shocks. We have identified the tax shock in equation (3.1) and (3.3) via the contemporaneous value of the Cloyne tax shock (ϵ_t^{Cloyne}). This specification is close to the Favero and Giavazzi (2012) one, given that we also include in vector y_t lagged values of revenues.[23] The reasons of this specification are twofold. Firstly, we treat ϵ_t^{Cloyne} as an observable and exogenous shock to revenues.[24] Secondly, this specification allows us to preserve degrees of freedom given our sample size. Notice that we have extended the linear analysis to the non-linear one. Whether the inclusion of lagged values of the ϵ^{Cloyne} may be not problematic in terms of degree of freedom in the linear specification, it will be in the non-linear one since the parameters to be estimated double. However, lagging the tax shock allows to account for the possibility of a partial revision to tax shocks.

Another issue rises since tax shock is constructed on the base of policymakers' intentions. For instance, policymakers may declare that tax changes are made for reason unrelated to movements in macroeconomic variables, while in reality they are concerned about these. We tackle the two above issues by regressing the ϵ_t^{Cloyne} on its own 12 lags and on the covariates that enter in vector y_t. Then,

[23]Romer and Romer (2010) study the effect of a tax shock on the US GDP regressing the GDP on the contemporaneous value and 12 lags of their tax measure. Favero and Giavazzi (2012) add to the Romer and Romer (2010) specification some fiscal variables. They show that the truncated moving average representation of Romer and Romer (2010) shocks give biased estimates of the output reaction because of correlation between the Romer and Romer (2010) shocks and distant lags of output and taxation. Because of that they identify the shock via the contemporaneous value of the Romer and Romer (2010) shock (exogenous term) and treat it as the structural shock of the one of the variable included in the VAR (revenues).

[24]Notice that correlation between the ϵ_t^{Cloyne} and the lagged values of tax revenues, included in vector y_t, is low and range between 0.02 and 0.05.

we select the number of terms of the MA(p) process by checking the statistical significance of such terms. It turns out that p=4 is the last significant term of the process. Thus, we identify the tax shock via the residual obtained regressing ϵ_t^{Cloyne} on its own 4 lags and on the covariates that enter in vector y_t. In this way, we identify a *residual* tax shock purged from the potential revision in tax changes and movements of some macroeconomic variables. The correlation between the ϵ_t^{Cloyne} and the alternative (residual) tax shock measure is 0.94. We plug such residual in equation (3.1) and (3.3) instead of ϵ_t^{Cloyne}, and we estimate them. Figure (3.7) plots the Cloyne tax measure (in our notation ϵ_t^{Cloyne}) versus the alternative measures of tax shock identified by the residual of the above exercise, whereas table 3.6 reports tax multipliers from this exercise. The results are in line with the ones obtained in our baseline model (table 3.4).

The Cloyne's measure used to identify tax shocks includes, among the subcategories, the tax shocks driven by "deficit consolidation" (DC, henceforth) motivations. As pointed out by Cloyne (2013), the "DC" subcategory is different from the Romer and Romer (2010) one. In Romer and Romer (2010), the "DC" category is treaded as exogenous because it reflects past shocks, not related to macroeconomic conditions. For instance, it captures an increase in taxes to reduce an inherited deficit to long-run economic reasons. Conversely, Cloyne (2013) notes that in the UK part of the tax changes due to fiscal consolidation are related to current macroeconomic conditions (endogenous). For that reason, the "DC" in Cloyne (2013) is more restrictive than in Romer and Romer (2010) one and, it includes only 12 observations. Once again, since tax shock is constructed on the base of policymakers' intentions, we verify the robustness of our results excluding from the ϵ^{Cloyne} the DC subcategory. Table 3.8 shows that our baseline results, both in the linear and non-linear specification, are not affected.

Alternative specification. Francis and Ramey (2009) and Owyang, Ramey, and Zubairy (2013) highlight the importance of including a quadratic trend in the US post-WWII period because of the slow-moving demographics. We address this issue for the UK replacing the linear time trend in equation (3.1) and (3.3) with the quadratic one. The results from these exercises are reported in table 3.7.

Furthermore, to control for monetary policy actions we add to the control vector y_t of equation (3.1) and (3.3) the policy rate and the inflation.[25] Table 3.11 shows that adding other variables to our baseline specification, both in the linear and non-linear specification, does not affect are results.

Alternative values of the smoothness parameter. We calibrate the smoothness parameter γ to match the frequencies of the UK recessions obtained via the BBQ algorithm (see Appendix 3B), in our sample 15%. Once again, we (re)calibrate γ in order to include in our sample a number of recessions ranging from 10% to 20%. The lower bound is set by the minimum amount of observations each regime should contain (Hansen, 1999). Table 3.9 and 3.10 show that our results are robust to alternative calibrations of γ parameters.

Figure 3.8 and 3.9 plot the impulse responses of GDP and its components from the above exercises. Our robustness checks confirm the non-linearity of tax shock effects on real variables.

3.5 Conclusion

This paper studies the non-linear effects of an exogenous tax increase in the UK. The tax shock is identified by the new measure proposed by Cloyne (2013). We model non-linearity via the combination of local projection technique (Jordá, 2005) and smooth transition regressions (Granger and Teräsvirta, 1994). We find that the sign and the size of tax multipliers on real variables change according to the states of economy. In recessions, a positive tax shock dampens the economic activity. The (negative) reaction of output is mainly driven by a fall in private consumption. In expansions, output and consumptions do not respond to a tax shock in the short-run. The reason can be found in the asymmetric reaction of government consumption across the business cycle. Since in expansions government consumption reacts positively to tax shocks, it plays an important role in

[25] The inflation rate is the annualized Retail Price Index, since the Consumer Price Index is not available from 1955.

counteracting the (negative) effect of such shock on output and consumption (but not on investment).

Other studies in the literature, as Perotti (2005) and Cloyne (2013), find contrasting results: a positive tax shock has expansionary effect (albeit close to zero) in Perotti (2005), and contractionary one in Cloyne (2013). We reconcile these differences considering the state of the business cycles the economy is when exogenous tax change occurs.

The results documented in this paper lead to new research questions. For example, we have seen that following a tax shock the reaction of imports is different according to the state of economy. It would be interesting to study whether the UK tax shock has some (non-linear) spillover effects on its trade partner countries. Moreover, Romer and Romer (2010) have constructed a narrative-based tax shock measure for the US. It may worth studying the asymmetric effects of a tax shock in the US and UK, to compare the non-linear effects of a tax shock in a relatively close (US) versus a small open economy (UK).

TABLE 3.1: Data Sources

Series	Description	Sources
GDP	Real GDP	ONS
Nominal GDP	GDP in current prices	ONS
Consumption	Final household consumption expenditure	ONS
Investment	Gross fixed capital formation	ONS
Imports	Trade in goods and services: Total imports	ONS
Exports	Trade in goods and services: Total exports	ONS
Population	UK total population	Eurostat
Inflation	Change in Retail Prices Index	ONS
Interest rate	Official Bank rate	Bank of England
Government consumption	Government consumption of goods and services	ONS
Tax revenues	Total tax and NI receipts	ONS
Cloyne's Tax Shocks	Exogenous tax changes (Cloyne, 2013)	AER's website

TABLE 3.2: Linear-Cumulative Tax Multipliers (baseline)

		Cumulative Multipliers Linear (1)			Cumulative Multipliers Linear (1)
	Horizon			Horizon	
GDP	4	**-0.46**	**Priv. Cons.**	4	**-1.56**
		[-3.00, 0.63]			[-3.45, -0.90]
	8	**-1.67**		8	**-1.85**
		[-2.87, -1.04]			[-3.06, -1.25]
	12	**-1.81**		12	**-2.09**
		[-2.79, -1.24]			[-3.12, -1.53]
	16	**-2.03**		16	**-2.37**
		[-2.88, -1.50]			[-3.29, -1.82]
	20	**-2.08**		20	**-2.24**
		[-2.91, -1.55]			[-3.00, -1.75]
Investments	4	**-1.00**	**Exports**	4	-0.19
		[-2.19, -0.57]			[-0.78, 0.25]
	8	**-0.90**		8	-0.24
		[-1.53, -0.59]			[-0.60, 0.01]
	12	**-0.72**		12	-0.17
		[-1.11, -0.49]			[-0.40, 0.02]
	16	**-0.63**		16	**-0.18**
		[-0.92, -0.45]			[-0.37, -0.02]
	20	**-0.53**		20	**-0.26**
		[-0.75, -0.38]			[-0.42, -0.14]
Imports	4	**-1.05**	**Gov. Cons.**	4	**0.45**
		[-2.37, -0.53]			[0.21, 1.05]
	8	**-1.04**		8	**0.47**
		[-1.77, -0.67]			[0.26, 0.84]
	12	**-0.92**		12	**0.42**
		[-1.40, -0.64]			[0.24, 0.70]
	16	**-0.84**		16	**0.47**
		[-1.21, -0.62]			[0.30, 0.70]
	20	**-0.75**		20	**0.50**
		[-1.03, -0.57]			[0.32, 0.71]

Notes: the table shows the cumulative multipliers on output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The multipliers are calculated over one to five years. The 90% confidence intervals are reported in brackets. Bold numbers indicate coefficients statistically significant at 90%.

TABLE 3.3: Linearity Tests

GDP	z_{t-i}	**MA(2)**	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
	t-1	**0.00**	0.03	0.15	0.03	47.25	84.03	82.97
	t-2	0.38	1.03	0.24	0.05	88.23	95.53	80.73
	t-3	7.68	0.18	3.32	1.71	95.17	93.58	98.44
	t-4	0.54	5.64	2.32	74.12	86.87	96.51	99.35
	t-5	52.3	39.04	6.65	17.13	98.79	99.90	96.69
Priv. Cons.	z_{t-i}	**MA(2)**	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
	t-1	**0.02**	1.34	0.83	0.83	0.83	2.40	3.35
	t-2	3.62	7.62	3.22	1.03	32.14	24.87	17.38
	t-3	10.05	0.61	6.65	12.69	33.94	42.82	19.95
	t-4	2.93	5.01	8.23	26.47	22.42	31.45	60.84
	t-5	28.92	22.69	47.32	47.68	51.51	70.82	60.15
Investments	z_{t-i}	**MA(2)**	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
	t-1	**0.27**	0.36	9.54	24.69	9.39	14.17	25.80
	t-2	25.71	27.62	56.52	42.30	65.88	72.11	66.68
	t-3	63.62	73.92	70.80	74.74	82.36	90.02	97.25
	t-4	28.91	42.24	60.84	60.08	98.49	99.81	98.58
	t-5	95.04	64.13	77.31	78.84	56.96	88.74	70.94
Exports	z_{t-i}	**MA(2)**	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
	t-1	**7.92**	42.68	65.28	68.47	47.25	84.03	82.97
	t-2	47.74	74.95	51.78	21.44	88.23	95.54	80.73
	t-3	94.88	92.55	55.31	81.69	95.17	93.58	98.44
	t-4	51.80	94.22	92.96	90.19	86.87	96.51	99.35
	t-5	46.86	98.17	86.31	95.35	98.79	99.90	96.69
Imports	z_{t-i}	**MA(2)**	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
	t-1	2.89	3.78	0.91	0.61	3.06	2.31	5.00
	t-2	8.33	26.78	23.52	2.99	10.72	24.38	64.44
	t-3	9.73	5.66	9.85	11.67	19.79	63.87	53.09
	t-4	11.30	2.39	26.53	17.06	68.56	50.59	29.23
	t-5	**0.56**	8.74	1.01	43.21	44.37	45.44	25.17
Gov. Cons.	z_{t-i}	**MA(2)**	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
	t-1	52.39	67.84	55.39	2.54	3.31	0.21	0.02
	t-2	61.03	9.04	3.12	2.23	0.63	0.02	1.34
	t-3	76.75	14.12	0.93	2.44	0.03	0.65	19.88
	t-4	12.82	0.25	0.84	0.03	2.20	1.91	1.78
	t-5	**0.00**	0.44	0.32	2.63	6.12	7.08	5.13

Notes: P-values (multiplied by 100) of linearity tests of variables of interest (GDP, private consumption, investments, exports, imports and government consumption). The linearity test is run considering different potential transition variable candidate z_{t-i}, such as the lagged *(t-i)* backward-looking moving average (MA) over *(j)* quarter(s) of the output growth rate with $i \in I = 1, .., 5$ and $j \in J = 2, ...8$. In bolds the transition variable corresponding to the smallest p-value that governs the non-linearity of variables indicated in the left column.

