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## **Nonlinear Effects of Macroeconomic Shocks**

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## **Preface**

This PhD. thesis was written in the period from July 2013 to April 2016 during my studies at the Department of Economics and Management of the University of Padova. There is a number of people who have contributed to the making of this thesis, to whom I am grateful.

First and foremost, I would like to express my gratitude to my thesis advisor Efre Castelnuovo for the continuous encouragement, patience, motivation and immense knowledge. His guidance helped me in all the time of research and writing this PhD. thesis. Besides my advisor, I would like to thank Giovanni Caggiano for his support and help. I have benefit greatly from his insightful comments and knowledge. I also thanks Timo Teräsvirta for his help, advices and guidance at the University of Aarhus.

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Juan Manuel Figueres , April 2016



# Contents

Preface	iii
Introduction	vii
Introduzione	ix
<b>Chapter 1</b>	
<i>News on State-Dependent Fiscal Multipliers: The Role of Confidence.</i>	1
<b>Chapter2</b>	
<i>Fiscal-Monetary policy Mix in Recessions and Expansions.</i>	41
<b>Chapter3</b>	
<i>Economic Policy Uncertainty Spillovers in Booms and Busts.</i>	75
with Giovanni Caggiano and Efrem Castelnuovo	



## Introduction

This thesis investigates the nonlinear macroeconomic effects of fiscal and uncertainty shocks. It comprises three contained chapters, each one of them being self-contained. In each chapter, theoretical predictions coming from theoretical models are presented and discussed. Such predictions are then tested using state-of-the-art econometric techniques.

The first chapter is titled “News in State-Dependent Fiscal Multipliers: The Role of Confidence”. This study scrutinizes the role of consumer confidence in determining the real effects that unanticipated (news) government spending shocks have on output in recessions and expansions by estimating a Smooth-Transition VAR model. To account for fiscal foresight, I employ a measure of anticipated fiscal shocks defined as the sums of expectations’ revisions over future fiscal spending. This variable is shown to carry relevant information about movements on government spending. My results indicate that fiscal multipliers during recession is both statistically larger than in expansions and greater than one. Importantly, consumer confidence is shown to play a decisive role on determining the effects of an anticipated spending shock within nonlinear framework. In particular, the response of confidence is key in explaining the statically larger fiscal multiplier during recessions. Moreover, the role of confidence is found to be relevant for the transmission of anticipated shocks only. These results qualify confidence as a key ingredient for understanding the transmission of fiscal news shocks (as opposed to unanticipated fiscal shocks).

The second chapter is titled “Fiscal-Monetary Policy Mix in Recessions and Expansions”. This study investigates the role of monetary policy in determining the size of the fiscal spending multiplier in recessions and expansions as for the U.S. economy. To this end, I quantify the size of state-dependent fiscal multipliers by using a nonlinear VAR model endowed with fiscal and monetary variables. I carefully separate anticipated and unexpected fiscal spending shocks by jointly modeling fiscal spending and the measure of spending news proposed by Ramey (2011 QJE). My results indicate that the fiscal multiplier in recessions is larger than one and statistically different from that corresponding to expansions. Importantly, the role of monetary policy during recessions triggers a crowding out effect. In particular, a counterfactual exercise clearly have the role played for the systematic policy to emerge. These findings highlight the importance of jointly consider monetary and fiscal indicators when studying the effects of a fiscal stimulus.

The third chapter titled “Economic Policy Uncertainty Spillovers in Booms and Busts” is joint paper with Giovanni Caggiano and Efrem Castelnuovo. This study aims at quantifying the impact of economic policy uncertainty shocks originating in the U.S. on the Canadian business cycle in booms and busts. It does so by employing a nonlinear Smooth-Transition VAR model to identify and simulate an increase in the U.S. economic policy uncertainty on a number of Canadian macroeconomics variables, including real activity indicators (industrial production and unemployment), inflation, a short-term interest rate, and the bilateral

exchange rate. Our results point to statistically and economically relevant nonlinear spillover effects. Uncertainty shocks originated in the U.S. explain about the 27% of the variance of the 2-years ahead forecast error of the Canadian unemployment rate in periods of slack vs. 8% during economic booms. Counterfactual simulations lead to the identification of a novel “economic policy uncertainty spillovers channel”. According to this channel, spikes in the U.S. economic policy uncertainty foster economic policy uncertainty in Canada in first place and, because of the latter, an increase in the Canadian rate of unemployment occurs.



## Introduzione

La tesi analizza gli effetti macroeconomici nonlineari di shock fiscali e di incertezza. Essa comprende tre capitoli, ciascuno dei quali è indipendente dagli altri. In ciascun capitolo, le predizioni teoriche derivanti da modelli macroeconomici vengono presentate e discusse. Tali predizioni sono poi testate empiricamente utilizzando tecniche econometriche all'avanguardia.

Il primo capitolo si intitola “News in State-Dependent Fiscal Multipliers: The Role of Confidence”. Questo studio analizza il ruolo giocato dalla fiducia dei consumatori nella determinazione degli effetti reali che shock di spesa pubblica non previsti hanno sul livello della produzione in recessione e in espansione utilizzando un modello vettoriale autoregressivo “Smooth-Transition”. Per tenere conto degli effetti di anticipazione sulla politica fiscale, utilizzo una misura di shock fiscali previsti, definita come la somma delle revisioni delle aspettative circa il livello futuro della spesa pubblica. Questa variabile risulta possedere rilevanti informazioni circa variazioni future effettive della spesa pubblica. I miei risultati indicano che il moltiplicatore fiscale durante le fasi recessive è statisticamente più elevato rispetto alle fasi espansive, oltre a essere maggiore di uno. In maniera importante, i risultati mostrano come il livello della fiducia dei consumatori giochi un ruolo decisivo nel determinare gli effetti di uno shock fiscale non previsto all’interno di un contesto non-lineare. In particolare, la risposta del livello di fiducia è cruciale nello spiegare la differenza statistica trovata in recessione. Inoltre, il ruolo del livello della fiducia è rilevante per la trasmissione soltanto degli shock previsti di politica fiscale. Questi risultati qualificano il livello di fiducia come un fattore determinante nel comprendere la trasmissione di shock fiscali previsti (a differenza degli shock fiscali non previsti).

Il secondo capitolo si intitola “Fiscal-Monetary Policy Mix in Recessions and Expansions”. Questo lavoro analizza il ruolo della politica monetaria nella determinazione della grandezza dei moltiplicatori fiscali in recessione e in espansione per l’economia degli Stati Uniti. A questo scopo, quantifico i moltiplicatori fiscali utilizzando un modello VAR non lineare che include variabili sia fiscali che monetarie. Per separare gli shock fiscali anticipati da quelli non anticipati, utilizzo sia variabili di spesa pubblica che la misura di “news” fiscale proposta da Ramey (2011 QJE). I miei risultati indicano che il moltiplicatore fiscale in recessione è maggiore di uno e statisticamente differente da quello che si ottiene in espansione. In maniera importante, il ruolo della politica monetaria in recessione comporta un effetto spiazzamento. In particolare, un esercizio controfattuale mostra in maniera chiara come emerga il ruolo giocato dalla politica monetaria. Questi risultati sottolineano l’importanza di considerare in maniera congiunta indicatori fiscali e monetari per analizzare gli effetti di politiche fiscali espansive.

Il terzo capitolo intitolato “Economic Policy Uncertainty Spillovers in Booms and Busts” è un lavoro congiunto con Giovanni Caggiano e Efrem Castelnuovo. Questo lavoro ha come

obiettivo la quantificazione dell'impatto di shock di incertezza politico-economica che hanno origine negli USA sull'andamento del ciclo economico canadese in recessione e in espansione. A tal fine, utilizziamo un modello vettoriale autoregressivo "Smooth-Transition" per identificare e analizzare gli effetti di un aumento del livello di incertezza economico-politica negli USA su una serie di variabili macroeconomiche canadesi, inclusi indicatori del livello dell'attività economica (produzione industriale e tasso di disoccupazione), tasso di inflazione, tasso di interesse a breve termine, e tasso di cambio bilaterale. I nostri risultati mostrano che ci sono effetti contagio non lineari rilevanti sia da un punto di vista statistico che economico. Gli shock di incertezza che hanno origine negli USA spiegano in recessione circa il 27% della varianza dell'errore di previsione a due anni del tasso di disoccupazione canadese, contro un valore pari a 8% in fasi di boom economico. Simulazioni controfattuali identificano un nuovo canale di contagio dell'incertezza economico-politica. In base a esso, aumenti del livello di incertezza economico-politica negli USA provocano in primo luogo un aumento del livello di incertezza in Canada e, per questo tramite, un aumento del tasso di disoccupazione canadese.

# **Chapter 1**

## **News on State-Dependent Fiscal Multipliers: The role of Confidence**



# News on State-Dependent Fiscal Multipliers: The role of Confidence\*

Juan Manuel Figueres<sup>†</sup>

## Abstract

This paper investigates the role of consumer confidence in determining the real effects that anticipated (news) government spending shocks have on output in recessions and in expansions as for the U.S. economy. To account for fiscal foresight, I employ a measure of anticipated fiscal shocks defined as the sums of expectations' revisions over future fiscal spending. This variable is shown to carry relevant information about movements on government spending. Results indicate that the fiscal multiplier during recession is both statistically larger than in expansions and greater than one. Importantly, consumer confidence is shown to play a decisive role in determining the real effects of an anticipated spending shock within a nonlinear framework. In particular, the response of confidence is key in explaining the statistically larger fiscal multiplier during recessions. Moreover, the role of confidence is found to be relevant for the transmission of anticipated shocks only. These results qualify confidence as a key ingredient for understanding the transmission of fiscal news shocks (as opposed to unanticipated fiscal shocks).

*Keywords:* Consumer confidence, Fiscal forecast, Fiscal spending multiplier, Nonlinear models, Smooth Transition Vector AutoRegressions.

*JEL Classification:* C32, E32, E62.

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## 1 Introduction

This paper quantifies the size of *state-dependent fiscal multipliers* to study the *role of consumer confidence* in determining the effects that an anticipated (*news*) government spending shock has on economic activity. In doing so I identify the fundamental fiscal shocks and I disentangle the effects that *anticipated* and *unanticipated* spending shocks have on confidence and output during recessions and expansions.

During the last years the debate about the role of consumers sentiment in determining the effectiveness of government policy has recovered impulse among economists and policy-makers. This idea relates to the Keynesian argument claiming that a fiscal stimulus boosts the economic activity during a recession through an improvement in confidence. In a recent paper Bachmann and Sims (2012) find empirical evidence indicating that consumers confidence is a critical factor in the transmission of spending shocks into the economic activity during a downturn. Importantly, they show that the main driver behind the relationship between a fiscal stimulus, confidence and the subsequent economic activity is the information regarding future improvements in fundamentals which follow spending shocks during recessions. Moreover, a fiscal issue that is also likely to critically affect the transmission of policy shocks is the anticipation effect, better known as fiscal foresight. This phenomenon arises from the fact that changes in fiscal policy are usually implemented with a lag so that agents might partially anticipate them and adjust their decisions before the policy changes take place. When studying consumers confidence, fiscal foresight implies that agents may anticipate a fiscal stimulus and update their expectations about the future fundamentals before the stimulus is actually implemented. Therefore, suggesting that “news” about a future fiscal stimulus may be more important in determining the role of confidence than the fiscal stimulus itself. The present paper is an attempt to shed some light on this last point by empirical studying the anticipation effect along with the role of confidence in determining the size of state-dependent fiscal multipliers.

I analyze the above mentioned relationship between confidence, fiscal multiplier and the anticipation effect in the framework of Structural Vector Autoregression (VAR) models. Given their considerable flexibility, these models have been widely used in literature on fiscal policy since the seminal contribution of Blanchard and Perotti (2002). Nevertheless, there are important issues to be considered when estimating fiscal multipliers by using VARs. First and foremost, in presence of fiscal foresight standard VAR models may not incorporate enough information to recover the fundamental fiscal shocks. This is because agents anticipate future changes (*news*) in the fiscal policy while the VAR econometrician can only observe the present and past values of fiscal variables. Forni and Gambetti (2010) and Ramey (2011) show that the government spending shocks estimated by using the standard fiscal variables are predicted by the government spending forecast, meaning that are at least partially anticipated (i.e., are non-fundamental). Importantly, Leeper, Walker, and Yang (2013) prove

that when the econometric analysis fails to address fiscal foresight, the estimated tax multiplier may exhibit quantitative important biases. Secondly, estimating the effects of fiscal policy by using linear VARs omits the possibility that the fiscal multiplier may vary across the business cycle as it is mentioned by the traditional Neo-Keynesian literature and New Keynesian models in presence of the zero lower bound.<sup>1</sup> Recent empirical studies have considered the possibility of government spending shocks having different effects depending on the state of economy. Among others, Auerbach and Gorodnichenko (2012), Bachmann and Sims (2012), Mittnik and Semmler (2012), and Baum, Poplawski-Ribeiro, and Weber (2012) find the fiscal multiplier to be significantly larger during recessions. Moreover, in a recent paper Caggiano *et al.* (2015) estimate state-dependent fiscal multipliers by explicitly addressing the fiscal foresight. To overcome the issue of non-fundamentalness they employ a measure of anticipated fiscal shocks proposed by Forni and Gambetti (2014). Their results indicate that the fiscal multiplier is statistically larger during periods of deep recession.

In the spirit of Auerbach and Gorodnichenko (2012), I compute state-dependent fiscal multipliers by employing a Smooth-Transition VAR model which allows me to consistently estimate the responses to a spending shock in recessions and in expansions. Moreover, following Forni and Gambetti (2014), I account for the fiscal foresight effect by implementing a measure of *anticipated* spending shocks that conveys relevant information about future movements (*news*) in government spending. This measure is defined as the sums of expectations' revisions about the growth rate of future government spending from the Survey of Professional Forecasters. As is shown in the present paper this News variable has a superior predictive power in comparison to others measures used in the literature. Finally, to isolate the role of confidence I compute the fiscal multipliers for the counterfactual situation where the level of confidence remains constant (i.e., it does not react to spending shocks).

My main results are the following. First, for an *anticipated* (news) spending shock the fiscal multiplier is statistically larger during recessions than over expansions. Moreover, the fiscal multiplier over recession is statistically larger than one. Second, a counterfactual exercise which holds the level of confidence constant gives as result fiscal multipliers that are not anymore statistically different across regimes. This points to the role confidence as key in determining the real effects that an anticipated spending shock has on output within

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<sup>1</sup> For example in the IS-LM-AD-AS the size of the fiscal multiplier exhibits large values during periods of economic slack (the AS curve is flat and there is a lower crowding out effect affecting investment and consumption) and small values in economic booms (the AS curve is steep, implying a higher crowding out effect). Moreover, Eggertsson (2009), Christiano, Eichenbaum, and Rebelo (2011) and Woodford (2011) show that when the nominal interest rate is held at the ZLB, a deficit financed increase in government spending leads to an increase in inflation expectations, which in turn leads to a decrease in real interest rates, boosting in this way investment and consumption. In such cases without crowding out effect the fiscal multiplier is around 3.

nonlinear framework. Third, for an *unanticipated* spending shock (i.e., an innovation in the fiscal variable) the multipliers are never statistically larger than one. Interestingly, in this case confidence does not turn out to be important in explaining non-linearity. These findings suggest that the reason behind the role of confidence is the information about future government spending provided by the anticipated (news) spending shocks and not contained in the fiscal variable itself.

The closest papers to mine are Bachmann and Sims (2012), Ramey and Zubairy (2014) and Caggiano *et al.* (2015). Bachmann and Sims (2012) show that consumers confidence is a key factor in the transmission of spending shocks into the economy activity during recessions. With respect to them, I study the role of confidence in determining the anticipated and unanticipated effects of a government spending shock. In contrast they focus only on the unanticipated effect of a fiscal shock. Importantly, I show that when disentangling the anticipated and unanticipated effects of spending shocks, confidence is found to be a relevant ingredient for the transmission of anticipated (news) government spending shocks only. This indicates that the news about future increases in government spending are critical in determining the relationship between the consumers confidence and the subsequent economic activity when adopting a spending-based fiscal stimulus. Ramey and Zubairy (2014) and Caggiano *et al.* (2015) study the non-linearity of fiscal multipliers by accounting for fiscal foresight. While the former find no evidence in favour of state-dependent fiscal multipliers, Caggiano *et al.* (2015) show that the fiscal multiplier is statistically larger only during severe economic conditions. My contribution complements these two papers by adding consumer confidence to the vector of modeled variables and considering the role that confidence plays in the transmission of anticipated fiscal shocks in good and bad times.

The rest of this paper is organized as follows. Section 2 studies the anticipation effect and the estimation of anticipated (news) spending shocks. Section 3 offers statistical evidence in favour of non-linearity and presents the Smooth-Transition VAR model along with the data used for its estimation. Section 4 describes the results. The last section concludes.

## **2 The Fiscal Foresight**

Fiscal Foresight arises because of the fact that changes in fiscal policy are usually implemented with a lag so that agents might partially anticipate them by early reacting to a change in spending and taxes (i.e., reacting before its implementation). When agents base their decisions on a larger information set than the econometrician has, the use of structural VAR models to recover the effects of changes in fiscal policy is likely to lead to non-fundamentalness problem (Beaudry and Portier, 2014). This means that the Vector Moving Average (VMA) representation of SVARs is not invertible in the past. Hence, present and past values of the fiscal variables would not convey enough information to recover the fiscal shocks. As Leeper, Walker, and Yang (2013) show, when agents' information set is larger



than the one of the econometrician, then agents and econometrician employ different discounting patterns. That is, while the econometrician discounts in the usual way and assign a larger weight to recent shocks, the private agents discount by assigning a smaller weight to more recent realizations of the shock. This is because, with fiscal foresight, the recent shocks are related with news informing about movements in the more distant future.

According to different empirical studies the government spending shocks estimated by using the standard fiscal variables are Granger-caused by the government spending forecast, i.e., the estimated shocks are non-fundamental because of the fiscal foresight (Forni and Gambetti, 2010 and Ramey, 2011). Therefore to properly assess the effects of the fiscal policy over the business cycle we have to first overcome the non-fundamentality problem. This issue may be solved by enlarging the information set used to estimate the spending shocks. Different approaches are proposed in the literature in order to do so. Ramey and Shapiro (1998) use a narrative approach to identify government spending shocks, they use the *Business Week* magazine to construct a dummy variable reflecting the major military episodes which anticipate an increase in the defense spending. Ramey (2011) employs additional sources of information plus the *Business Week*, she proposes the use of a variable measuring the expected discounted value of government expending changes resulting from foreign political events. Leeper, Richter and Walker (2012) implement a calibrated DSGE model and government spending forecast from the Survey of Professionals Forecasters to account for the fiscal foresight. Forni and Gambetti (2010) adopt a structural, large dimensional, dynamic factor model in order to enlarge the information set used in the estimation of the government spending shocks.

In the present study I use the approach developed by Forni and Gambetti (2014). They propose the use of VAR models endowed with an supplementary variable, the “government spending news”, containing additional information about future government spending that accounts for the fiscal foresight, hence solving a fiscal issue with the right fiscal data. This variable is defined as difference between the expectation of the agents about the growth rate of government spending for  $t + j$  at time  $t$  and the expectation at time  $t - 1$ , that is  $news_t = E_t g_{t+j} - E_{t-1} g_{t+j}$ . This is the expectation revision representing the new information that becomes available at time  $t$  proportional to the anticipated shock not contained in the fiscal variable. Thus when a government spending shock occurs at time  $t$ , even if the government spending measure remains unchanged due to the implementation lag, the agents know that government spending will change in the future so that they react by updating their expectations. In order for this variable to convey the information needed to recover the anticipated shock, it is necessary to consider the expectation revision of the spending growth rate over a horizon equivalent to the  $h$  periods of foresight (i.e.,  $j$  equal to the  $h$  periods ahead for which the agents anticipate the fiscal movements).<sup>2</sup> But in general

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<sup>2</sup> Perotti (2011) concludes that the expectation revision  $E_t g_t - E_{t-1} g_t$  conveys little information on future

the periods of foresight are unknown. Nevertheless, as proposed by Forni and Gambetti (2014), this problem can be overcome by using the *sum of expectations' revisions* for a horizon long enough to ensure that the revision variable is proportional to the anticipated fiscal shock. Therefore the “government spending news” variable is defined as follows:

$$news_t(j, J) = \sum_{j=1}^J (E_t g_{t+j} - E_{t-1} g_{t+j}) \quad (1)$$

where  $E_t g_{t+j}$  represents agents' expectations at time  $t$  for the growth rate of government spending from quarter  $t + j - 1$  to quarter  $t + j$ , and  $E_{t-1} g_{t+j}$  represents agents' expectations for the same variable and period at time  $t - 1$ . Consequently,  $E_t g_{t+j} - E_{t-1} g_{t+j}$  represents the new information that becomes available to the agents at time  $t$  about the growth rate of government spending  $j$  quarters ahead. When  $J$  is large enough (i.e.,  $J \geq h$ .) the variable  $news_t$  is proportional to the *anticipated* government spending shock. The expectations' revisions are constructed by using the forecast for the growth rate of government spending from the Survey of Professionals Forecasters.<sup>3</sup> This survey contains the forecasts of the annualized growth rates of government spending up to four quarters ahead starting from 1981:Q3. Caggiano *et al.* (2015) employs the above defined measure of government spending news in order to address the fiscal foresight effect when quantifying the size of the state-dependent fiscal multipliers. They find the News variable to convey significant information about future movements in government spending.

Next I perform a test in order to assess the predictive power of the different specifications of the  $news_t(j, J)$ , and then I analyze the main advantages of using the expectations' revisions approach to identify anticipated government spending shocks.

**The predictive power of the News variable.** In order to statistically test the information content of the News variable I perform a Granger-causality test between the News variable computed for different specifications of  $j$  and  $J$ , and the VAR estimated government spending shocks. The aim of this test is twofold, first is to prove that the shocks estimated with standard variables can be predicted by the expectations' revisions (i.e., are non-fundamental shocks), and the second is to assess the proper specification of News variable that maximize its predictive power. Notes that when analyzing the different specifications of the News variable one should take into account all the variables included in the system under study. This is because the informational power about the movements in government spending of each specification for  $news_t(j, J)$  depends upon the economic system in which the News variable is embedded. Therefore, to be consistent with the

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government spending. The reason is that such expectation revision does not cover a horizon long enough to capture the anticipation effect, i.e.,  $j < h$ .

<sup>3</sup> As Perotti (2011) points out, constructing measures of expectations of government spending by using the forecast of the growth rate instead the forecast of the levels helps to avoid inconsistencies resulting from the frequent changes in the base years affecting the SPF forecast of the variable in levels.

variables used in the main analysis of this paper, the spending shocks are drawn from a linear VAR(4) endowed with the log of real per capita government spending, the confidence index and the log of real per capita output. Moreover, given that the SPF collects the forecast for the growth rate of government spending up to four quarters ahead, the largest horizon for  $news_t(j,J)$  is  $J = 3$ .