TABLE 3.4: Cumulative Tax Multipliers (baseline)

		Cumulative Multipliers					Cumulative Multipliers		
		Recession (1)	Expansion (2)	Linear (3)			Recession (1)	Expansion (2)	Linear (3)
	Horizon					Horizon			
GDP	4	**-1.77**	0.29	**-0.46**	**Priv. Cons.**	4	**-2.00**	0.14	**-1.56**
		[-5.02, -0.79]	[-1.30, 2.21]	[-3.00, 0.63]			[-5.60, -0.98]	[-1.06, 1.57]	[-3.45, -0.90]
	8	**-2.93**	0.68	**-1.67**		8	**-2.94**	0.47	**-1.85**
		[-6.68, -1.73]	[-0.30, 2.60]	[-2.87, -1.04]			[-6.64, -1.80]	[-0.20, 1.83]	[-3.06, -1.25]
	12	**-2.47**	0.14	**-1.81**		12	**-2.50**	0.17	**-2.09**
		[-4.25, -1.64]	[-1.05, 1.62]	[-2.79, -1.24]			[-4.22, -1.17]	[-0.68, 1.31]	[-3.12, -1.53]
	16	**-2.01**	-0.71	**-2.03**		16	**-1.94**	-1.05	**-2.37**
		[-3.05, -1.42]	[-3.13, 0.84]	[-2.88, -1.50]			[-2.89, -1.40]	[-3.86, 0.20]	[-3.29, -1.82]
	20	**-1.72**	**-2.23**	**-2.08**		20	**-1.71**	**-2.40**	**-2.24**
		[-2.47, -1.29]	[-7.82, -0.40]	[-2.91, -1.55]			[-2.41, -1.26]	[-8.28, -0.37]	[-3.00, -1.75]
Investments	4	**-0.88**	-0.42	**-1.00**	**Exports**	4	**-1.98**	1.65	-0.19
		[-2.49, -0.36]	[-1.92, 0.61]	[-2.19, -0.57]			[-5.59, -0.98]	[-2.57, 6.31]	[-0.78, 0.25]
	8	**-1.04**	-0.39	**-0.90**		8	**-1.85**	**1.59**	-0.24
		[-2.35, -0.59]	[-1.36, 0.05]	[-1.53, -0.59]			[-4.16, -1.08]	[0.71, 4.80]	[-0.60, 0.01]
	12	**-0.63**	**-0.61**	**-0.72**		12	**-1.50**	**1.94**	-0.17
		[-1.15, -0.37]	[-2.05, -0.11]	[-1.11, -0.49]			[-2.55, -1.01]	[0.92, 5.94]	[-0.40, 0.02]
	16	**-0.30**	**-1.19**	**-0.63**		16	**-1.43**	**2.58**	**-0.18**
		[-0.55, -0.13]	[-4.15, -0.35]	[-0.92, -0.45]			[-2.12, -1.03]	[1.01, 8.90]	[-0.37, -0.02]
	20	-0.016	**-1.17**	**-0.53**		20	**-1.31**	**2.69**	**-0.26**
		[-0.35, 0.00]	[-6.03, -0.65]	[-0.75, -0.38]			[-1.82, -0.99]	[1.01, 9.22]	[-0.42, -0.14]
Imports	4	**-2.64**	1.21	**-1.05**	**Gov. Cons.**	4	-0.23	1.10	**0.45**
		[-7.30, -1.38]	[-1.74, 4.77]	[-2.37, -0.53]			[-0.78, 0.05]	[-1.70, 4.22]	[0.21, 1.05]
	8	**-2.78**	**1.23**	**-1.04**		8	**-0.62**	**1.68**	**0.47**
		[-6.17, -1.71]	[0.52, 3.80]	[-1.77, -0.67]			[-1.46, -0.32]	[0.81, 5.07]	[0.26, 0.84]
	12	**-2.18**	**1.52**	**-0.92**		12	**-0.50**	**1.95**	**0.42**
		[-3.71, -1.52]	[0.69, 4.79]	[-1.40, -0.64]			[-0.90, -0.29]	[0.94, 6.07]	[0.24, 0.70]
	16	**-1.83**	**1.84**	**-0.84**		16	**-0.22**	**2.26**	**0.47**
		[-2.69, -1.36]	[0.67, 6.38]	[-1.21, -0.62]			[-0.42, -0.07]	[0.86, 7.84]	[0.30, 0.70]
	20	**-1.63**	**1.57**	**-0.75**		20	-0.05	**2.61**	**0.50**
		[-2.27, -1.26]	[0.56, 5.46]	[-1.03, -0.57]			[-0.19, 0.08]	[1.03, 8.90]	[0.32, 0.71]

Notes: the table shows the cumulative multipliers on output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The multipliers are calculated over one to five years and in recessions, expansions, and the linear case, columns (1) to (3) respectively. The 90% confidence intervals are reported in brackets. Bold numbers indicate coefficients statistically significant at the 90% level.

TABLE 3.5: Test

	Horizon	Difference between Multipliers in Recessions and in Expansions (1)		Horizon	(1)
GDP	4	-2.4	**Priv. Cons.**	4	-2.47
		[-7.50, 1.33]			[-7.29, 0.53]
	8	**-4.13**		8	**-3.82**
		[-8.80, -1.72]			[-8.07, -1.86]
	12	**-2.90**		12	**-2.93**
		[-5.51, -0.88]			[-5.18, -1.29]
	16	-1.45		16	-1.02
		[-3.71, -1.42]			[-3.26, 2.22]
	20	0.43		20	0.64
		[-2.51, 6.17]			[-2.24, 6.75]
Investments	4	-0.59	**Exports**	4	-4.47
		[-3.32, 2.06]			[-8.86, 4]
	8	-0.71		8	**-4.03**
		[-2.58, 0.71]			[-8.52, -2.34]
	12	-0.05		12	**-3.82**
		[-1.15, 1.53]			[-7.74, -2.47]
	16	0.88		16	**-4.27**
		[-0.39, 3.87]			[-8.31, -2.54]
	20	**1.60**		20	**-4.18**
		[0.27, 5.89]			[-8.43, -2.53]
Imports	4	**-4.65**	**Gov. Cons.**	4	-1.53
		[-8.19, 2.66]			[-5.00, 2.18]
	8	**-4.65**		8	**-2.61**
		[-8.38, -2.80]			[-6.26, -1.40]
	12	**-4.18**		12	**-2.62**
		[-7.56, -2.81]			[-6.60, -1.50]
	16	**-4.00**		16	**-2.5**
		[-8.21, -2.50]			[-8.06, -1.00]
	20	**-3.45**		20	**-2.68**
		[-7.01, -2.20]			[-8.04, -0.98]

Notes: Empirical densities of the difference between multipliers in recessions and in expansions. Densities based on 10,000 realizations of such differences for each horizon of interest. Bold numbers refer to coefficients statistically significant at the 90% level.

TABLE 3.6: Robustness checks-Tax shock identified via the residual obtained regressing the Cloyne's measure on its own lags and on covariates

		Cumulative Tax Multipliers					Cumulative Tax Multipliers		
		Recession (1)	Expansion (2)	Linear (3)			Recession (1)	Expansion (2)	Linear (3)
	Horizon					Horizon			
GDP	4	**-1.57**	0.04	**-1.35**	**Priv. Cons.**	4	**-2.00**	0.17	**-1.56**
		[-4.78, -0.59]	[-1.80, 1.89]	[-3.04, -0.63]			[-6.00, -0.95]	[-1.76, 1.11]	[-3.41, -0.90]
	8	**-3.19**	0.23	**-1.76**		8	**-3.52**	0.16	**-1.84**
		[-9.00, -1.60]	[-0.96, 1.66]	[-3.03, -1.11]			[-9.81, -1.86]	[-0.65, 1.20]	[-3.08, -1.26]
	12	**-3.20**	0.42	**-1.98**		12	**-3.64**	-0.21	**-2.09**
		[-7.86, -1.83]	[-2.00, 0.67]	[-3.03, -1.37]			[-8.85, -2.13]	[-1.27, 0.57]	[-3.12, -1.54]
	16	**-2.62**	**-1.26**	**-2.23**		16	**-2.75**	**-1.46**	**-2.37**
		[-4.92, -1.66]	[-3.92, -0.18]	[-3.15, -1.63]			[-5.09, -1.81]	[-4.36, -0.56]	[-3.30, -1.81]
	20	**-2.36**	**-2.28**	**-1.93**		20	**-2.49**	**-2.45**	**-2.23**
		[-4.00, -1.53]	[-5.80, -1.11]	[-2.67, -1.44]			[-4.24, -1.70]	[-6.04, -1.33]	[-3.02, -1.75]
Investments	4	**-0.95**	-0.43	**-1.00**	**Exports**	4	**-2.24**	1.79	-0.18
		[-2.83, -0.37]	[-2.02, 0.79]	[-2.14, -0.56]			[-6.56, -1.06]	[-3.41, 6.94]	[-0.77, 0.25]
	8	**-1.32**	-0.45	**-0.91**		8	**-2.45**	**1.69**	-0.24
		[-3.69, -0.563]	[-1.53, 0.03]	[-1.53, -0.60]			[-6.91, -1.26]	[0.74, 5.08]	[-0.60, 0.02]
	12	**-0.90**	**-0.63**	**-0.72**		12	**-2.38**	**1.95**	-0.17
		[-2.25, -0.45]	[-1.95, -0.21]	[-1.10, -0.48]			[-5.76, -1.038]	[0.97, 5.43]	[-0.40, 0.02]
	16	**-0.41**	**-1.11**	**-0.64**		16	**-2.15**	**2.37**	**-0.18**
		[-0.90, -0.15]	[-3.19, -0.50]	[-0.92, -0.44]			[-3.97, -1.42]	[1.25, 6.67]	[-0.37, -0.02]
	20	-0.24	**-1.42**	**-0.53**		20	**-1.97**	**2.17**	**-0.26**
		[-0.57, 0.00]	[-3.59, -0.77]	[-0.76, -0.38]			[-3.30, -1.37]	[1.26, 5.35]	[-0.42, -0.15]
Imports	4	**-3.02**	1.29	**-1.06**	**Gov. Cons.**	4	**-0.37**	1.23	**0.45**
		[-8.81, -1.54]	[-2.38, 5.15]	[-2.36, -0.53]			[-1.23, -0.04]	[-2.28, 4.61]	[0.20, 1.03]
	8	**-3.67**	**1.32**	**-1.05**		8	**-1.00**	**1.85**	**0.47**
		[-8.34, -1.99]	[0.54, 4.02]	[-1.80, -0.68]			[-2.87, -0.48]	[0.90, 5.49]	[0.26, 0.84]
	12	**-3.44**	**1.51**	**-0.92**		12	**-0.92**	**1.93**	**0.42**
		[-8.25, -12.05]	[0.74, 4.37]	[-1.41, -0.65]			[-2.29, -0.49]	0.98, 5.40]	[0.25, 0.70]
	16	**-2.78**	**1.67**	**-0.85**		16	**-0.44**	**2.03**	**0.46**
		[-5.07, -1.84]	[0.85, 4.75]	[-1.21, -0.62]			[-0.90, -0.19]	[1.05, 5.74]	[0.30, 0.70]
	20	**-2.55**	**1.26**	**-0.75**		20	-0.19	**2.15**	**0.49**
		[-4.21, -1.80]	[0.68, 3.11]	[-1.04, -0.57]			[-0.47, 0.01]	[1.22, 5.21]	[0.33, 0.71]

Notes: the table shows the cumulative multipliers on output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The tax shock is identified via the residual obtained regressing the Cloyne's tax shock series on its own four lags and on covariates of the baseline specification. The multipliers are calculated over one to five years and in recessions, expansions, and the linear case, columns (1) to (3) respectively. The 90% confidence intervals are reported in brackets. Bold numbers indicate coefficients statistically significant at the 90% level.