Table 1 shows the p-values for the Granger-causality test of the one period-lagged News variable. The top panel contains the expectation revisions and the bottom panel the *sum of expectations'* revisions. Observe that only  $news(1,1)$  and  $news(2,2)$  turn out to be informative about the government spending shocks, while the expectation revision for the shortest and the longest horizon,  $news(0,0)$  and  $news(3,3)$  respectively, have not predictive power since the null hypothesis is always accepted.<sup>4</sup> Consequently, when examining the sum of expectations' revisions the specification  $news(1,2)$  ( i.e.,  $news(1,1) + news(2,2)$  ) results to be the most informative one.

Figure 1 plots  $news(1,2)$  for the sample 1981:Q4-2013:Q1. We can observe that the series exhibits spikes related with exogenous fiscal policy episodes. For example the positive spikes coincide with episodes related to significant increase in government spending as the beginning of the War in Afghanistan (2001:Q4) and 2009 Fiscal Stimulus package (2009:Q1). While the negative spike at 1989:Q4 coincides with the government spending cut resulting from the end of the Cold War associated with the fall of the Berlin Wall.

**Comparison with Ramey's narrative approach.** Another widely used measure to overcome the fiscal foresight effect is the variable developed by Ramey (2011). This variable estimates the expected present value of government expending changes due to foreign political events, being constructed by using the *Business Week* magazine (mainly) and additional newspaper sources. Below I show that the News measure conveys information to predict the Ramey's variable. To do so I run a bivariate VAR with Ramey's and the News variable  $news(1,2)$  regressing both variables on their first lags.<sup>5</sup> Table 2 reports the p-values of the t-test corresponding to the exclusion of the specified variable. I employ the longest possible sample of 1981:Q4-2013:Q1. Moreover, given that the first twenty observations of this sample are all zero for Ramey's variable, I also use a shorter sample starting from 1986:Q4. Note that only the null hypotheses for the News's variable coefficient explaining the Ramey's variable are rejected, meaning that the News variable Granger causes the Ramey's variable while the reverse direction of causality is rejected. Furthermore Figure 2 shows the News variable together with Ramey's variable. Observe that the largest spikes in  $news(1,2)$  tend to anticipate the changes of the Ramey's variable, being this behavior in line with the Granger-causality test.

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<sup>4</sup> As Forni and Gambetti (2014) point out, when the expectation revision  $E_t g_{t+j} - E_{t-1} g_{t+j}$  spans over a horizon  $j$  too small the revision does not provide information about the government spending shocks.

<sup>5</sup> As Ramey (2011) does, the Ramey's variable at time  $t$  is divided by the nominal GDP of the previous period.

### 3 Methodology and Data

#### 3.1 Nonlinear Model

With the purpose of study the role of confidence in determining the effects of fiscal shocks during recessions and expansions I implement a Smooth-Transition VAR model (for a detailed presentation, see Teräsvirta, Tjøstheim and Granger, 2010). The most important advantage of this model is that it allows for responses differentiated across states of the economy (i.e., recession and expansion) having an smooth transition from one state to another. The model is described below:

$$\mathbf{X}_t = F(z_{t-1})\mathbf{\Pi}_R(L)\mathbf{X}_t + [1 - F(z_{t-1})]\mathbf{\Pi}_E(L)\mathbf{X}_t + \boldsymbol{\varepsilon}_t, \quad (2)$$

$$\boldsymbol{\varepsilon}_t \sim N(0, \boldsymbol{\Omega}_t), \quad (3)$$

$$\boldsymbol{\Omega}_t = F(z_{t-1})\boldsymbol{\Omega}_R + [1 - F(z_{t-1})] \boldsymbol{\Omega}_E, \quad (4)$$

$$F(z_t) = \exp(-\gamma z_t) / [1 + \exp(-\gamma z_t)], \gamma > 0, z_t \sim N(0, 1). \quad (5)$$

where  $\mathbf{X}_t$  indicates the vector of endogenous variables,  $\mathbf{\Pi}_R(L)$  and  $\mathbf{\Pi}_E(L)$  are the matrices of coefficients accounting for the dynamic of the variables in  $\mathbf{X}_t$  and  $\boldsymbol{\varepsilon}_t$  indicates the vector of residuals from the reduce form, with zero mean and state-depended variance-covariance matrix  $\boldsymbol{\Omega}_t$ . Moreover  $\boldsymbol{\Omega}_R$  and  $\boldsymbol{\Omega}_E$  are the reduced-form residuals variance-covariance matrices during recession and expansion. Notice that the above presented model accounts for nonlinearities coming from the dynamics of the system as well as from the contemporaneous relationships. Finally, one of the most important feature of this model is the transition function  $F(z_t)$ . This function indicates the probability of being in a recession, where  $z_t$  is the switching variable represented by an index of the business cycle and  $\gamma$  is the smoothness parameter regulating the transition from a regime to another.<sup>6</sup> In order  $\gamma$  to be scale invariant the index  $z_t$  is normalized to have unit variance and zero mean. Note that if  $\mathbf{\Pi}_R(L) = \mathbf{\Pi}_E(L)$ ,  $\boldsymbol{\Omega}_R = \boldsymbol{\Omega}_E$ , the model falls back to the linear framework.

In addition the index  $z_t$  is dated at  $t - 1$  to avoid the contemporaneous feedbacks resulting from policy actions taken whenever the economy is in an expansion or a recession. In line with Auerbach and Gorodnichenko (2012), Bachmann and Sims (2012), Caggiano, Castelnuovo, and Groshenny (2014), and Berger and Vavra (2014) I define the switching variable  $z_t$  as the standardized seven-quarter moving average of output growth rate. The parameter  $\gamma$  is calibrated to 1.88 to ensure that the economy will be in recession regime about 15 percent of the times, a frequency in line with the NBER business cycle dates for the sample 1981:Q4-2013:Q1. Therefore the economy is defined to be in a recession

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<sup>6</sup> Lower values of parameter  $\gamma$  will insure smoother switches from one regime to another.

when  $F(z_t) > 0.85$  in such a way that with  $\gamma = 1.88$  the probability to be in recession is  $Pr(F(z_t) > 0.85) \approx 15\%$ . This calibration implies  $z_t \leq -0.92\%$  during the recessionary regime. Figure 3 contrasts the transition function  $F(z_t)$  with the recessions dated by the NBER.

The variable  $z_t$  is assumed to be exogenous to the system, hence is not included in the vector of endogenous variables  $\mathbf{X}_t$  so that there is no feedback from the exogenous variable to the dynamic of the system (i.e., the system can remain for a long time in deep recessions or in strong expansions, being the model linear in each fixed regime). The advantage of this last assumption is that the estimated impulse responses are linear and do not depend either on the initial conditions, the sign of the shock or the size of the shock (Koop, Pesaran and Potter 1996). Nevertheless, as it was pointed out in Owyang, Ramey and Zubairy (2013) this method of computing the impulse responses has two main drawbacks. First, in reality the economy is hardly to remain either in a deep recession or in a strong expansion for long terms of time. Secondly, even if the economy starts in one of the regimes, a shock affecting  $Y_{t+h}$  would indirectly affects  $z_{t+h-1}$  too, and, thereby the future state of the economy (i.e., the responses of output affects the future regimes which in turn affects the dynamic of the futures responses). Then even if I compute the responses for an horizon of 20 quarters, in order to overcome the issues above described I focus my attention on the responses during the first 5 quarters, being this horizon consistent with the average duration of a recession for the sample used.<sup>7</sup> My focus in the short run responses to a fiscal shock renders much unproblematic the use of conditionally linear impulse responses.

The baseline specification of the vector of endogenous variables is given by  $\mathbf{X}_t = [g_t \text{ conf}_t y_t \text{ news}_t]'$ , where  $g_t$  is the log of real per capita government spending,  $\text{conf}_t$  is the confidence measure,  $y_t$  is the log of real per capita output and  $\text{news}_t$  is the government spending News variable.

**Model Estimation.** Because of the high non-linearity of the model (2)-(5), I estimate it by using Monte Carlo Markov Chain algorithm developed by Chernozhukov and Hong (2003). Since nonlinear estimation becomes problematic when too many parameters are being estimated, I employ a parsimonious specification of the STVAR model that includes two lags. Moreover, in order to construct the confidence bands I use bootstrap procedure to obtain the distribution of the generated impulse responses. See Appendix A.

**Testing Non-linearity.** In order to assess the presence of non-linearity at a multivariate level, I carry out two tests for the baseline vector of endogenous variables  $\mathbf{X}_t$ . First, following Teräsvirta and Yang (2014), I test the null hypothesis of linearity for the dynamics of the system in (2) against the alternative of (Logistic Vector) STVAR with a single switching variable. The result of the test points out to a clear rejection of the null hypothesis in favour

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<sup>7</sup> During the period 1981:Q4-2013:Q1 the NBER dates 4 recessions: 1981:Q4-1982:Q4; 1990:Q3-1991:Q1; 2001:Q1-2001:Q4, and 2007:Q4-2009:Q2.

of the STVAR specification. See Appendix B. Secondly, I test the constancy of the error covariance matrix in (4) against the alternative of Smooth-Transition via the test proposed by Yang (2014). For this last test, the null hypothesis of constant covariance matrix is rejected in favour of the Smooth-Transition specification. See Appendix C.

### 3.2 Data

The sample period used in the estimation is 1981:Q4-2013:Q1, being 1981:Q4 the first observation available for the News variable.<sup>8</sup> Note that this sample does not include the large variation of government spending associated with the Second World War and the Korean War. Nevertheless, as it was pointed out in Blanchard and Perotti (2002) and Christiano (2013) these two war episodes had very special characteristics and effects in the economy,<sup>9</sup> making difficult to think of them as generated by the same stochastic process related with the rest of spending variations observed in the sample. Therefore using a shorter sample helps to avoid inconsistent estimation of the fiscal multiplier.

In line with Auerbach and Gorodnichenko (2013) government spending is the real government (federal, state and local) purchases (consumption and investments), and output is the real gross domestic product (GDP) measured in chained 2000 dollars.<sup>10</sup> These variables are expressed in per capita terms by dividing by the civilian non-institutionalized population age 16 and over. As suggested by Bachmann and Sims (2012), the measure of confidence is the Index of Consumers Expectations from the Michigan Survey of Consumers. This index represents an average of three different forward-looking survey questions related with the expectations about the business and personal financial conditions.<sup>11</sup> Basically, higher values of the index involves more confidence. By comparing the index series with the recessions dated by NBER it is easy to note that the Consumers Expectations has a procyclical behavior, exhibiting the lowest values in coincidence with the recession dates (see Figure 4). Moreover the government spending News variable (*news*) is constructed according to the equation (1) as proposed by Forni and Gambetti (2014). The variables  $g_t$  and  $y_t$  are taken in log levels due to possible cointegration relationships. Consequently, the variable  $news_t$  is expressed in

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<sup>8</sup> The Survey of Professional Forecasters provides forecast for the growth rate of government spending since 1981:Q3. Given that to construct the News variable a time  $t$  we need the forecast made at  $t - 1$ , the first observation of the constructed series for News variable is at 1981:Q4.

<sup>9</sup> For example, main durables goods were rationed during the Second World War, something that constrained the government spending from increasing further. Moreover during the Korean War taxes were significantly raised in order to finance the increase in the military spending.

<sup>10</sup> The series for government purchases are drawn from the table 3.1 of the Bureau of Economic Analysis and calculated as the sum of consumption expenditures and gross investments, minus the consumption of fixed capital. The series are converted in real terms by using the GDP deflator. The series for real GDP and its implicit deflator are obtained from the Federal Reserve Bank of St. Luis website.

<sup>11</sup> For details about the computation of the Index of Consumers Expectations see Appendix D.

cumulative sums to preserve the same order of integration. Moreover, it has to anticipate spending levels, and recall that the news is expressed in growth rates.