TABLE 3.7: Robustness checks: baseline specification with a quadratic trend

		Cumulative Tax Multipliers					Cumulative Tax Multipliers		
		Recession (1)	Expansion (2)	Linear (3)			Recession (1)	Expansion (2)	Linear (3)
	Horizon					Horizon			
GDP	4	**-1.66**	0.31	**-1.19**	**Priv. Cons.**	4	**-1.64**	-0.04	**-1.34**
		[-4.42, -0.78]	[-1.40, 2.51]	[-2.58, -0.57]			[-4.26, -0.82]	[-1.43, 1.22]	[-2.76, -0.80]
	8	**-2.98**	0.61	**-1.54**		8	**-2.68**	0.14	**-1.68**
		[-6.98, -1.71]	[-0.34, 2.33]	[-2.58, -0.97]			[-6.19, -1.57]	[-0.57, 1.02]	[-2.71, -1.16]
	12	**-2.91**	0.23	**-1.70**		12	**-2.54**	-0.20	**-1.90**
		[-5.76, -1.83]	[-0.87, 1.60]	[-2.56, -1.14]			[-5.01, -1.62]	[-1.16, 0.53]	[-2.82, -1.38]
	16	**-2.72**	-0.36	**-1.87**		16	**-2.11**	**-1.37**	**-2.13**
		[-4.66, -1.80]	[-2.13, 0.95]	[-2.67, -1.34]			[-3.65, -1.42]	[-4.40, -0.40]	[-2.98, -1.62]
	20	**-2.65**	**-1.42**	**-1.80**		20	**-1.97**	**-2.57**	**-2.14**
		[-4.36, -1.81]	[-4.85, -0.01]	[-2.52, -1.31]			[-3.23, -1.35]	[-8.10, -1.12]	[-2.93, -1.65]
Investments	4	**-0.91**	-0.33	**-0.93**	**Exports**	4	**-1.76**	1.63	-0.14
		[-2.40, -0.43]	[-1.71, 0.18]	[-1.92, -0.55]			[-4.63, -0.92]	[-2.63, 6.49]	[-0.63, 0.26]
	8	**-1.24**	-0.27	**-0.88**		8	**-1.81**	**1.47**	-0.20
		[-2.97, -0.68]	[-1.12, 0.18]	[-1.44, -0.60]			[-4.21, -1.03]	[0.68, 4.38]	[-0.51, 0.005]
	12	**-0.97**	-0.36	**-0.72**		12	**-1.72**	**1.83**	-0.14
		[-1.99, -0.57]	[-1.30, 0.07]	[-0.36, -0.05]			[-3.42, -1.09]	[0.97, 5.15]	[-0.16, 0.00]
	16	**-0.67**	**-0.78**	**-0.66**		16	**-1.83**	**2.47**	-0.15
		[-1.21, -0.37]	[-2.58, -0.20]	[-0.33, 0.00]			[-3.12, -1.26]	[1.18, 7.53]	[1.08, 9.22]
	20	**-0.55**	**-1.22**	**-0.58**		20	**-1.81**	**2.63**	**-0.27**
		[-0.96, -0.29]	[-4.16, -0.44]	[-0.85, -0.241]			[-2.90, -1.28]	[1.29, 8.06]	[-0.44, -0.13]
Imports	4	**-2.40**	1.17	**-0.92**	**Gov. Cons.**	4	-0.23	1.21	**0.47**
		[-6.18, -1.33]	[-1.75, 4.77]	[-1.98, -0.47]			[-0.76, 0.01]	[-1.95, 4.59]	[0.24, 0.97]
	8	**-2.83**	**1.17**	**-0.96**		8	**-0.6**	**1.57**	**0.46**
		[-6.50, -1.69]	[0.49, 3.41]	[-1.59, -0.62]			[-1.47, -0.28]	[0.77, 4.54]	[0.27, 0.80]
	12	**-2.58**	**1.47**	**-0.89**		12	**-0.50**	**1.66**	**0.41**
		[-5.08, 1.70]	[0.71, 4.18]	[-1.34, -0.62]			[-1.05, -0.23]	[0.82, 4.79]	[0.124, 0.68]
	16	**-2.40**	**1.85**	**-0.83**		16	-0.10	**1.65**	**0.50**
		[-4.08, -1.68]	[0.86, 5.68]	[-1.18, -0.60]			[-0.35, 0.10]	[0.71, 4.96]	[0.32, 0.73]
	20	**-2.37**	**1.73**	**-0.82**		20	0.20	**1.55**	**0.48**
		[-3.78, -1.72]	[0.79, 5.41]	[-1.14, -0.61]			[0.00, 0.45]	[0.63, 4.97]	[0.32, 0.71]

Notes: the table shows the cumulative response of output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The cumulative multipliers are calculated over one and two years and in recessions, expansions, and in the linear case. The 90% confidence intervals are reported in brackets. Bold numbers indicate the coefficients statistically significant at the 90% level

TABLE 3.8: Robustness check: Tax shock identified by excluding the DC subcomponent from ϵ_t^{Cloyne}

		Cumulative Tax Multipliers					Cumulative Tax Multipliers		
		Recession (1)	Expansion (2)	Linear (3)			Recession (1)	Expansion (2)	Linear (3)
	Horizon					Horizon			
GDP	4	**-1.89**	0.42	**-1.31**	**Priv. Cons.**	4	**-2.00**	0.28	**-1.57**
		[-5.61, -0.82]	[-0.58, 2.01]	[-3.05, 0.60]			[-6.00, -1.01]	[-0.51, 1.42]	[-3.41, -0.90]
	8	**-3.14**	0.72	**-1.73**		8	**-3.03**	**0.58**	**-1.84**
		[-7.02, -1.87]	[-0.01, 2.11]	[-2.98, -1.09]			[-6.79, -1.85]	[0.06, 1.72]	[-3.04, -1.25]
	12	**-2.57**	0.25	**-1.90**		12	**-2.52**	0.39	**-2.11**
		[-4.34, -1.72]	[-0.61, 1.26]	[-2.97, -1.30]			[-4.24, -1.73]	[-0.18, 1.32]	[-3.15, -1.52]
	16	**-2.09**	-0.34	**-2.11**		16	**-1.96**	-0.45	**-2.38**
		[-3.17, -1.48]	[-1.46, 0.46]	[-3.04, -1.54]			[-2.94, -1.43]	[-1.45, 0.13]	[-3.30, -1.81]
	20	**-1.78**	**-1.12**	**-1.80**		20	**-1.71**	**-1.12**	**-2.23**
		[-2.57, -1.28]	[-2.67, -0.37]	[-2.49, -1.35]			[-2.42, -1.27]	[-2.54, -0.51]	[-3.02, -1.76]
Investments	4	**-0.84**	-0.31	**-1.00**	**Exports**	4	**-2.24**	**1.44**	-0.18
		[-2.47, -0.31]	[-1.23, 0.21]	[-0.38, -0.21]			[-6.54, -1.09]	[0.54, 4.63]	[-0.77, 0.24]
	8	**-1.00**	-0.25	**-0.90**		8	**-2.08**	**1.35**	-0.24
		[-2.27, -0.54]	[-0.81, 0.09]	[-0.43, -0.27]			[-4.62, -1.25]	[0.71, 3.28]	[-0.60, 0.01]
	12	**-0.60**	**-0.38**	**-0.72**		12	**-1.68**	**1.60**	-0.16
		[-1.06, -0.32]	[-1.07, -0.03]	[-0.39, -0.24]			[-2.84, -1.13]	[0.90, 3.85]	[-0.41, 0.02]
	16	**-0.26**	**-0.70**	**-0.64**		16	**-1.60**	**1.93**	**0.18**
		[-0.50, -0.08]	[-1.68, -0.32]	[-0.39, -0.24]			[-2.35, -1.16]	[1.13, 4.63]	[-0.37,-0.03]
	20	-0.10	**-0.90**	**-0.53**		20	**-1.44**	**1.77**	**-0.26**
		[-0.30, 0.05]	[-1.88, -0.50]	[-0.40, -0.24]			[-2.00, -1.10]	[1.09, 3.75]	[-0.42, -0.15]
Imports	4	**-2.87**	**1.17**	**-1.04**	**Gov. Cons.**	4	-0.25	**0.95**	**0.45**
		[-8.12, -1.48]	[0.39, 3.88]	[-2.34, -0.54]			[-0.88, 0.03]	[0.37, 3.02]	[0.20,1.03]
	8	**-2.94**	**1.18**	**-1.04**		8	**-0.62**	**1.38**	**0.46**
		[-6.63, -1.82]	[0.61, 2.92]	[-1.75, -0.67]			[-1.46, -0.32]	[0.78, 3.36]	[0.26, 0.83]
	12	**-2.29**	**1.43**	**-0.93**		12	**-0.47**	**1.59**	**0.42**
		[-3.82, -1.59]	[0.79, 3.55]	[-1.40, -0.64]			[-0.84, -0.26]	[0.91, 3.86]	[0.25, 0.69]
	16	**-1.92**	**1.62**	**-0.85**		16	**-0.20**	**1.71**	**0.46**
		[-2.84, -1.43]	[0.94, 3.91]	[-1.20, -0.62]			[-0.37, -0.04]	[0.99, 4.11]	[0.30, 0.70]
	20	**-1.70**	**1.33**	**-0.75**		20	-0.01	**1.77**	**0.48**
		[-2.35, -1.31]	[0.80, 2.84]	[-1.03, -0.57]			[-0.15, 0.11]	[1.09, 3.79]	[0.48, 0.71]

Notes: the table shows the cumulative multipliers of output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The tax shock is identified summing up the tax change components motived by long-run (LR), ideological (ID), and externel (ET) reasons and excluding the deficit consolidation (DC) one. The mean responses are calculated over one and two years and in recessions, expansions, and in the linear case. The 90% confidence intervals are reported in brackets. Bold numbers indicate the coefficients statistically significant at the 90% level.