### 3.3 The Predictive Power of the News Variable within a Nonlinear Framework

In order to statistically test the informative power of the News variable within a nonlinear framework I perform a Granger-causality test involving the News variable computed for different specifications of  $j$  and  $J$ , and the fiscal shocks estimated with a STVAR not modeling News. First, I estimate the fiscal spending shocks by employing the Smooth-Transition VAR model (2)-(5) endowed only with the log of real per capita government spending, the confidence index and the log of real per capita output. Then I test whether or not these shocks can be predicted by the News variable. Table 3 contains the p-values for the Granger-causality test of the one period-lagged News variable. Observe that, alike Section 2, the specification  $news(1,2)$  is the most informative about government spending shocks. Therefore from now on I define the News variable as  $news(1,2)$ .

### 3.4 Identification of the Government Spending Shock

Following Forni and Gambetti (2014), I estimate the anticipated government spending shocks by including the government spending News variable in the vector of endogenous variables  $X_t$ . It is important to note that the forecasts used to construct the News variable are likely to be driven by non-fiscal shocks as well. Therefore, as Forni and Gambetti (2014) indicate, a proper identification scheme would be to order the news measure as the last variable of the Cholesky decomposition. Ordering the news last allows me to control for shocks others than the fiscal news ones which may affect the forecast revisions.<sup>12</sup> Hence, an *anticipated* government spending shock is defined as an innovation in the News variable. Differently, an *unanticipated* government spending shock is defined as an innovation in the fiscal variable itself. This identification strategy allows me to disentangle the effects that *anticipated* and *unanticipated* spending shocks have on confidence and output.

## 4 Results

This section presents the main results of the paper. For all the estimations I present the reaction of the system to a government spending shock and the respective fiscal multiplier. I compute the fiscal multiplier in two different ways. First, I compute the max multiplier as the maximum response of output divided by the maximum response of government spending.<sup>13</sup>

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<sup>12</sup> This specification implies that, by construction, innovations in  $news_t$  have no impact effect on the rest of the endogenous variables.

<sup>13</sup> A similar measure is proposed by Blanchard and Perotti (2002). Differently, they use the ratio of the

Second, I calculate the sum multiplier defined as the ratio of the sum of output response (to a spending shock) to the sum of government spending response (to a spending shock). This latter measure is proposed by Woodford (2011) and widely used in the literature given that it takes into account the persistence of the fiscal shock. Both types of multiplier are computed for the short run horizon of 5 quarters (length of time consistent with the NBER recessions), and the short-medium run horizons of 8 and 16 quarters.<sup>14</sup> Moreover given that the variables enter in the system in logs, the estimated multipliers are scaled by the sample average of  $Y/G$  in order to transform percent changes into dollars changes<sup>15</sup>. Section 4.1 shows the estimates of the system (2)-(5) for the baseline specification of  $X_t$  with an *anticipated* (news) spending shock. Additionally, for reasons of comparison I also present the estimates of the linear model. In Section 4.2 I study the role of confidence by computing the counterfactual multipliers conditional to a fixed level of confidence. Finally, in Section 4.3 I compare the previous results against the reaction of the system to an *unanticipated* government spending shock.

#### 4.1 Anticipated (News) Spending Shocks

This section presents the estimates for the baseline  $X_t$  containing the log of real per capita government spending ( $g_t$ ), the index of confidence ( $conf_t$ ), the log of real per capita GDP ( $y_t$ ) and the News variable ( $news_t$ ) with an *anticipated* (news) government spending shock defined as the last shock of the Cholesky scheme. Figure 5 compares the impulse responses of the system for the Smooth-Transition VAR model over recessions and expansions with those for the linear model. As we can see in the linear framework output has a small positive reaction in the short-medium run which becomes negative after 14 quarters. Confidence exhibits a behavior similar to output, having a positive reaction over the first quarters which is reverted and becomes negative after 7 quarters. Moreover the reaction of government spending is smooth and positive reaching its maximum at 11 quarters before starting to decrease. This responses would indicate a modest effect of a spending shock on output, nevertheless when accounting for nonlinearities the responses of

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maximum response of output to the *impact* response of government spending rather than to the maximum response of government spending. As Ramey and Zubairy (2014) point out, this kind of multipliers is not informative for the policy makers given that it does not consider the evolution of the cost of government spending associated with the path of output.

<sup>14</sup> These horizon lengths are proposed by Ramey and Zubairy (2014), given that the multipliers for two-year and four-year horizons are the most relevant for the short-run stimulus policy.

<sup>15</sup> This *ex post* conversion factor has been criticized by Ramey and Zubairy (2014) who argue that the  $Y/G$  ratio for the U.S. data sample 1889-2013 varies from 2 to 24 with a mean of 8. Therefore the use of a constant value for  $Y/G$  may lead to inflated, or at least distorted, multiplier estimates. In the sample used in this paper the  $Y/G$  ratio varies from 5.39 to 6.76 with a mean of 5.99 and a variance of 0.13. Hence, given its small variation, the adoption of a constant value for  $Y/G$  does not seem to be problematic in my case.



the system become markedly different depending on the state of the economy. Observe that for the nonlinear model, at the short-medium run, the reaction of output during recessions is statistically larger than over expansions. In recessions output significantly increases over the first 5 quarters and then decreases with some persistence, in contrast during the expansion regime output has a small positive reaction at the short run that is never statistically different from zero. Similar to output, confidence has a strong and positive reaction during recessions which is rapidly reverted after 4 quarters, while over expansions its reaction is slightly negative and statistically different from zero only at the long run. In addition the reaction of government spending is positive for both regimes and larger during recessions. Table 3 contains the estimated fiscal multipliers for the baseline specification  $X_t$ . The multiplier during recessions is much larger than over expansions across the three different horizons of 5, 8, and 16 quarters, exhibiting its maximum values of 3.41 (max) and 3.70 (sum) at the short run, and being also statistically larger than one. While in expansions the sum (max) multiplier is never larger than 0.39 (0.67).<sup>16</sup> The multipliers (max and sum) corresponding to the linear VAR are always lower than the ones related with recessions but larger than those corresponding to expansions, thus suggesting that the linear model captures the average effect of an increase in government spending between the two different states of the economy.

Furthermore even if at first sight the above results suggest the existence of nonlinearities, it is not clear whether or not the multiplier is statistically different across regimes. Therefore in order to address this last point I run a test by computing the distribution of the difference between the multiplier estimated during recessions and that estimated over expansions.<sup>17</sup> The aim of this exercise is to test if the difference in multipliers between regimes is statistically different from zero. Given that my focus is on the short run, I present the results of the test for the horizon of 5 quarters. This length of time is consistent with the average duration of a recession in the data. Nevertheless, the results here presented are robust to the different horizons of 8 and 16 quarters (results not shown here, but available upon request). The top levels of Figure 6 depicts the distribution of the difference for the max and sum multipliers with 68 % confidence intervals. Note that in both cases the zero line lies outside the confidence intervals, therefore providing evidence in favour of state-dependent multipliers from the statistical standpoint. Moreover given the importance that controlling for taxes may have in measuring the effects of a government spending shock, like when there is a fiscal consolidation or a stimulus package, I perform a further check (not shown here) by

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<sup>16</sup> One should read the reported values as upper bounds for extremes states of the economy due to the assumption that the economy remaining in a recession/expansion forever.

<sup>17</sup> The empirical density of the difference between multipliers is obtained by subtracting a realization of the multiplier in expansions from a realization of the multiplier in recessions for a number of times equal to 5,000. Moreover, each realization of the multiplier is obtained via bootstrap procedure.

enlarging the estimated system with a measure of taxes.<sup>18</sup> I found that the baseline results are robust to this specification containing taxes.

## 4.2 Does Confidence Matter?

What is the role of confidence within nonlinear framework? Does it matter for the real effects of anticipated (news) fiscal shocks? The role of consumers confidence on the business cycle has been widely discussed in the literature since Keynes featured the concept of animal spirits. This concept relies on the idea that changes in agents' sentiment about economic activity account for important fluctuations in aggregate consumption, which in turn account for large fluctuations in output. Observe from Figure 5 that confidence and output positively and largely reacts to an unanticipated government spending shock during recessions while during expansion the reaction of both variables is negligible, thus suggesting a possible connection between both reactions. Therefore having in mind the idea of animal spirits, the answers to the above questions are key when implementing fiscal stimulus. Then, to address this point I perform a counterfactual exercise by computing the multipliers for the system  $X_t$  conditional to a fixed level of confidence (i.e., the confidence response to an increase in government spending (news) is offset by another shock such that the level of confidence remains unchanged). Following the approach adopted by Sims and Zha (2006) I generate a hypothetical sequence of confidence shocks in order to held the response of confidence fixed to zero at each horizon, in such a way that the output response reflects the effect of an anticipated (news) government spending shock in a hypothetical situation where confidence is held constant.<sup>19</sup> The last rows of each panel in Table 3 shows the counterfactual fiscal

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<sup>18</sup> The series for taxes are drawn from the table 3.1 of the Bureau of Economic Analysis and constructed by subtracting from the current receipts the social benefits. The nominal series are converted in real terms by using the GDP deflator. Moreover the variable is expressed in per capita terms by dividing by the civilian non-institutionalized population age 16 and over, and then taken in logs levels. Taxes is ordered second in the Cholesky decomposition, after government spending and before confidence. The results are documented in an Appendix available upon request.

<sup>19</sup> Sims and Zha (2006) study the role of endogenous monetary policy in the transmission other shocks. They combine an initial shock with a hypothetical sequence of policy innovations enough to offset the endogenous policy response at each horizon. A drawback of using this approach is that ignore the Lucas critique by assuming that the agents are repeatedly surprised by the hypothetical policy shocks without adapting their forecast process of the economy to the new policy. Nevertheless, as Sims and Zha point out, this is an acceptable assumption to entertain. This is because it would take some time for the agents to learn that policy will not respond, since it is illogical to assume that they will immediately and fully understand the policy change and take it as permanent. Therefore this kind of approach is more suitable for a short run analysis like mine, given that it is reasonable to assume that the agents will be surprised by, in my case, confidence shocks for 5 quarters, while the same would not be true for 20 quarters. A more detailed explanation about how to compute the hypothetical shocks is done by Bachmann and Sims (2012).

multipliers (max and sum) when level of confidence remains fixed.<sup>20</sup> Note that, at the short run, during recessions the counterfactual multipliers (max and sum) are significantly lower than the baseline multipliers, while over expansions the constrained multipliers show a modest variation with respect to its unconstrained counterpart. As a consequence, the difference in multipliers between regimes shrinks. Then, as in the previous section, I test whether or not the counterfactual multipliers are statistically different across regimes. From the bottom levels of Figure 6 it is easy to observe that now the difference in multipliers between recessions and expansions is not statistically different from zero, suggesting that when confidence is held constant multipliers do not depend on the state of the economy.<sup>21</sup> These findings indicate that confidence plays a critical role in determining the real effects of anticipated spending shocks within nonlinear framework, in such a way that the confidence response is key in explaining the statistically different fiscal multipliers. A possible explanation to this might be given by the fact that during recessions the level of confidence is lower than usual (see Figure 4), hence an anticipated (news) government spending shock generates a boost in confidence, which in turn stimulates output. While during expansions an innovation in government spending does not further increase confidence which is already at normal levels, thus having a modest effect on output. Importantly, my results are robust to the different horizons of 8 and 16 quarters (figures not shown here, but available upon request). These findings are in line with those from Bachmann and Sims (2012). In addition to them, my analysis account for the fiscal foresight effect, what allows me to properly identify the fundamental fiscal shocks differentiating the *anticipated* from the *unanticipated* fiscal effects.<sup>22</sup> This identification approach also permits me to detect which of the two effects is the relevant one for explaining the role of confidence in determining state-dependent spending multipliers. So far, results suggest that the anticipated (news) fiscal effect would be the main driven force behind the role of confidence.

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<sup>20</sup> Given that I focus my attention in nonlinearities I only present the counterfactual multipliers for the Smooth-Transition VAR model.

<sup>21</sup> This results are also robust to the specification controlling for taxes (figures available upon request).

<sup>22</sup> Bachmann and Sims (2012) perform a robustness check to control for the fiscal foresight by estimating a nonlinear VAR endowed with the Ramey's variable for the sample 1960:Q1-2011:Q1. Nevertheless there exist two important objections to their exercise. First, they estimate the fiscal multipliers for an *unanticipated* government spending shock defined as an innovation in the government spending variable. As Ramey (2011) indicates, this procedure is not valid given that her *News* variable does not fully capture all the anticipated changes in government spending, it only considers changes related with military events. Therefore if one realize an exercise as Bachmann and Sims do, the estimated spending shocks will include *anticipated* changes in government spending that are not captured by the Ramey's variable (not military related), i.e., the shocks are non-fundamental. Secondly, Ramey (2011) shows that her variable has a low predictive power about government spending in a sample that excludes the WWII and the Korean War, what worsen the non-fundamentalness problem in Bachmann and Sims exercise.

### 4.3 Unanticipated Government Spending Shocks

Which is the effect of an *unanticipated* spending shock? Disentangling the effects of unanticipated and anticipated fiscal shocks may be key in explaining why confidence matters for the state-dependent fiscal multipliers. Aiming to do so, I compute the IRFs and multipliers for an unanticipated government spending shock defined as the first of the Cholesky decomposition for the baseline specification  $\mathbf{X}_t$ . Then I compare the results from this section with the case of anticipated (news) spending shocks. Figure 7 depicts the impulse responses of the system. Note that in contrast with the previous sections, during recessions output immediately reacts on impact and remains almost constant for a few quarters to then significantly fall. Government spending itself behaves similar to output during recessions, strongly increasing at the very short horizon to then start to fall. Observe that the above listed differences are more marked in the linear model. While the responses of output and government spending differ from the anticipated fiscal shock, the shape of the confidence reaction does not exhibit important alterations. Table 4 contains the estimated fiscal multipliers for the unanticipated government spending shock. Clearly the multipliers (sum and max) are far lower at all horizons and over both regimes than the ones corresponding to the anticipated spending shock, and even though the multipliers during recessions are still larger than over expansions the difference in multipliers markedly narrows.

The last rows of each panel in the Table 4 shows the counterfactual multipliers conditional to a fixed level of confidence. Observe that under expansions the counterfactual multipliers are so much lower than the unconstrained ones, while during recessions the difference between the counterfactual and the baseline multipliers is not that large. Hence, unlike the previous section, during recessions the size of the fiscal multiplier does not seem to be significantly reduced when confidence is held constant. Following this analysis Figure 8 shows the distribution of difference in multipliers between recessions and expansions for the unconstrained (top panel) and the counterfactual (bottom panel) multipliers. Note that now the difference in multipliers (max and sum) is always different from zero even for the counterfactual case, suggesting that for an unanticipated government spending shock the confidence reaction does not explain nonlinear fiscal multipliers.

Recalling that the measure of confidence conveys consumers expectations about future economic activity, these results indicate that a news shock provides information related to future movements in government purchases which significantly influences the consumers expectations about the economy, which in turn determines an important fraction of output level, and hence the fiscal multiplier, during recessions. While an innovation in the fiscal spending variable lacks this kind of information, being the consumer expectation reaction unable to explain the difference in fiscal multipliers. Therefore the overall findings suggest that the reason behind the role of confidence is the information about future government spending contained by the news shocks.

## 5 Conclusions

This paper investigates the role of consumer confidence in determining the effects that an anticipated (news) government spending shock has on the economic activity within a nonlinear framework. To do so I quantify the size of the fiscal multiplier by implementing a Smooth-Transition VAR model endowed with government spending, confidence, output and a measure of government spending news. This exercise allows me to identify the fundamental fiscal shocks and disentangle the effects that *anticipated* and *unanticipated* spending shocks have on confidence and output during recessions and expansions. Following Forni and Gambetti (2014), I overcome the issue of non-fundamentality by including in the estimated system a measure of government spending news defined as the sum of forecast revisions from the Survey of Professional Forecasters. I show that such a measure of spending news is able to predict both the future movements in government spending and other measure of fiscal news used in the literature.

My results point to a positive and significant response of confidence and output to an anticipated (news) spending shock during recessions. Differently, over expansions, the responses are statistically insignificant. The fiscal multiplier during recessions is found to be statistically larger than one and different from the one estimated over expansions. Importantly, I show that when confidence is held constant the multipliers are not anymore statistically different across regimes. This result points to the role of confidence as a key driver of the response of output to anticipated fiscal stimulus during recessions.

Finally, I contrast the previous results with those conditional on an unanticipated government spending shock. I find the fiscal multiplier in general to be lower than that corresponding to the anticipated spending shock, and never statistically larger than one. Interestingly, for an unanticipated spending shock confidence does not turn out to be important in explaining nonlinear fiscal multipliers. These findings indicate that an anticipated (news) spending shock provides relevant information related to future movements in government spending which significantly influences the consumers confidence, which in turn determines an important fraction of output during recessions. While an unanticipated spending shock does not convey this kind of information. Hence, the reason behind the role of confidence is the information about a future fiscal stimulus conveyed by the news shocks rather than the fiscal stimulus itself. It follows, therefore, that confidence plays an important role in the transmission of news about future fiscal policy into the economic activity.

The results of this paper highlight the importance of providing information about future public spending when taking expansionary fiscal policy in order to stimulate the economic activity during recessionary phases. Credible announcements about concrete increases in government purchases may be key in boosting aggregate confidence, and thus boosting output, during a period of economic slack.

## Appendix A - Estimation procedure of the nonlinear model

The STVAR model (2)-(5) is estimated by using maximum likelihood methods.<sup>23</sup> The log-likelihood of the model is the following:

$$\log L = \text{const} - \frac{1}{2} \sum_{t=1}^T \log |\mathbf{\Omega}_t| - \frac{1}{2} \sum_{t=1}^T \mathbf{u}'_t \mathbf{\Omega}_t^{-1} \mathbf{\mu}_t \quad (\text{A1})$$

where  $\mathbf{\mu}_t = \mathbf{X}_t - (1 - F(z_{t-1}))\mathbf{\Pi}_E(L)\mathbf{X}_{t-1} - F(z_{t-1})\mathbf{\Pi}_R(L)\mathbf{X}_{t-1}$  is the vector of residuals. Given the high non-linearity of the model and its many parameters  $\mathbf{\Psi} = \{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E, \mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$ , the estimation by using standard optimization routines becomes problematic. Therefore I estimate the model by following the procedure used by Auerbach and Gorodnichenko (2012) which is described below.

Note that conditional on  $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$  the model is linear in the lag polynomials  $\{\mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$ . Thus, for a given guess on the parameters  $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$  I can estimate the coefficients  $\{\mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$  by using weighted least squares where the estimates of the coefficients must minimize  $\frac{1}{2} \sum_{t=1}^T \mathbf{\mu}'_t \mathbf{\Omega}_t^{-1} \mathbf{\mu}_t$ . First we rewrite the regressors in the following way:

Let  $\mathbf{W}_t = [F(z_{t-1})\mathbf{X}_{t-1} \quad (1 - F(z_{t-1}))\mathbf{X}_{t-1} \quad \dots \quad F(z_{t-1})\mathbf{X}_{t-p} \quad (1 - F(z_{t-1}))\mathbf{X}_{t-p}]$  be the extended vector of regressors and  $\mathbf{\Pi} = [\mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)]$ , so we can write  $\mathbf{\mu}_t = \mathbf{X}_t - \mathbf{\Pi}\mathbf{W}'_t$ . Therefore the objective function is:

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

Then, it can be proved that the first order condition to obtain  $\mathbf{\Pi}$  is:

$$\text{vec}\mathbf{\Pi}' = (\sum_{t=1}^T [\mathbf{\Omega}_t^{-1} \otimes \mathbf{W}'_t \mathbf{W}_t])^{-1} \text{vec}(\sum_{t=1}^T \mathbf{W}'_t \mathbf{X}_t \mathbf{\Omega}_t^{-1}) \quad (\text{A2})$$

This procedure works iterating on  $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$ , obtaining  $\mathbf{\Pi}$  and the likelihood (A1) for each set of values for  $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$  until the optimum is achieved. Because the model is highly nonlinear in its parameters, several local optima might be founded; therefore one should try different starting values for  $\{\gamma, \mathbf{\Omega}_R, \mathbf{\Omega}_E\}$ .

To ensure that the matrices  $\{\mathbf{\Omega}_R, \mathbf{\Omega}_E\}$  are positive definite I work with an alternative vectors of parameters,  $\mathbf{\Psi} = \{\gamma, \text{chol}(\mathbf{\Omega}_R), \text{chol}(\mathbf{\Omega}_E), \mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L)\}$ , where *chol* indicates the Cholesky decomposition operator. Moreover and given the non-linearity of the model I estimate the parameters by using Markov Chain Monte Carlo (MCMC) algorithm developed by Chernozhukov and Hong (2003) (henceforth CH). The advantage of this method is that not only deliver a global optima but also the densities for the parameters estimates.

<sup>23</sup> This section highly reflects the Auerbach and Gorodnichenko's (2012) "Appendix: Estimation Procedure".

To implement CH we use Metropolis-Hastings algorithm. For a starting value  $\Psi^{(0)}$ , the procedure to construct chains of length  $N$  is as follows :

**Step 1:**

Draw a candidate vector of parameters values as  $\Theta^{(n)} = \Psi^{(n)} + \psi^{(n)}$  for the chain's  $n+1$  state, where  $\Psi^{(n)}$  is the current  $n$  state of the vector of parameters values in the chain and  $\psi^{(n)}$  is a vector of i.i.d. shocks taken from  $N(0; \Omega_\Psi)$  where  $\Omega_\Psi$  is a diagonal matrix.

**Step 2:**

Take the chain's  $n+1$  state as  $\Psi^{(n+1)} = \Theta^{(n)}$  with probability  $\min\{1, L(\Theta^{(n)})/L(\Psi^{(n)})\}$ , where  $L(\Theta^{(n)})$  is the value of the objective function conditional on the candidate vector of parameters values, and  $L(\Psi^{(n)})$  the value of the objective function conditional on the current state of the chain. Otherwise, take  $\Psi^{(n+1)} = \Psi^{(n)}$ .

The starting value  $\Psi^{(0)}$  is computed by approximating the model so that it can be written as regressing  $X_t$  on lags of  $X_t, X_t z_t, X_t z_t^2$ . Then the residuals from this regression are used to fit the equation for the reduced-form time-varying variance-covariance matrix of the STVAR by using maximum likelihood to estimate  $\Omega_R$  and  $\Omega_E$ , these estimates are used as starting values. By using the estimates  $\Omega_R$  and  $\Omega_E$  and a calibrated  $\gamma$  I can obtain  $\Omega_t$ . Finally, conditional on  $\Omega_t$  we compute the starting values for the lag polynomials  $\{\Pi_R(L), \Pi_E(L)\}$  using the equation (A2). The initial matrix  $\Omega_\Psi$  is calibrated to one percent of the parameters values, then is adjusted "on the fly" for the first 20.000 draws in order to generate an acceptance rate of around 0.3, as is proposed for this kind of simulations. I employ 100.000 draws for my estimates, and drop the first 20.000 draws.

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

In order to construct the confidence bands I use bootstrap procedure with 5000 interactions to obtain the distribution of the generated impulse responses. Then the confidence bands are computed as the selected percentiles from the bootstrapped distributions.