TABLE 3.9: Robustness check: changing the value of γ such that $F(z_t) \geq 0.90 \approx 0.10$

		Cumulative Tax Multipliers				Cumulative Tax Multipliers	
		Recession	Expansion			Recession	Expansion
		(1)	(2)			(1)	(2)
	Horizon				Horizon		
GDP	4	**-1.65**	0.34	**Priv. Cons.**	4	**-2.03**	0.40
		[-4.93, -0.67]	[-1.99, 2.84]			[-5.81, -0.98]	[-1.31, 2.47]
	8	**-2.94**	0.93		8	**-3.06**	**0.90**
		[-7.02, -1.67]	[-0.32, 3.56]			[-7.33, -1.80]	[0.01, 3.13]
	12	**-2.46**	0.44		12	**-2.61**	0.66
		[-4.50, -1.60]	[-1.10, 2.66]			[-4.69, -1.75]	[-0.50, 2.82]
	16	**-2.01**	-0.40		16	**-1.99**	-0.54
		[-3.179, -1.37]	[-3.16, 2.00]			[-3.13, -1.431]	[-3.08, 1.53]
	20	**-1.71**	-2.00		20	**-1.73**	-2.04
		[-2.55, -1.18]	[-8.01, 4.79]			[-2.53, -1.25]	[-8.62, 4.81]
Investments	4	**-0.88**	-0.34	**Exports**	4	**-2.06**	2.08
		[-2.84, -0.34]	[-1.92, 0.96]			[-6.04, -1.00]	[-5.14, 8.62]
	8	**-1.04**	-0.34		8	**-1.97**	**2.00**
		[-2.59, -0.56]	[-1.44, 0.22]			[-4.78, -1.12]	[0.84, 6.52]
	12	**-0.63**	**-0.63**		12	**-1.58**	**2.48**
		[-0.67, -0.08]	[-3.82, -0.22]			[-2.87, -1.04]	[0.92, 8.42]
	16	**-0.28**	-1.31		16	**-1.50**	3.31
		[-0.55, -0.09]	[-5.19, 3.28]			[-2.33, -1.07]	[-7.20, 8.63]
	20	-0.13	-2.00		20	**-1.37**	3.56
		[-0.33, 0.03]	[-8.44, 4.85]			[-1.97, -1.02]	[-8.48, 8.22]
Imports	4	**-2.77**	**1.67**	**Gov. Cons.**	4	-0.23	**1.06**
		[-8.00, -1.44]	[-4.07, 7.24]			[-0.88, 0.04]	[-2.60, 4.48]
	8	**-2.96**	**1.74**		8	**-0.63**	**1.64**
		[-7.01, -1.7]	[0.71, 5.67]			[-1.57, -0.31]	[0.70, 5.32]
	12	**-2.30**	**2.10**		12	**-0.51**	**1.96**
		[-4.19, -1.58]	[0.77, 7.08]			[-0.97, -0.28]	[0.71, 6.67]
	16	**-1.90**	2.54		16	**-0.22**	2.30
		[-2.94, -1.39]	[-5.66,8.41]			[-0.43, -0.047]	[-5.06, 9.36]
	20	**-1.70**	2.30		20	-0.05	2.17
		[-2.43, -1.28]	[-5.73, 8.25]			[-0.20, 0.08]	[-6.73, 8.88]

Notes: the table shows the mean (cumulative) response of output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The mean responses are calculated over one and two years and in recessions and expansions. The 90% confidence intervals are reported in brackets. Bold numbers indicate the coefficients statistically significant at the 90% level

TABLE 3.10: Robustness check: changing the value of γ such that $F(z_t) \geq 0.80 \approx 0.20$

		Cumulative Tax Multipliers				Cumulative Tax Multipliers	
		Recession (1)	Expansion (2)			Recession (1)	Expansion (2)
	Horizon				Horizon		
GDP	4	**-2.11**	0.11	**Priv. Cons.**	4	**-2.20**	-0.14
		[-5.82, -1.00]	[-1.00, 1.41]			[-5.96, -1.11]	[-1.09, 0.61]
	8	**-3.00**	0.36		8	**-2.91**	0.03
		[-6.20, -1.88]	[-0.41, 1.50]			[-5.83, -1.85]	[-0.57, 0.68]
	12	**-2.52**	-0.20		12	**-2.43**	-0.32
		[-4.00, -1.76]	[-1.20, 0.65]			[-3.81, -1.74]	[-1.19, 0.23]
	16	**-2.07**	**-1.04**		16	**-1.93**	**-1.43**
		[-2.99, -1.50]	[-2.92, 0.22]			[-2.75, -1.44]	[-3.57, -0.70]
	20	**-1.81**	**-2.13**		20	**-1.75**	**-2.40**
		[-2.53, -1.33]	[-4.91, -1.11]			[-2.42, -1.32]	[-5.37, -1.40]
Investments	4	**-0.96**	-0.38	**Exports**	4	**-2.06**	**1.12**
		[-2.76, -0.40]	[-1.56, 0.26]			[-5.68, -1.03]	[0.36, 3.65]
	8	**-1.02**	-0.37		8	**-1.75**	**1.09**
		[-2.11, -0.57]	[-1.18, 0.06]			[-3.62, -1.07]	[0.54, 2.74]
	12	**-0.59**	**-0.61**		12	**-1.42**	**1.33**
		[-1.04, -0.33]	[-1.61, -0.16]			[-2.26, -1.00]	[0.72, 3.17]
	16	**-0.27**	**-1.20**		16	**-1.36**	**1.65**
		[-0.51, -0.08]	[-3.04, -0.60]			[1.94, -1.02]	[0.94, 4.00]
	20	-0.13	**-1.66**		20	**-1.27**	**1.57**
		[-0.33, 0.03]	[-3.77, -0.94]			[-1.73, -0.98]	[0.93, 3.51]
Imports	4	**-2.69**	0.60	**Gov. Cons.**	4	-0.17	**0.91**
		[-7.26, -1.44]	[-0.02, 2.09]			[-0.64, 0.11]	[0.39, 2.87]
	8	**-2.62**	**0.63**		8	**-0.49**	**1.31**
		[-5.35, -1.68]	[0.20, 1.69]			[-1.08, -0.25]	[0.73, 3.23]
	12	**-2.08**	**0.83**		12	**-0.40**	**1.52**
		[-3.27, -1.50]	[0.38, 2.02]			[-0.70, -0.23]	[0.86, 3.57]
	16	**-1.76**	**0.95**		16	**-0.17**	**1.75**
		[-2.47, -1.34]	[0.48, 2.40]			[-0.37, -0.04]	[1.01, 4.28]
	20	**-1.61**	**0.65**		20	-0.01	**1.91**
		[-2.18, -1.27]	[0.28, 1.57]			[-0.14, 0.11]	[1.16, 4.27]

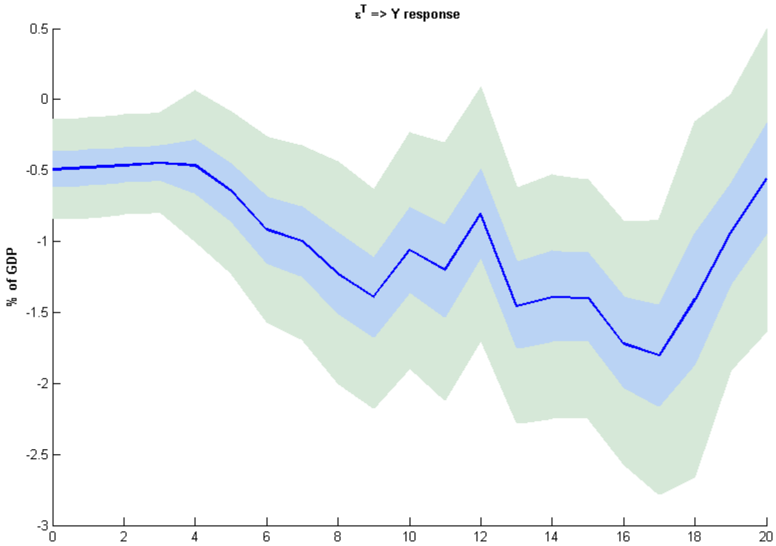
Notes: the table shows the cumulative multipliers of output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The multipliers are calculated over one and two years and in recessions and expansions. The 90% confidence intervals are reported in brackets. Bold numbers indicate the coefficients statistically significant at the 90% level

TABLE 3.11: Robustness checks: baseline augmented by additional variables (interest rate and inflation rate)

		Cumulative Multipliers					Cumulative Multipliers		
		Recession (1)	Expansion (2)	Linear (3)			Recession (1)	Expansion (2)	Linear (3)
	Horizon					Horizon			
GDP	4	**-1.03**	0.02	**-0.90**	**Priv. Cons.**	4	**-1.22**	0.07	**-1.56**
		[-2.27, -0.40]	[-2.24,2.26]	[-1.83,-0.40]			[-2.36, -0.59]	[-1.49, 1.84]	[-2.28, -0.74]
	8	**-2.37**	0.39	**-1.29**		8	**-2.31**	0.21	**-1.60**
		[-4.66, -1.40]	[-1.26, 2.36]	[-2.11, -0.82]			[-4.50, -1.46]	[-0.89, 1.50]	[-2.50, -1.14]
	12	**-2.71**	-0.44	**-1.42**		12	**-2.83**	**-0.14**	**-1.85**
		[-5.90, -1.54]	[-1.63, 0.32]	[-2.10, -0.98]			[-6.10, -1.72]	[-0.82, 0.38]	[-2.63, -1.38]
	16	**-1.84**	**-1.90**	**-1.66**		16	**-1.77**	**-1.47**	**-2.07**
		[-3.22, -1.12]	[-6.68, -0.37]	[-2.29, -1.23]			[-3.02, -1.15]	[-5.10, -0.34]	[-2.78, -1.62]
	20	**-1.36**	**-3.52**	**-1.54**		20	**-1.40**	**-2.61**	**-2.00**
		[-2.26, -0.77]	[-9.81, -1.53]	[-2.07, -1.16]			[-2.20, -0.90]	[-8.81, -1.13]	[-2.60, -1.60]
Investments	4	-0.22	-0.73	**0.82**	**Exports**	4	**-2.11**	1.96	-0.19
		[-0.71, 0.12]	[-4.04, 2.71]	[-1.54, -0.49]			[-4.09, -1.28]	[-7.11, 8.24]	[-0.54, 0.23]
	8	**-0.55**	-0.81	**-0.81**		8	**-2.45**	2.20	-0.24
		[-1.21, -0.21]	[-3.27, 1.32]	[-1.29, -0.55]			[-4.73, -1.56]	[-3.77, 8.28]	[-0.50, 0.02]
	12	-0.39	**-0.92**	**-0.66**		12	**-3.23**	**1.90**	-0.17
		[-1.06, 0.00]	[-0.98, -0.46]	[-1.11, -0.49]			[-6.78, -2.01]	[1.03, 4.99]	[-0.38, 0.01]
	16	0.08	**-1.85**	**-0.60**		16	**-2.80**	**2.76**	**-0.18**
		[-0.19, 0.39]	[-6.40, -0.72]	[-0.84, -0.43]			[-4.56, -1.96]	[1.11, 9.37]	[-0.33, -0.03]
	20	**0.34**	**-2.45**	**-0.53**		20	**-2.49**	**0.50**	**-0.26**
		[0.10, 0.65]	[-8.08, -1.15]	[-0.69, -0.37]			[-3.66, -1.83]	[1.29, 9.27]	[-0.37, -0.12]
Imports	4	**-2.09**	1.13	**-0.83**	**Gov. Cons.**	4	**-0.26**	1.26	**0.42**
		[-4.00, -1.30]	[-4.30, 6.00]	[-1.63, -0.44]			[-0.65, -0.02]	[-4.49, 6.43]	[0.22, 10.83]
	8	**-2.88**	1.38	**-0.93**		8	**-0.90**	2.09	**0.38**
		[-5.54, -1.87]	[-2.38, 5.21]	[-1.49, -0.62]			[-1.78, -0.53]	[-3.60, 8.02]	[0.22, 0.66]
	12	**-3.68**	1.30	**-0.85**		12	**-1.26**	**1.84**	**0.37**
		[-7.74, -2.33]	[0.67, 3.39]	[-1.25, -0.61]			[-2.73, -0.74]	[1.02, 4.76]	[0.21, 0.58]
	16	**-2.86**	**1.71**	**-0.80**		16	**-0.69**	**2.44**	**0.44**
		[-4.63, -2.02]	[0.64, 5.90]	[-1.10, -0.60]			[-1.20, -0.41]	[1.01, 8.27]	[0.30, 0.64]
	20	**-2.49**	**1.41**	**-0.75**		20	**-0.35**	**2.58**	**0.48**
		[-3.66, -1.85]	[0.59, 4.65]	[-0.98, -0.58]			[-0.63, -0.15]	[1.20, 8.48]	[0.33, 0.68]