## Appendix B- Linearity Test

In order to test for nonlinear dynamics at a multivariate framework, I employ the linearity test described by Teräsvirta and Yang (2014). They propose to test of the null hypothesis of linearity against a (Logistic) Smooth Transition Vector Autoregressive with a single switching variable for the whole system.

Consider the  $p$  –dimensional  $n$ -order Taylor approximation around  $\gamma = 0$  of the logistic STVAR model (2):

$$\mathbf{X}_t = \boldsymbol{\Theta}'_0 \mathbf{Y}_t + \sum_{i=1}^n \boldsymbol{\Theta}'_i \mathbf{Y}_t z_t^i + \boldsymbol{\varepsilon}_t \quad (\text{A3})$$

where  $\mathbf{X}_t = [g_t \text{ conf}_t \ y_t \ \text{news}_t]'$  is the  $(p \times 1)$  baseline specification of vector of endogenous variables,  $\mathbf{Y}_t = [\mathbf{X}_{t-1}, \dots, \mathbf{X}_{t-k}, \boldsymbol{\alpha}]$  is the  $(k \times p + q) \times 1$  vector of exogenous variables including endogenous variables lagged  $k$  times and a column vector of constants  $\boldsymbol{\alpha}$ , and  $z_t$  is the switching variable. Moreover  $\boldsymbol{\Theta}_0$  and  $\boldsymbol{\Theta}_i$  are matrices of parameters. Following Teräsvirta and Yang (2014), the null hypothesis of linearity is  $H_0 : \boldsymbol{\Theta}_i = \mathbf{0} \ \forall i$ . In the present paper I fix the value of the order of the Taylor approximation to  $n = 1$ . Furthermore the number of endogenous variables is  $p = 4$ , the number of lags is  $k = 2$  and the number of exogenous variables is  $q = 1$ .

The test for linearity against the STVAR model is performed as follows:

- 1- Estimate the model under the null  $H_0 : \boldsymbol{\Theta}_i = \mathbf{0} \ \forall i$  (estimate the linear model) by regressing  $\mathbf{X}_t$  on  $\mathbf{Y}_t$ . Compute the residuals  $\tilde{\mathbf{E}}$  and the matrix residuals sum of squares  $\mathbf{RSS}_0 = \tilde{\mathbf{E}}' \tilde{\mathbf{E}}$ .
- 2- Regress  $\tilde{\mathbf{E}}$  on  $\mathbf{Y}_t$  and  $\mathbf{Z}_n$  where  $\mathbf{Z}_n = [\mathbf{Y}'_t z_t | \mathbf{Y}'_t z_t^2 | \dots | \mathbf{Y}'_t z_t^n]$ . Compute the residuals  $\tilde{\tilde{\mathbf{E}}}$  and the matrix residuals sum of squares  $\mathbf{RSS}_1 = \tilde{\tilde{\mathbf{E}}}' \tilde{\tilde{\mathbf{E}}}$ .
- 3- Compute the test-statistic

$$\begin{aligned} LM_{\chi^2} &= T \text{tr}\{\mathbf{RSS}_0^{-1}(\mathbf{RSS}_0 - \mathbf{RSS}_1)\} \\ &= T(p - \text{tr}\{\mathbf{RSS}_0^{-1} - \mathbf{RSS}_1\}) \end{aligned} \quad (\text{A4})$$

where  $\text{tr}\{\cdot\}$  indicates the trace of a matrix. Note that under the null hypothesis, the test statistic has an asymptotic  $\chi^2$  distribution with  $np(kp + q)$  degrees of freedom (36 in my case). The value of the test for the model in (2) is  $LM = 125$  with a corresponding p-value equal to zero. Therefore, I reject the null hypothesis of linearity in favour of a STVAR specification of the model. Furthermore, the null hypothesis of linearity can be rejected also for an order of the Taylor approximation  $n = 2$  and  $n = 3$ .



## Appendix C- Constancy of the Error Covariance Matrix Test

Following Yang (2014) I carry out a test of constancy of the error covariance matrix against the alternative of Smooth Transition. The proposed test assumes that an spectral decomposition of the time-varying error covariance matrix exists such that:

$$\mathbf{\Omega}_t = \mathbf{P}\mathbf{\Lambda}_t\mathbf{P}' \quad (\text{A5})$$

where the  $\mathbf{P}$  is a time-invariant orthogonal matrix such that  $\mathbf{P}\mathbf{P}' = \mathbf{I}_p$ ,  $\mathbf{I}_p$  being an identity matrix, and  $\mathbf{\Lambda}_t = \text{diag}(\lambda_{1t}, \dots, \lambda_{pt})$  whose elements are all positive. Notice that the above equations implies that the covariance matrix is time-varying in the way that the eigenvectors remain constant while the corresponding eigenvalues are allowed to vary over time.

Under this assumption, the log-likelihood function for observation  $t = 1, \dots, T$  with Gaussian distributed errors is:

$$\begin{aligned} \log L_t &= c - \frac{1}{2} \log |\mathbf{\Omega}_t| - \frac{1}{2} \boldsymbol{\mu}_t' \mathbf{\Omega}_t^{-1} \boldsymbol{\mu}_t \\ &= c - \frac{1}{2} \log |\mathbf{\Lambda}_t| - \frac{1}{2} \boldsymbol{\omega}_t' \mathbf{\Lambda}_t^{-1} \boldsymbol{\omega}_t \\ &= c - \frac{1}{2} \sum_{i=1}^p (\log \lambda_{it} + \omega_{it}^2 \lambda_{it}^{-1}), \end{aligned}$$

where  $\boldsymbol{\omega}_t = \mathbf{P}' \boldsymbol{\mu}_t = (\omega_{1t}, \dots, \omega_{pt})'$  contains the errors. The null hypothesis to be tested is:

$$H_0: \lambda_{it} = \lambda_i, \quad i = 1, \dots, p. \quad (\text{A6})$$

Moreover, the  $LM$  test-statistic has the following form:

$$LM_{\chi^2} = \frac{1}{2} \sum_{i=1}^p \left[ \left( \sum_{t=1}^T \tilde{g}_{it} \tilde{\mathbf{z}}_{it}' \right) \left( \sum_{t=1}^T \tilde{\mathbf{z}}_{it} \tilde{\mathbf{z}}_{it}' \right)^{-1} \left( \sum_{t=1}^T \tilde{g}_{it} \tilde{\mathbf{z}}_{it} \right) \right] \quad (\text{A7})$$

where  $\tilde{g}_{it} = \tilde{\omega}_{it}^2 / \tilde{\lambda}_{it} - 1$  and  $\tilde{\mathbf{z}}_{it}$  is a vector of variables determining the time-varying components  $\lambda_{it}$ . To test for the constancy of the covariance matrix against a Smooth Transition specification  $\tilde{\mathbf{z}}_{it}$  is defined as the  $n$ -order Taylor approximation of the of the transition function (5) around  $\gamma = 0$ . In the present paper I use a second-order approximation.

As Yang (2014) shows the test can be computed in the following way:

- 1- Estimate the model under the null hypothesis of constant covariance matrix. Collect the estimated residuals  $\tilde{\boldsymbol{\mu}}_t$ ,  $t = 1, \dots, T$ . Compute the corresponding covariance matrix  $\tilde{\boldsymbol{\Omega}}$  and the eigenvalue decomposition  $\tilde{\boldsymbol{\Omega}} = \tilde{\mathbf{P}}\tilde{\boldsymbol{\Lambda}}\tilde{\mathbf{P}}'$ , where  $\tilde{\boldsymbol{\Lambda}} = \text{diag}(\tilde{\lambda}_1, \dots, \tilde{\lambda}_p)$ .
- 2- Compute the transformed residuals  $\tilde{\boldsymbol{\omega}}_t = \tilde{\mathbf{P}}'\tilde{\boldsymbol{\mu}}_t$  and  $\tilde{g}_{it} = \tilde{\omega}_{it}^2/\tilde{\lambda}_{it} - 1$ , for  $t = 1, \dots, T$ ,  $i = 1, \dots, p$ . Compute the sum of squared  $\tilde{\boldsymbol{g}}_i$  as  $SSG_i = \tilde{\boldsymbol{g}}_i'\tilde{\boldsymbol{g}}_i$ .
- 3- For each equation, regress  $\tilde{g}_{it}$  on  $\tilde{\boldsymbol{z}}_{it}$ . Collect the residuals  $\tilde{\boldsymbol{v}}$  and compute the residuals sum of squares  $RSS_i = \tilde{\boldsymbol{v}}_i'\tilde{\boldsymbol{v}}_i$ .
- 4- Compute the  $LM$  test-statistic as follows:

$$LM_{\chi^2} = \sum_{i=1}^p T \frac{SSG_i - RSS_i}{SSG_i} \quad (\text{A8})$$

It can be proven that under regularity conditions the  $LM$  statistic is asymptotically  $\chi^2$  distributed with  $p \times n$  degrees of freedom (8 in my case). The value of the test for the baseline vector of endogenous variables  $\mathbf{X}_t = [g_t \text{ conf}_t \ y_t \ \text{news}_t]'$  is  $LM = 27.46$ , with a corresponding p-value approximately equal to zero. Therefore, the null hypothesis of constant error covariance matrix is rejected in favour of a Smooth Transition alternative specification.

## Appendix D- Confidence Index:

The index of Consumer Expectation is composed by following three forward-looking questions:<sup>24</sup>

Q1= Looking ahead, do you think that a year from now you (and your family living there) will be better off financially, or worse off, or just about the same as now?

**Answer choice:** Better now, Same, Worse, Don't know.

Q2= Now turning to the business conditions in the country as a whole, do you think that during the next twelve months we'll have good times financially, or bad times, or what?

**Answer choice:** Will be better off, Same, Will be worse Off; Don't know.

Q3= Looking ahead, which would you say is more likely-that in the country as a whole we'll have continuous good times during the next five years or so, or that we will have periods of widespread unemployment or depression, or what?

**Answer choice:** Good times, Good with qualifications, Pro-Con, Bad with qualifications, Bad times, Don't know.

The index of Consumers Expectations is computed as follows:

-First compute the relative scores for each of the three questions as the percent giving favorable replies minus the percent giving unfavorable replies, plus 100;

- Then apply the formula bellow:

$$ICE = \frac{Q_1 + Q_2 + Q_3}{4.1134} + 2 \quad (A9)$$

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<sup>24</sup> For further details see <http://www.sca.isr.umich.edu/fetchdoc.php?docid=24770>.

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Independent variable	Dependent variable
	<i>g. shock<sub>t</sub></i>
<b>Expectation revision</b>	
<i>news(0,0)<sub>t-1</sub></i>	0.38
<i>news(1,1)<sub>t-1</sub></i>	<b>0.01</b>
<i>news(2,2)<sub>t-1</sub></i>	<b>0.02</b>
<i>news(3,3)<sub>t-1</sub></i>	0.13
<b>Sum of expectation's revisions</b>	
<i>news(1,2)<sub>t-1</sub></i>	<b>0.00</b>
<i>news(1,3)<sub>t-1</sub></i>	<b>0.05</b>

**Table 1. Granger-causality test of government spending shocks: Linear model.** P-values of Granger-causality test corresponding to the prediction of the VAR estimated government spending shocks by the different specifications of one-period lagged News variable. Values in bold indicate a predictive power found to be significant at a 10% confidence level. The structural spending shocks are drawn from VAR(4) containing, in the following order, the log of real per capita government spending, the confidence index and the log of real per capita output. The sample used is 1981:Q4-2013:Q1. The test considers standard errors robust to heteroskedasticity and serial correlation.