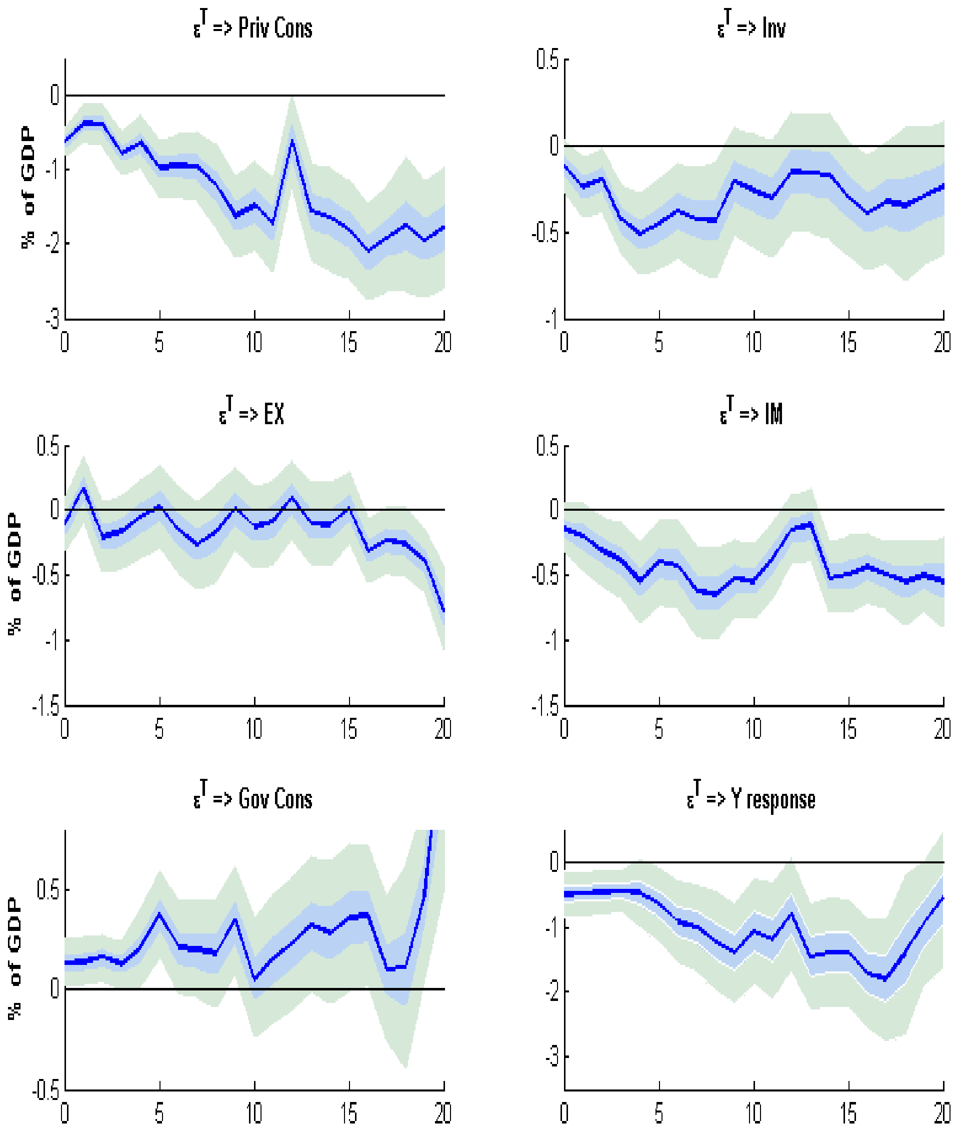
Notes: the table shows the cumulative multipliers on output, private consumption, investments, exports, imports, government consumption to a tax shock corresponding to 1% of GDP. The multipliers are calculated over one to five years and in recessions, expansions, and the linear case, columns (1) to (3) respectively. The 90% confidence intervals are reported in brackets. Bold numbers indicate coefficients statistically significant at the 90% level.

FIGURE 3.1: Response of GDP to an exogenous tax shock



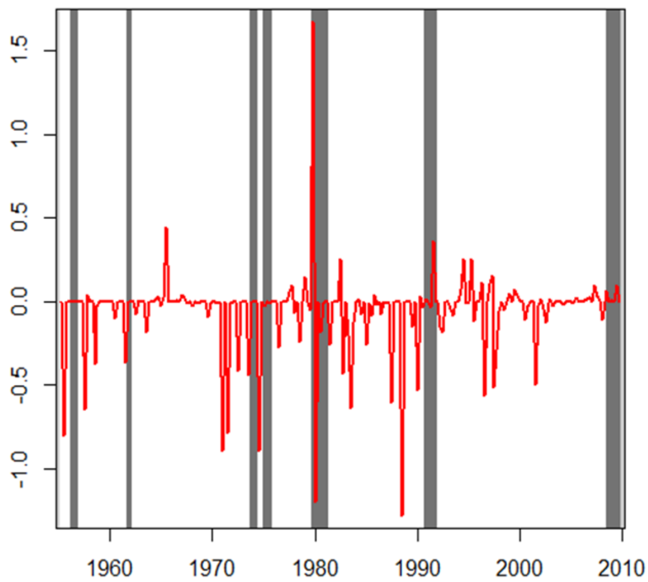
Notes: Figure shows the response of output to a tax shock corresponding to 1% of GDP. Blue lines indicate the IRFs. The lightand dark shaded areas refer to the 68% and 90% confidence intervals, respectively.

FIGURE 3.2: Response of GDP components to an exogenous tax shock



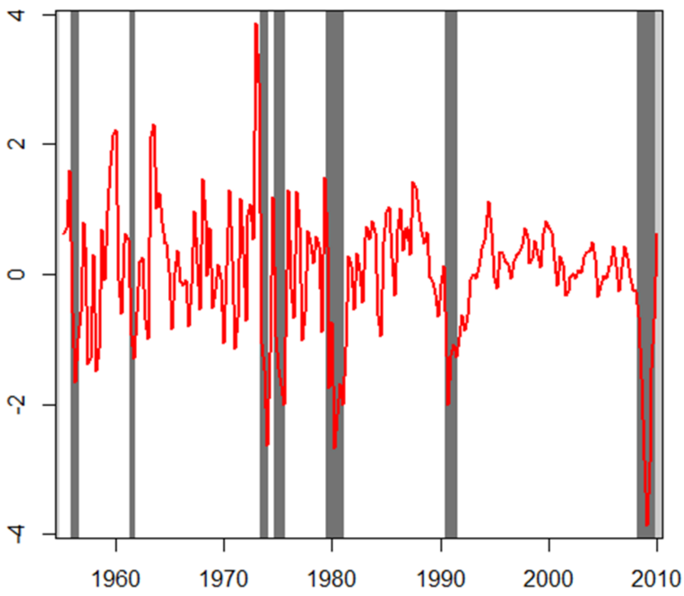
Notes: Figure shows the response of private consumption, investments, exports, imports, government consumption and output to a tax shock corresponding to 1% of GDP. Blue lines indicate the IRFs. The light and dark shaded areas refer to the 68% and 90% confidence intervals, respectively.

FIGURE 3.3: Tax shock vs Business cycle



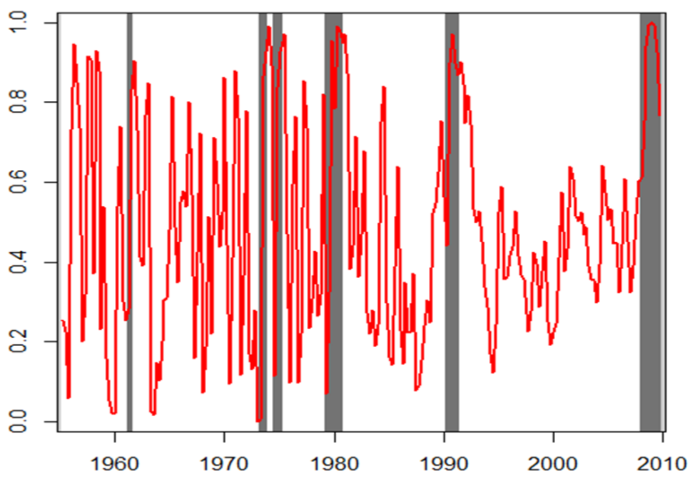
Notes: The shaded area indicate the UK recessionary phases (1955:I-2009:IV) identified by applying the BBQ algorithm, whereas the red lines refers to the tax shock measure of Cloyne (2013).

FIGURE 3.4: Transition variable versus Business Cycle Dates



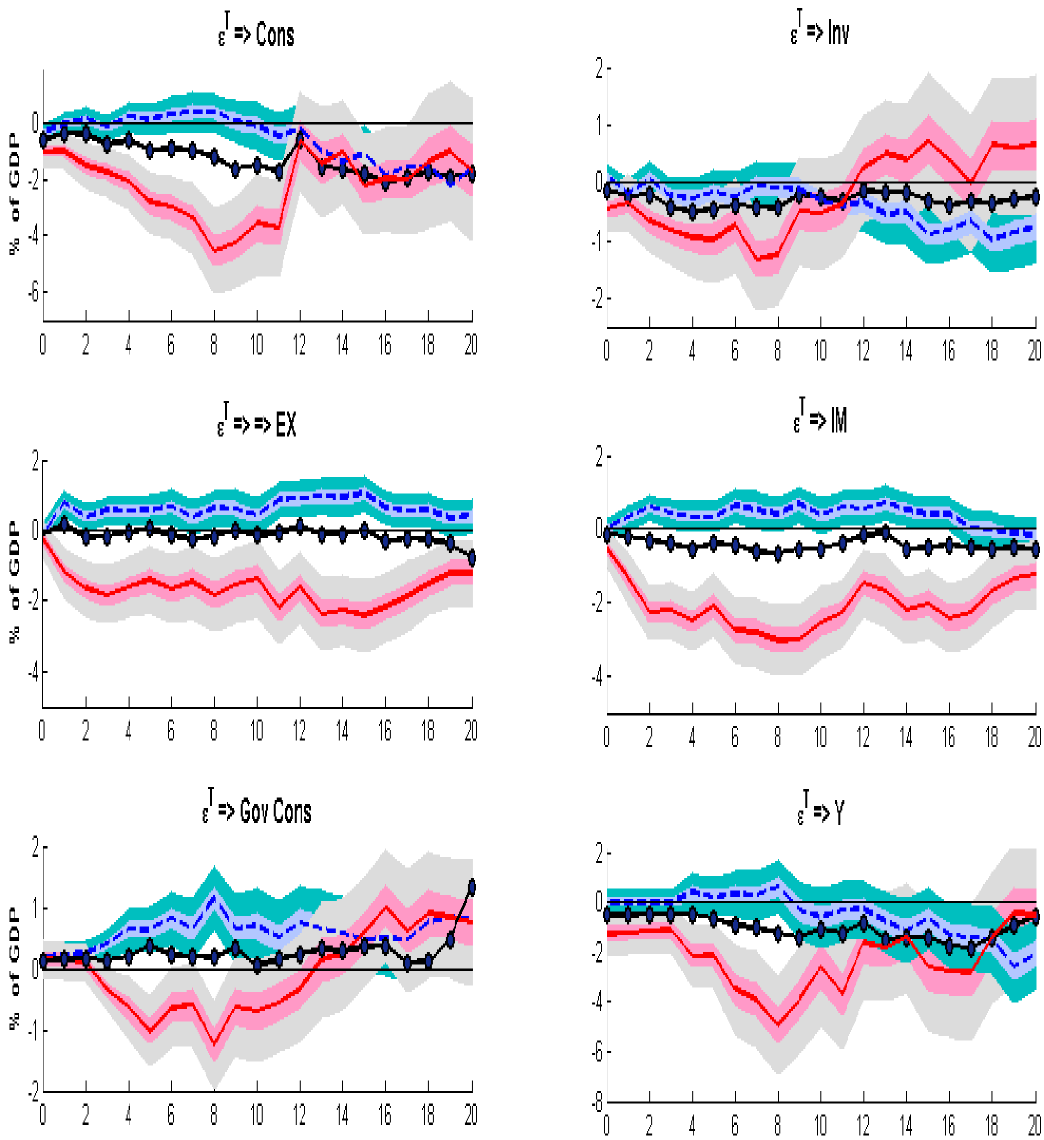
Notes: the transition variable is the standardized backward-looking moving average constructed with two realizations of the quarter-on-quarter real GDP growth rate. Shaded area referes to the recessionary phase identified applying the BBQ algorithm (Harding and Pagan, 2002)

FIGURE 3.5: Probability of being in a recessionary phase



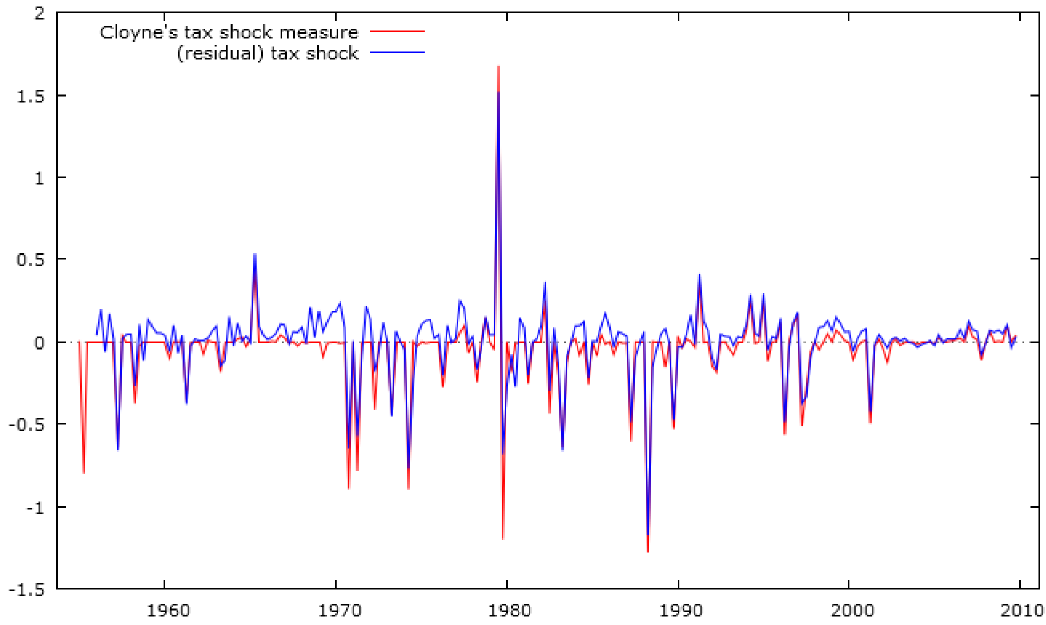
Notes: $F(z_t)$ computed according to the logistic function presented in the text. The transition variable is the standardized backward-looking moving average constructed with two realizations of the quarter-on-quarter real GDP growth rate. The value of the slope parameter is 1.7.

FIGURE 3.6: Response of GDP components to an exogenous tax shock in recessions and expansions



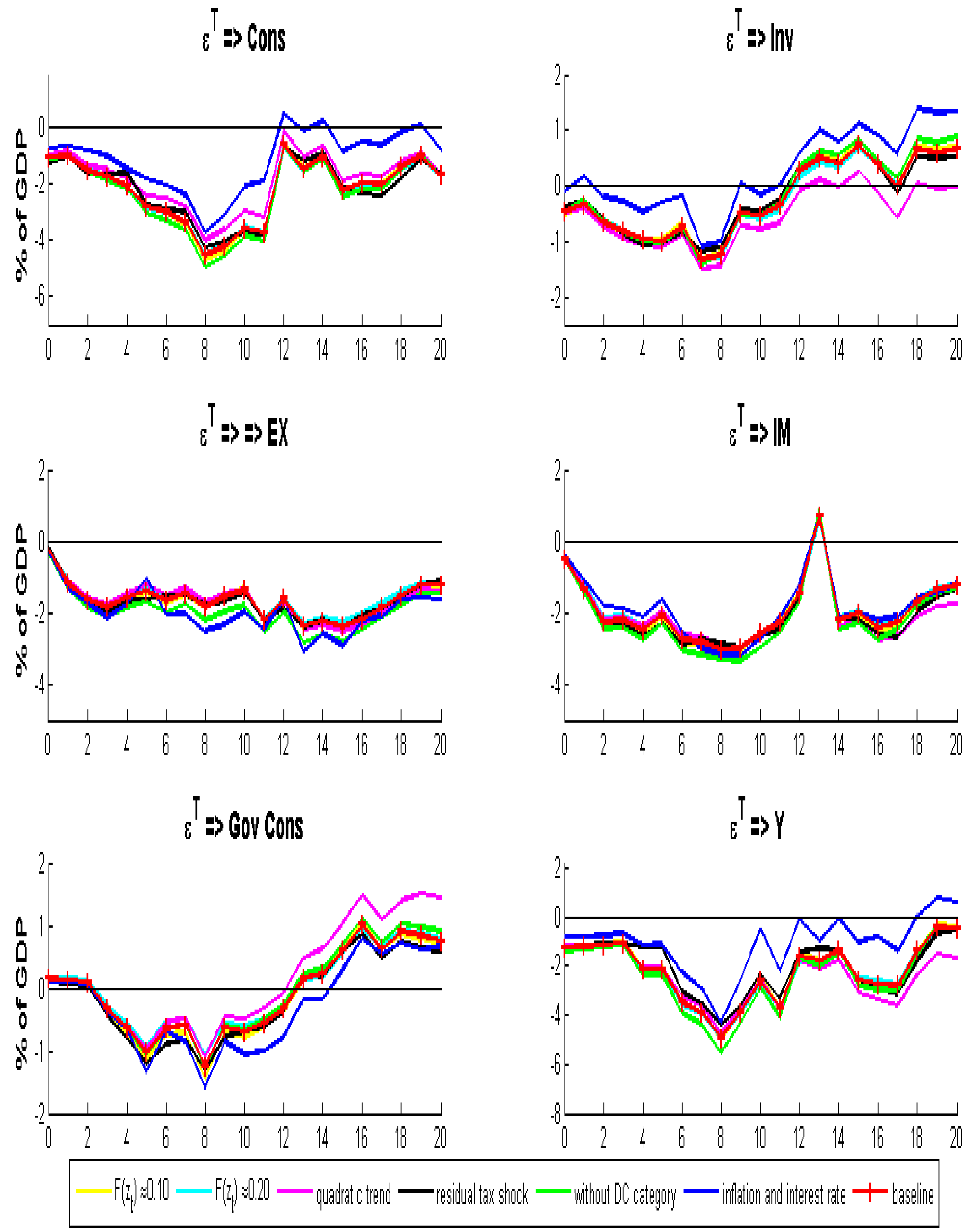
Notes: Figure shows the response of private consumption, investments, exports, imports, government consumption and output to a tax shock corresponding to 1% of GDP. Black dotted lines refer to IRFs computed in a linear specification, whereas the pink and the blue dotted ones to IRFs in recessions and expansions, respectively. The dark and light shaded area refer to the 68% and 90% confidence intervals, respectively.

FIGURE 3.7: Cloyne shocks and its exogeneity



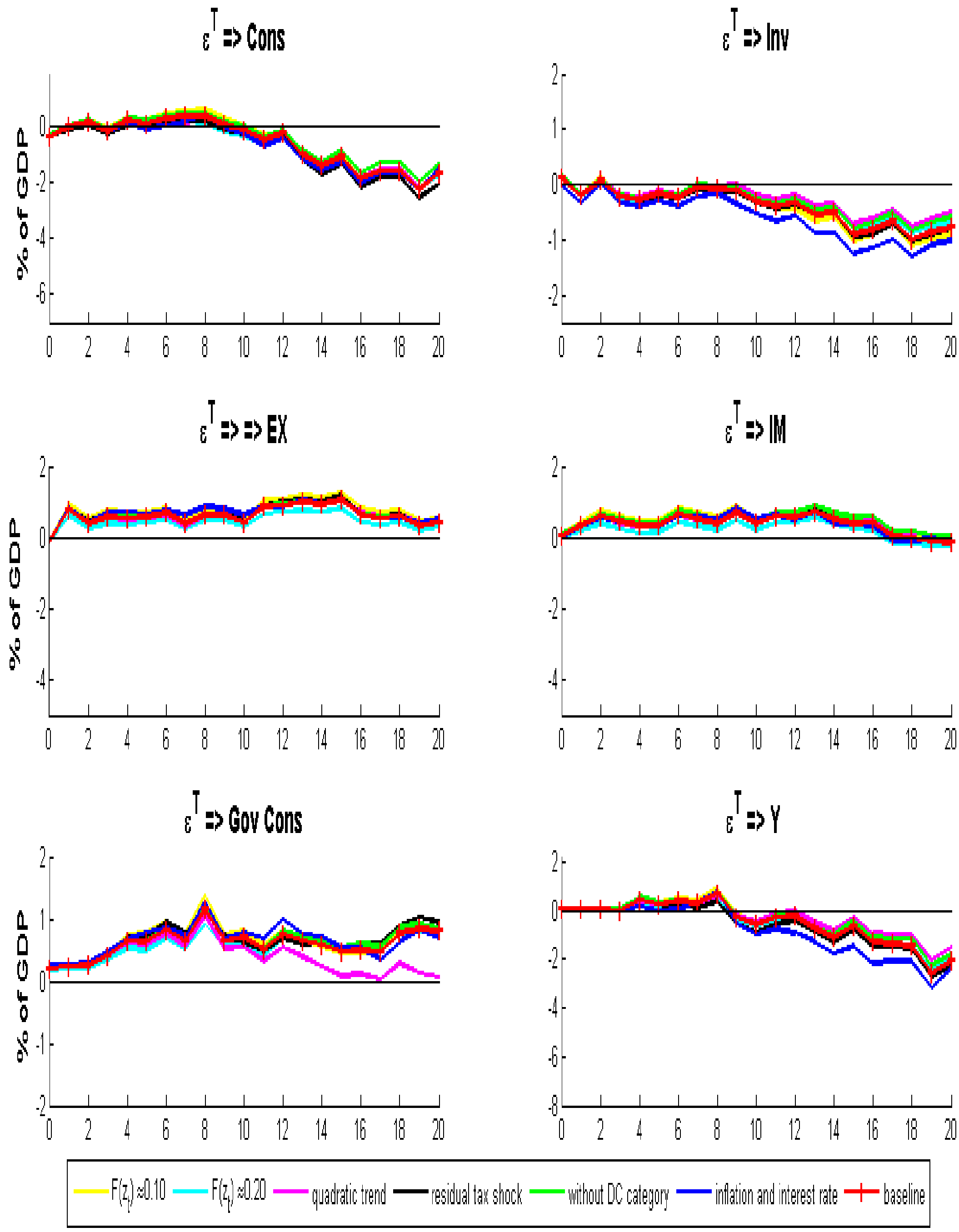
Notes: The figure plots the tax shock series constructed by Cloyne (2013), in our notation ϵ_t^{Cloyne}, versus the (residual) tax shock series obtained regressing the ϵ_t^{Cloyne} on its own four lags and four lags of the log real GDP, revenues and government consumption.