Explained variable	<i>news<sub>t-1</sub></i>	<i>Ramey<sub>t-1</sub></i>
<i>Ramey<sub>t</sub></i> (1981:Q4-2013:Q1)	<b>0.07</b>	
<i>Ramey<sub>t</sub></i> (1986:Q4-2013:Q1)	<b>0.02</b>	
<i>news<sub>t</sub></i> (1981:Q4-2013:Q1)		0.94
<i>news<sub>t</sub></i> (1986:Q4-2013:Q1)		0.92

**Table 2. Granger-causality test: Ramey's vs. News variable.** P-values of Granger-causality test for VAR (1) including the Ramey's and News variable. Values in bold indicate a predictive power found to be significant at a 10% confidence level. The VAR is estimated for the sample 1981:Q4-2013:Q1. Moreover, given that the first twenty observations of this sample are all zero for Ramey's variable, I also use a shorter sample starting from 1986:Q4. The Ramey variable series is the one employed in Ramey and Zubairy (2014).

Independent variable	Dependent variable
	<i>g. shock. STVAR<sub>t</sub></i>
<b>Expectation revision</b>	
<i>news(0,0)<sub>t-1</sub></i>	0.30
<i>news(1,1)<sub>t-1</sub></i>	<b>0.06</b>
<i>news(2,2)<sub>t-1</sub></i>	<b>0.00</b>
<i>news(3,3)<sub>t-1</sub></i>	0.18
<b>Sum of expectation's revisions</b>	
<i>news(1,2)<sub>t-1</sub></i>	<b>0.00</b>
<i>news(1,3)<sub>t-1</sub></i>	<b>0.08</b>

**Table 3. Granger-causality test of government spending shocks: Nonlinear model.** P-values of Granger-causality test corresponding to the prediction of the STVAR estimated government spending shocks by the different specifications of one-period lagged News variable. Values in bold indicate a predictive power found to be significant at a 10% confidence level. The structural spending shocks are drawn from the Smooth-Transition VAR model containing, in the following order, the log of real per capita government spending, the confidence index and the log of real per capita output. The sample used is 1981:Q4-2013:Q1. The test considers standard errors robust to heteroskedasticity and serial correlation.

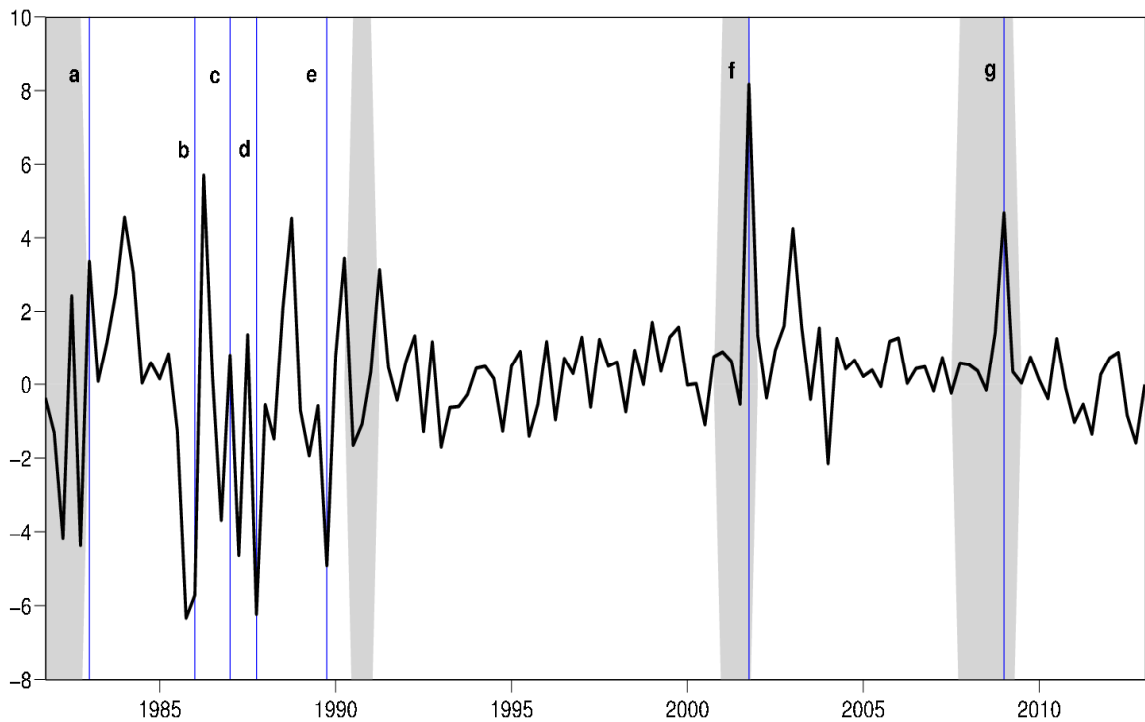


<b>Max multipliers</b>			
	$\frac{\max\{y_h\}_{h=1}^5}{\max\{g_h\}_{h=1}^5}$	$\frac{\max\{y_h\}_{h=1}^8}{\max\{g_h\}_{h=1}^8}$	$\frac{\max\{y_h\}_{h=1}^{16}}{\max\{g_h\}_{h=1}^{16}}$
Linear	2.14 [0.88 4.45]	1.96 [0.77 4.53]	1.80 [0.75 4.19]
Expansion	0.67 [0 1.87]	0.67 [0 1.93]	0.67 [0 1.93]
Recession	3.41 [2.62 4.41]	3.09 [2.33 4.01]	3.09 [2.33 4.04]
Expansion w/o conf.	0.98 [0.08 3.74]	0.99 [0.10 5.21]	0.86 [0.07 6.51]
Recession w/o conf.	2.71 [2.08 3.50]	2.79 [2.13 3.60]	2.88 [2.18 3.79]
<b>Sum multipliers</b>			
	$\frac{\sum_{h=1}^5 Y_h}{\sum_{h=1}^5 G_h}$	$\frac{\sum_{h=1}^8 Y_h}{\sum_{h=1}^8 G_h}$	$\frac{\sum_{h=1}^{16} Y_h}{\sum_{h=1}^{16} G_h}$
Linear	1.89 [0.06 4.34]	1.59 [-0.47 4.38]	0.79 [-1.81 4.09]
Expansion	0.39 [-0.63 1.61]	0.41 [-0.71 1.77]	0.20 [-1.29 1.89]
Recession	3.70 [2.75 5.08]	3.22 [2.26 4.47]	3.16 [1.69 4.95]
Expansion w/o conf.	0.58 [-0.62 2.79]	0.46 [-1.14 3.34]	-0.64 [-3.40 3.19]
Recession w/o conf.	2.43 [1.78 3.24]	2.54 [1.74 3.68]	3.07 [1.79 4.93]

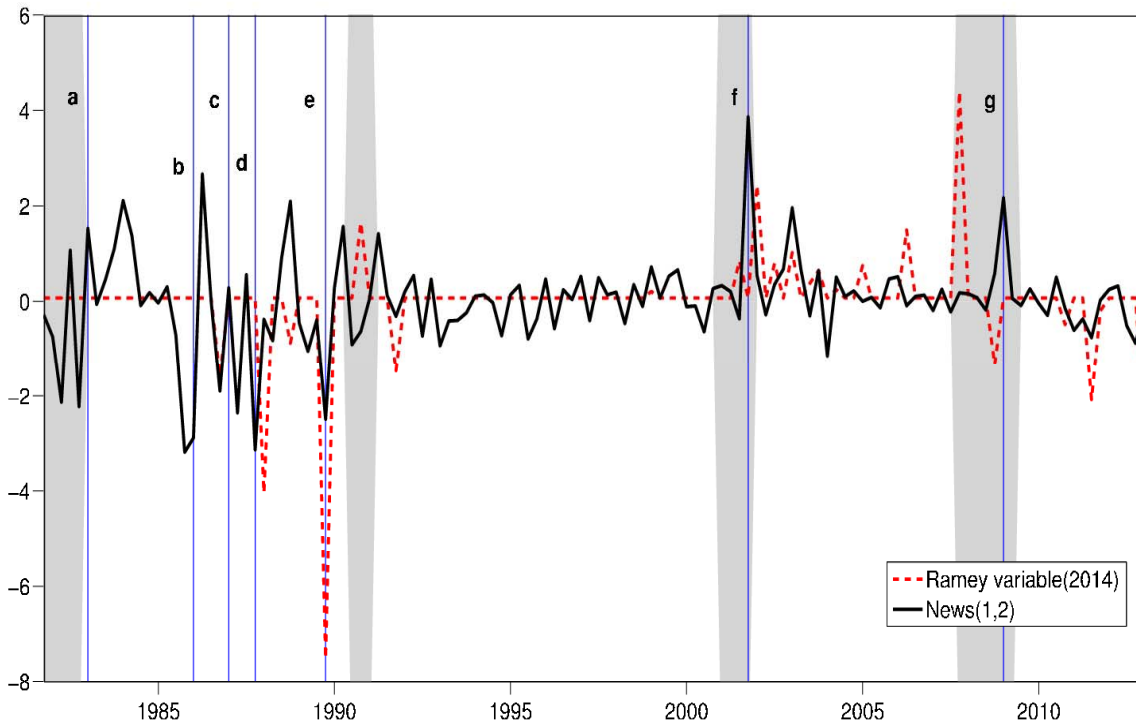
**Table 4. Fiscal Multiplier: Anticipated (news) government spending shock.** Fiscal multipliers for the baseline specification containing, in that order, the log of real per capita government spending, the confidence index, the log of real per capita GDP and the News variable. The shock is the last of the Cholesky decomposition. The last rows of each panel (max and sum) shows the fiscal multipliers conditional to a fixed level of confidence. The estimated multipliers are scaled by the sample average of  $Y/G$  in order to transform elasticities into dollars changes. The numbers in brackets indicate the 68% confidence intervals from the distribution of multipliers.

<b>Max multipliers</b>			
	$\frac{\max\{y_h\}_{h=1}^5}{\max\{g_h\}_{h=1}^5}$	$\frac{\max\{y_h\}_{h=1}^8}{\max\{g_h\}_{h=1}^8}$	$\frac{\max\{y_h\}_{h=1}^{16}}{\max\{g_h\}_{h=1}^{16}}$
Linear	0.85 [0.49 1.24]	0.85 [0.49 1.25]	0.86 [0.50 1.27]
Expansion	0.62 [0.45 0.79]	0.62 [0.45 0.79]	0.62 [0.49 0.79]
Recession	1.24 [0.80 1.73]	1.24 [0.81 1.73]	1.24 [0.81 1.74]
Expansion w/o conf.	0.40 [0.21 0.60]	0.37 [0.18 0.58]	0.31 [0.09 0.55]
Recession w/o conf.	0.79 [0.52 1.07]	0.81 [0.54 1.12]	0.82 [0.55 1.54]
<b>Sum multipliers</b>			
	$\frac{\sum_{h=1}^5 Y_h}{\sum_{h=1}^5 G_h}$	$\frac{\sum_{h=1}^8 Y_h}{\sum_{h=1}^8 G_h}$	$\frac{\sum_{h=1}^{16} Y_h}{\sum_{h=1}^{16} G_h}$
Linear	0.10 [-0.64 0.86]	-0.33 [-1.25 0.64]	-1.45 [-3.31 0.23]
Expansion	0.14 [-0.16 0.44]	-0.03 [-0.43 0.39]	-0.49 [-1.30 0.29]
Recession	1.17 [0.74 1.60]	1.15 [0.60 1.71]	1.07 [-0.18 2.05]
Expansion w/o conf.	-1.02 [-1.57 -0.51]	-1.80 [-2.81 -0.97]	-3.92 [-8.20 -1.84]
Recession w/o conf.	0.72 [0.37 1.05]	0.81 [0.33 1.28]	1.13 [0.18 1.91]

**Table 5. Fiscal Multipliers: Unanticipated government spending shock.** Estimated fiscal multipliers for a shock on the first variable of the baseline specification. The last rows of each panel (max and sum) shows the fiscal multipliers conditional to a fixed level of confidence. The estimated multipliers are scaled by the sample average of  $Y/G$  in order to transform elasticities into dollars changes. The numbers in brackets indicate the 68% confidence intervals from the distribution of multipliers.