FIGURE 3.8: Robustness checks: response of GDP and its components to an exogenous tax change in recessions



Notes: Figures show the impulse responses of private consumption, investments, exports, imports, government consumption and output to a tax shock corresponding to 1% of GDP under alternative specifications.

FIGURE 3.9: Robustness checks: response of GDP and its components to an exogenous tax change in expansions



Notes: Figures show the impulse responses of private consumption, investments, exports, imports, government consumption and output to a tax shock corresponding to 1% of GDP under alternative specifications.

Appendix 3A

Output Elasticity of Revenues and conversion factors of elasticity into multiplier: Do they matter for the Tax Multiplier?

The main challenge in estimating the tax multiplier is to disentangle a tax change due to a discretionary fiscal policy from a nondiscretionary component, e.g. the change in taxes due to a change in output. Two methods have been proposed in the literature. The first one relies on the SVAR model and based mainly on the identification assumption scheme pioneered by Blanchard and Perotti (2002). The second one identifies an exogenous tax change using a narrative method (Romer and Romer, 2010; Mertens and Ravn, 2014; Cloyne, 2013). Despite several studies investigating tax multipliers there is not a shared view. Perotti (2005), identifying a tax shock via coefficient restrictions, finds that following a tax shock the UK GDP decrease. Cloyne (2013) identify a tax shock through the narrative approach and finds that a tax cut stimulates the economy. In general, the size and duration of a tax shock vary across studies and the estimated tax multiplier via a SVAR model tend to be lower than the narrative approach.

Caldara and Kamps (2008) show that contrasting US fiscal multiplier estimations are likely due to different assumptions on the size of the elasticity of tax revenues to GDP. The first question addressed in this section is whether the UK output tax multiplier is "output elasticity of taxes dependent". To do that we employ in very basic (linear) SVARs two measure of the UK automatic stabilizer proposed in the literature by Perotti (2005) and Cloyne (2013).

Consider a simple three-variate VAR as in Blanchard and Perotti (2002)[26] of the form:

$$X_t = C(L)X_{t-1} + u_t \quad t = 1, ..., T \tag{3A.1}$$

where X_t includes four lags of the log real per capita government consumption, tax revenues and GDP and u_t is a three-dimensional vector of residuals. Equation (3A.1) includes also a constant and a linear time trend. Following the approach proposed by Perotti (2005), the reduced form innovations of vector u_t is expressed as a linear combination of the structural shocks such that:

$$u_t^G = \alpha_Y^G u_t^Y + \beta_T^G \epsilon_t^T + \epsilon_t^G \tag{3A.2}$$

$$u_t^T = \alpha_Y^T u_t^Y + \beta_G^T \epsilon_t^G + \epsilon_t^T \tag{3A.3}$$

$$u_t^Y = \alpha_G^Y u_t^G + \alpha_T^Y u_t^T + \epsilon_t^Y \tag{3A.4}$$

Since the aim is estimating the effect of a tax shock ϵ_t^T on the GDP, let us focus on equation (3A.3). It states that unexpected movement in taxes at time t may be due to output innovations (u_t^Y), structural shocks to government consumption (ϵ_t^G) or to taxes (ϵ_t^T) . Hence, the coefficients α_j^i capture the elasticity of variable i to the variable j, while coefficients β_j^i capture possible link between structural shocks to fiscal variable which may arise whether, for instance, government consumption instantaneously responds to revenues change with government consumption adjustment. The identification of a tax shock is based on the Au_t=Bϵ_t scheme and on some restrictions on the matrix A and B to map from innovations u_t^T to the

[26] Notice that Perotti (2005) estimates for the UK a five-variable VAR. Caldara and Kamps (2008) show that the different results in the literature about the US fiscal multipliers are not due to difference in the specification of the reduced-form models but to the different identification strategies. We stress that the exercise provided in this section is not aimed at choosing the best specification for our analysis but to understand whether the Caldara and Kamps' result is valid also for the UK economy. Hence, we estimate a more parsimonious VAR, as in Blanchard and Perotti (2002).

structural shocks ϵ_t^T. Expressed in matrix notation:

$$\begin{bmatrix} 1 & 0 & -\alpha_Y^G \\ 0 & 1 & -\alpha_Y^T \\ -\alpha_Y^G & -\alpha_T^Y & 1 \end{bmatrix} \begin{bmatrix} u_t^G \\ u_t^T \\ u_t^Y \end{bmatrix} = \begin{bmatrix} \sigma^G & \beta_T^G & 0 \\ \beta_G^T & \sigma^T & 0 \\ 0 & 0 & \sigma^Y \end{bmatrix} \begin{bmatrix} \epsilon_t^G \\ \epsilon_t^T \\ \epsilon_t^Y \end{bmatrix} \tag{3A.5}$$

The identification of fiscal shock in Blanchard and Perotti (2002) relies on some assumptions about the reaction lags and the structural elasticity. Suppose that there is a negative output shock. To offset such shock a fiscal policy action should be planned, approved by the House of Commons and then implemented. It should take more than one quarter to apply a discretionary fiscal policy. With quarterly data the contemporaneous response of the government spending to an output shock can be set to zero. Also, the implementation lags imply that coefficient α_Y^T captures only the automatic elasticity of the tax revenues to GDP due to a fluctuation in the tax base. Thus, the cyclically-adjusted tax innovation, u_t^{CAT}, is given by the difference between the tax innovation (u_t^T) and the output elasticity of revenues (α_Y^T). Notice that the restricted value of the coefficient α_Y^T is obtained by an out-of-model information.[27] Hence, the cyclical adjusted tax innovation derives from an instrumental variable estimation. The structural shock ϵ_t^T is recovered imposing a recursive order on matrix A, on which we assume that tax shock "comes first" than government spending one. Thus, we set β_G^T=0.[28]

Estimated impulse response functions obtained via the estimation of (3A.5) allow us to address a second problem, the *ex post* conversion factor's, raised by Owyang, Ramey, and Zubairy (2013). It is related to the estimation of fiscal multipliers. In general, it is common practice in the fiscal multiplier literature to run SVARs using

[27]Perotti (2005) calibrates the value of the automatic stabilizer for the UK economy through the OECD method and assumptions proposed by Giorno, Richardson, Roseveare, and van den Noord (1995) and van den Noord (2002). The output elasticity of output is calibrated combining the estimation of elasticity of tax revenues to their tax base with the elasticity of tax base to output. The corporate and indirect taxes is equal to 1 by assumption. Moreover, the computation of the automatic stabilizer excludes output elasticity to GDP cyclical effects on tax expenditure, income of self-employed, capital gains, for example. See Perotti (2005) and Mertens and Ravn (2014) in depth analysis.

[28] Our results are robust to the alternative specification that government spending "comes first" than taxes shock. The results are available upon request.

log-transformed variables, and then to convert estimated elasticities into multipliers via (*ex post*) conversion factors, e.g. the average of the ratio GDP/(fiscal variable). However, Owyang, Ramey, and Zubairy (2013) highlight that different sample size may rend different conversion factor values, which may lead to biased fiscal multiplier estimates.

Output Elasticity of Revenues and conversion factors of elasticity into multiplier: Do they matter for the Tax Multiplier? This paper address the two problems through simple exercises. We consider two sample sizes spanning one from 1963:I to 2001:II, and the other one from 1955:I to 2009:IV. For each sample size, we estimate two SVARs including the quarterly log of the real government consumption, tax revenues and GDP,[29]and imposing the implementation lag coefficient restrictions discussed above. Notice that for each sample size two alternative coefficient restrictions of output elasticity of taxes (α_Y^T) are set. On the one hand, we set α_Y^T=0.76 as in Perotti (2005). On the other hand, we fix α_Y^T=1.61 as estimated by Cloyne (2013) using narrative data on tax changes.[30] To obtain tax multipliers from (four) estimated SVARs, we convert estimated elasticities (since our variables are expressed in logarithm terms) into multipliers via *ex post* conversion factors. For each sample size and value of α_Y^T, we convert the elasticity into multipliers using the minimum, the mean and the maximum value of the average of the ratio GDP/T of the sample size under analysis.[31] The 90% confidence intervals are computed using 10,000 bootstrapping replications. Figure 3A.1 depicts the IRFs. The top panels show the response of output to a tax shock for the period 1963:I to 2001:II, whereas the bottom ones for the period spanning from 1955:I to 2009:IV. The left hand side panels depict the results for the two different sample when $\alpha_Y^T = 0.76$ (Perotti), whereas the right ones when $\alpha_Y^T = 1.61$ (Cloyne). Each plot reports the point estimates multiplied by different conversion factors: the mean (blue diamond line), the minimum (red line), and the maximum (dotted

[29]All SVARs are estimated including a constant and a linear time trend.

[30]The exercises on two samples are justified because the Perotti's output elasticity of taxes is calibrated for the period 1963:I to 2001:II, whereas the estimation of Cloyne is related to the sample 1955:I-2009:IV.

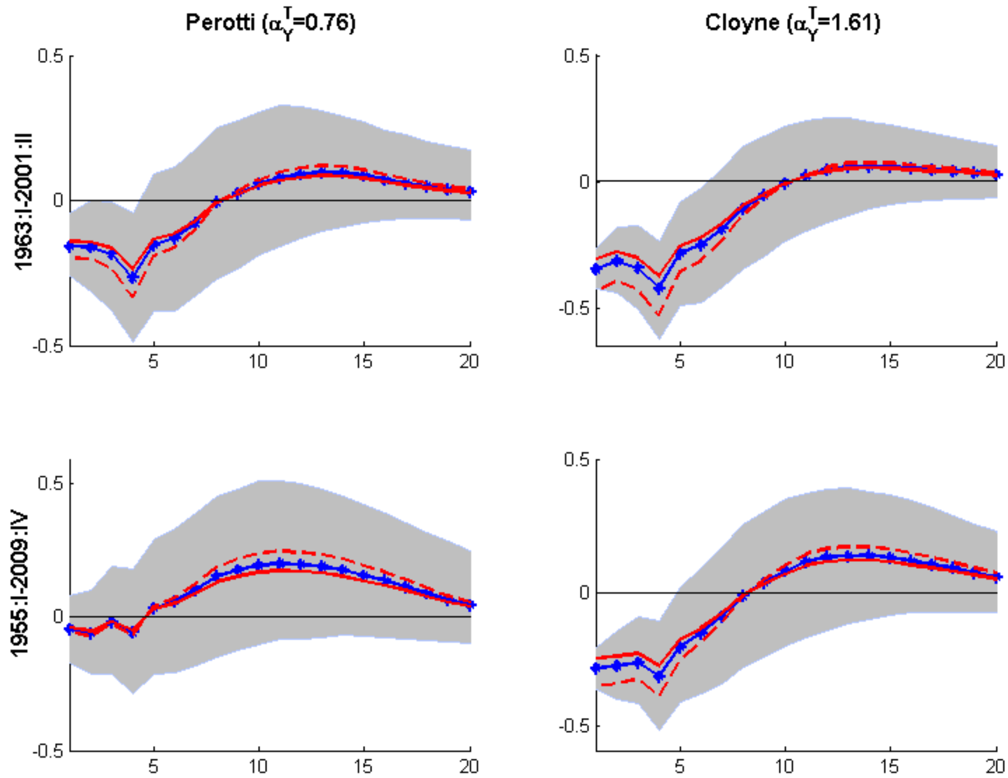
[31]For the sample size 1963:I-2001:II the mean, minimum and maximum of the ratio GDP/T are 3.19, 2.83 and 3.9, respectively. Regard to the sample size 1955:I to 2009:IV, the mean, minimum and maximum of the above ratio are 3.25, 2.83 and 4.02, respectively.

red line). Figure 3A.1 show that, for each sample size, increasing the value of α_Y^T the impact of tax shock on output increases. Moreover, the value of α_Y^T has an impact on the persistence of the shock. That is evident whether we consider table 3A.1 which reports the cumulative multipliers, for the different sample size, coefficient restrictions, and conversion factors. Let us focus on the sample size A (1963:I -2001:II) and on the row reporting tax multipliers using as conversion factor the mean of GDP/T, mean (A). The estimated 1-year integral multiplier (4Q) is -0.3 setting $\alpha_Y^T = 0.76$, whereas doubles setting $\alpha_Y^T = 1.61$. Moreover, the value of α_Y^T has effects on the persistence of the shock. Indeed, whereas the 2-year integral multiplier (8Q) tax multiplier is not statistically significant for $\alpha_Y^T = 0.76$, it is for $\alpha_Y^T = 1.61$. Turning on panel B, the 1-year integral tax multiplier (4Q) is statistically significant only when $\alpha_Y^T = 1.61$. Further, using different value of conversion factors affects the size of output tax multipliers. For example, this bias is evidence focusing on the 2-year integral multipliers (8Q) of panel A, for which tax multipliers range below and above one.

The results show that using the same dataset, the same estimation's method but different coefficient restrictions on the output stabilizer yield different results. This is consistent with Caldara and Kamps (2008): the fiscal multipliers change according to the calibration of the output elasticity of taxes. Moreover, the combination of coefficient restrictions with the value of *ex post* conversion factors may lead other bias on tax multiplier estimates.

An exogenous tax change measure based on the narrative method does not require imposing restrictions on the output elasticity of taxes. A solution to avoid *ex post* conversion problem is to convert GDP and taxes to the same units *ex ante* the estimation. Hence, we identify the tax shock via the tax shock measure proposed by proposed by Cloyne (2013), and to avoid bias on tax multipliers we transform the variables as in Hall (2009), Barro and Redlick (2011) and Owyang, Ramey, and Zubairy (2013).

FIGURE 3A.1: Perotti and Cloyne's output elasticity of revenues in a SVAR specification



Notes: Top panels show the response of output to a tax shock for the period 1963:I to 2001:II, whereas the bottom panels the one for the period 1955:I-2009:IV. The left panels refer to the case in which the output elasticity is set to 0.76 Perotti (2005), whereas the right ones refer to an automatic stabilizer set to 1.61 Cloyne (2013). The blue, red and dotted red lines depict the IRFs obtained using as *ex post* conversion factor the average, the minimum and the maximum of the ratio of GDP to revenues, respectively.

TABLE 3A.1: Multipliers estimated relying on different sample size, coefficient restrictions, and conversion factors

Sample size	CF	Perotti ($\alpha_Y^T = 0.76$)		Cloyne ($\alpha_Y^T = 1.61$)	
		4Q	8Q	4Q	8Q
(A) 1963:I-2001:II	min (A)	**-0.22**	-0.39	**-0.56**	**-0.88**
	mean (A)	**-0.30**	-0.44	**-0.62**	**-1.00**
	max (A)	**-0.37**	-0.55	**-0.79**	**-1.24**
(B) 1955:I-2009:IV	min (B)	-0.06	0.05	**-0.40**	-0.59
	mean (B)	-0.07	0.06	**-0.48**	-0.67
	max (B)	-0.09	0.07	**-0.60**	-0.83

Notes: Top rows show the response of output to a tax shock for the period 1963:I to 2001:II, whereas the bottom rows the one for the period 1955:I-2009:IV. The left column refers to the case in which the output elasticity is set to 0.76 Perotti (2005), whereas the right ones refers to an automatic stabilizer set to 1.61 Cloyne (2013). Bold numbers indicate the coefficients statistically significant at 90%.

Appendix 3B

Business cycle identification via BQQ

There is not in the UK an official dating Committee, as the NBER, which has established an expansion and recession chronology and which has been recognized as an authoritative dating of the cycle.[32] The NBER (2001) defines a recession as "a significant decline in activity spread across the economy, lasting more than few months, visible in industrial production, employment, real income, and wholesale-retail trade. A recession begins just after the economy reaches a peak of activity and ends as the economy reaches its trough". According to the literature turning points can be defined in terms of absolute decline in output (classical cycle) or in terms of deviation of GDP growth rate from its trend (deviation cycle). The deviation-from-trend approach, as in Cooley and Prescott (1995) and Stock and Watson (2008), requires detrending a series. However, several detrending methods exist. For example, the NBER uses the phase-average trend method (PAT), the macroeconomists use Hodrick-Prescott filter or the "band pass" to remove deterministic/stochastic trend. According to Canova (1998) the identified business cycle may depend on the filter used. Moreover, Harding and Pagan (2002) highlight that smoothing methods are aimed at simplifying turning point identification removing idiosyncratic variation. Thus, if turning points are detected using quarterly data series the utility of smoothing methods decreases with such frequency data. Hence, we identify turning points relying on the classical cycle approach. We use the dating algorithm proposed by Harding and Pagan (2002) which is the quarterly version of the well-known monthly Bry and Boschan (1971) algorithm.[33]

[32] In 2002 CEPR established a Business Cycle Dating Committee for the euro area.

[33] The BBQ is one of the most widespread algorithm in detecting turning points. For example, Artis, Marcellino, and Proietti (2002) rely on the BBQ to analyse the characteristic of the business cycle. However, there are other algorithm that we may use to date turning points, for example a Markov Switching model Hamilton (1998). As point out by Harding and Pagan (2002) the Markov Switching model depends on the relative statistical framework.

The BBQ algorithm isolates local minimum and maximum points in a quarterly series, via some constraints. First of all, a local peak (trough) occurs at time t when

$$y_t > (<) y_{t \pm k} \tag{3A.6}$$

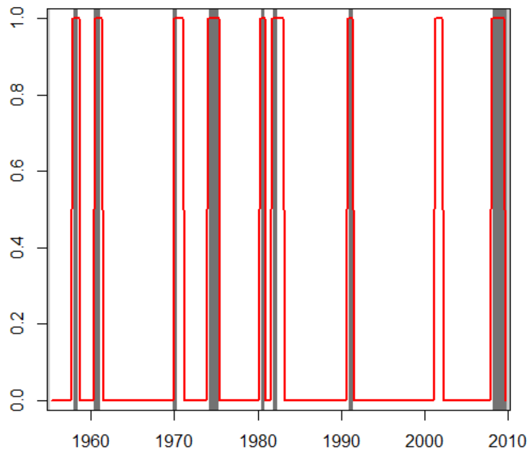
where k=1,2,..K. K allows y_t to be a local local peak (trough) to two quarters on either side.[34] Secondly, the phases alternate between peak and trough. This because whether the phases alternate, then it is possible to distinguish the phase of recession (from peak to trough) from the expansion one (from trough to peak).[35] Thirdly, a complete cycle (from peak to peak or from trough to trough) lasts at least n quarters. The last two rules are known as censoring rule.

To verify the validity of the BBQ a natural exercise is to apply the algorithm to the US for which exists an official chronology. We set for the US k=2 and the duration of the complete cycle to five quarters, as in Harding and Pagan (2002), and we apply the algorithm to the log-real GDP. Then, the turning points are compared with the NBER data. Figure 3B.1 plots the NBER turning points (red lines) versus the turning point identified by the BBQ ones (shaded area). Since 1955 to 2009 the NBER has recorded 9 recessions, whereas the algorithm does not capture the turning points of 2001. Stock and Watson (2010) report that the NBER committee for dating relies on the quarterly real GDP and on four monthly variables, such as real personal income less transfer, real manufacturing and wholesaleretail trade sales, industrial production, and nonfarm employment. They highlight that those series do not receive same weight in the dating procedure.

[34] Notice that larger is the value of K the more restrictive is the definition of the turning points (Harding (2008)).

[35] Harding (2008) show that in UK the frequency of non-alternating turning points is four times higher than the US

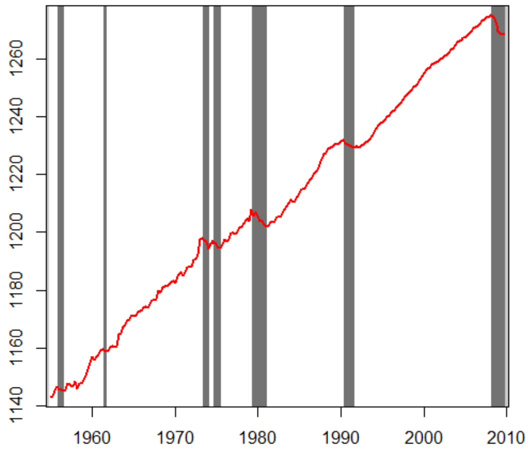
FIGURE 3B.1: US Business Cycle Chronology (NBER vs BBQ algorithm)



Notes: The shaded area indicate the recession phases identified by the BBQ algorithm, whereas the red lines show the NBER business cycle chronology. The sample size spans from 1955:I to 2009:IV

Moreover, Harding (2003) shows that the procedure and variables used by the NBER for the business cycle chronology has changed over time. Also, Harding (2008) report that in detecting the turning points the NBER uses not only committee's procedure but also a voting procedure that can complicate the perfect matching of the BBQ dating turning points with the NBER one. Thus, some difference between the two procedures mat be due to such reasons. Apart of turning point of 2001 that is not captured because it does not exhibit two quarters of negative growth (Harding (2008)) and keeping in mind the above problems, our exercise reproduces the turning points from the NBER. Overall, the BBQ algorithm performs well on the US. After having run the above test, we apply the BBQ algorithm to the UK. We apply the BBQ algorithm to log-real GDP and we set k=2 and fix the duration of the business cycle to four quarters differently from the US exercise. This because Harding and Pagan (2002) find difficulties to identify for the UK the strong downturn of 1974 with a complete cycle of five quarter. The recession of 1974 was characterized by a complete cycle of 4. Hence, for the UK a duration of the complete cycle of four can be applied.

FIGURE 3B.2: UK Turning Points



Notes: The shaded area indicate the UK recession phases (1955:I-2009:IV) identified by the application of the BBQ algorithm on the log-real UK GDP (red line).

Figure 3B.2 shows the identified turnings points for the UK via the application of BBQ algorithm to the log-real GDP. From 1955 to 2009 we identify seven recessions. The number of UK recessions identified via the BBQ matches that ones reported from the Bank of England in the Inflation Report (Bank of England (2014)). The only exception concerns the recessions of 1970s that are treated as a single recession by the BoE. The turning points identified by the BBQ algorithm is going to be used as benchmark for studying the state-dependent tax multiplier.

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