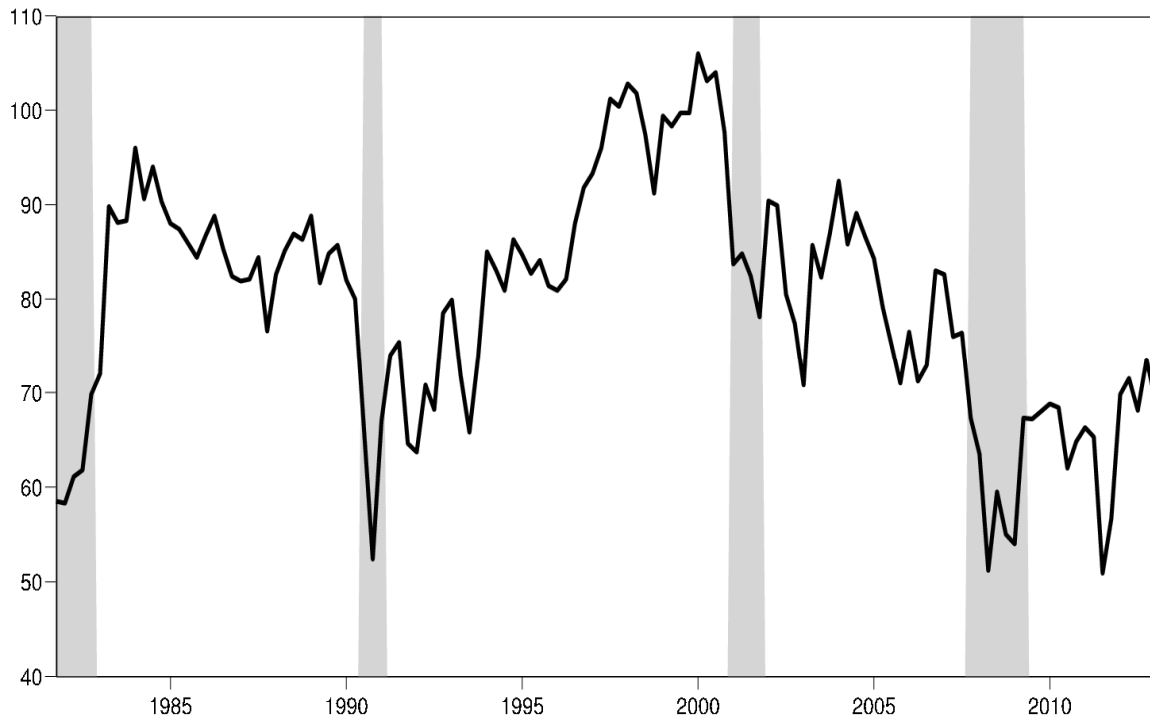
**Figure 1. News variable series and exogenous fiscal policy episodes.** The black solid line depicts the series for  $news_t(1,2)$ . The vertical blue lines correspond to the following episodes: (a) 1983Q1: Reagan’s “Evil Empire” and “Star Wars” speeches; (b) 1986Q1: Perestrojka; (c) 1987Q1: Senate elections won by Democrats a quarter before; (d) 1987Q4: Spending cuts as for the Pentagon; (e) 1989Q4: The fall of the Berlin Wall; (f) 2001Q4: War in Afghanistan; (g) 2010Q4: Obama’s Stimulus package. The shaded regions indicate the recessions as dated by the NBER.



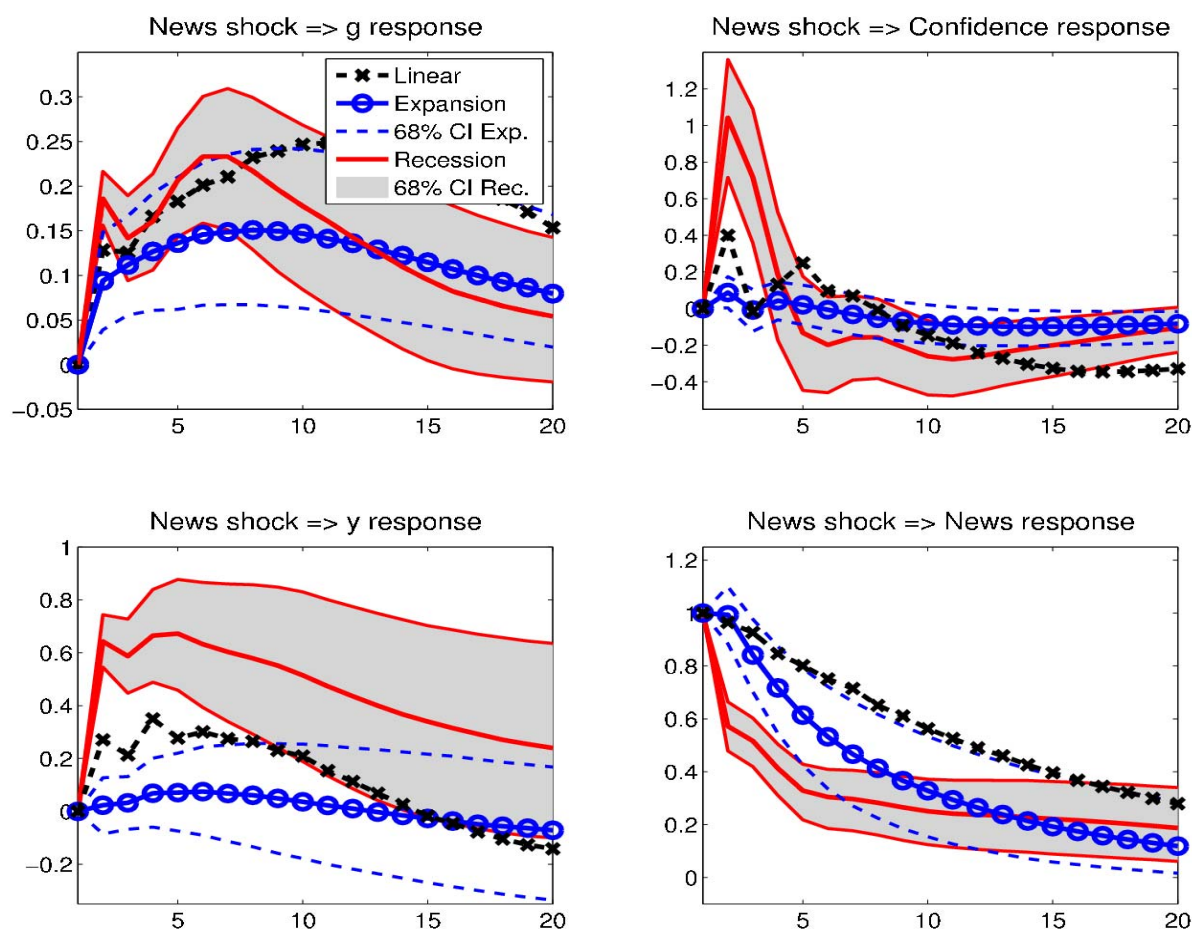
**Figure 2. News variable vs. Ramey’s variable.** The black solid line depicts the series for  $news_t(1,2)$  and the red dashed line draws the Ramey’s variable. The Ramey’s variable is computed as the present value of the expected government expending changes due to foreign political events (following Ramey (2011), each observation is divided by nominal GDP of the previous period). Both series shown in this Figure are standardized. The vertical blue lines correspond to the following episodes: (a) 1983Q1: Reagan’s “Evil Empire” and “Star Wars” speeches; (b) 1986Q1: Perestrojka; (c) 1987Q1: Senate elections won by Democrats a quarter before; (d) 1987Q4: Spending cuts as for the Pentagon; (e) 1989Q4: The fall of the Berlin Wall; (f) 2001Q4: War in Afghanistan; (g) 2010Q4: Obama’s Stimulus package. The shaded regions indicate the recessions as dated by the NBER.



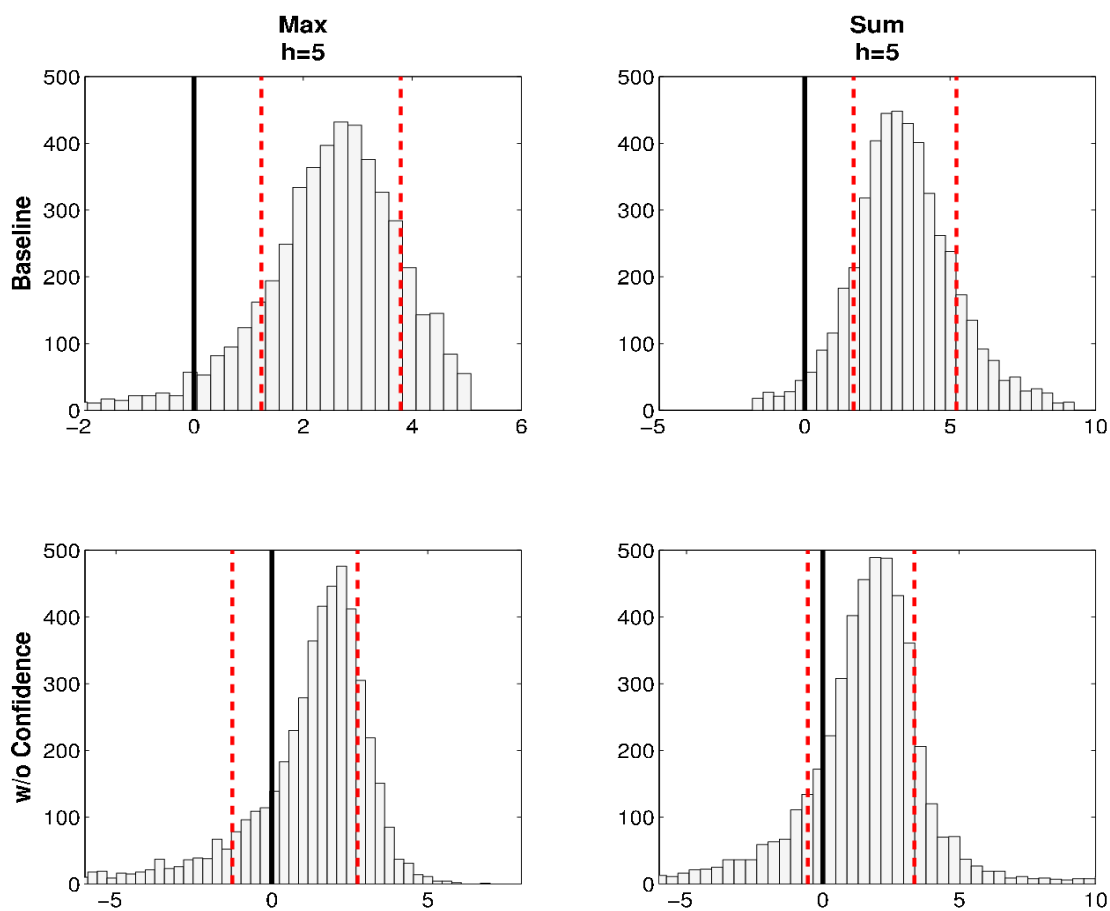
**Figure 3. Transition Function.**  $F(z_t)$  and the NBER recession dates, we can note how the shaded regions indicating the recessions defined by the NBER coincide with the picks of the black solid line indicating the probability of being in the recessionary regime  $F(z_t)$ .



**Figure 4. Consumers Confidence.** The index of Consumers Expectations and NBER recession dates. Note that the negative spikes of the confidence index (black solid line) coincide with the recessions defined by the NBER (shaded region).

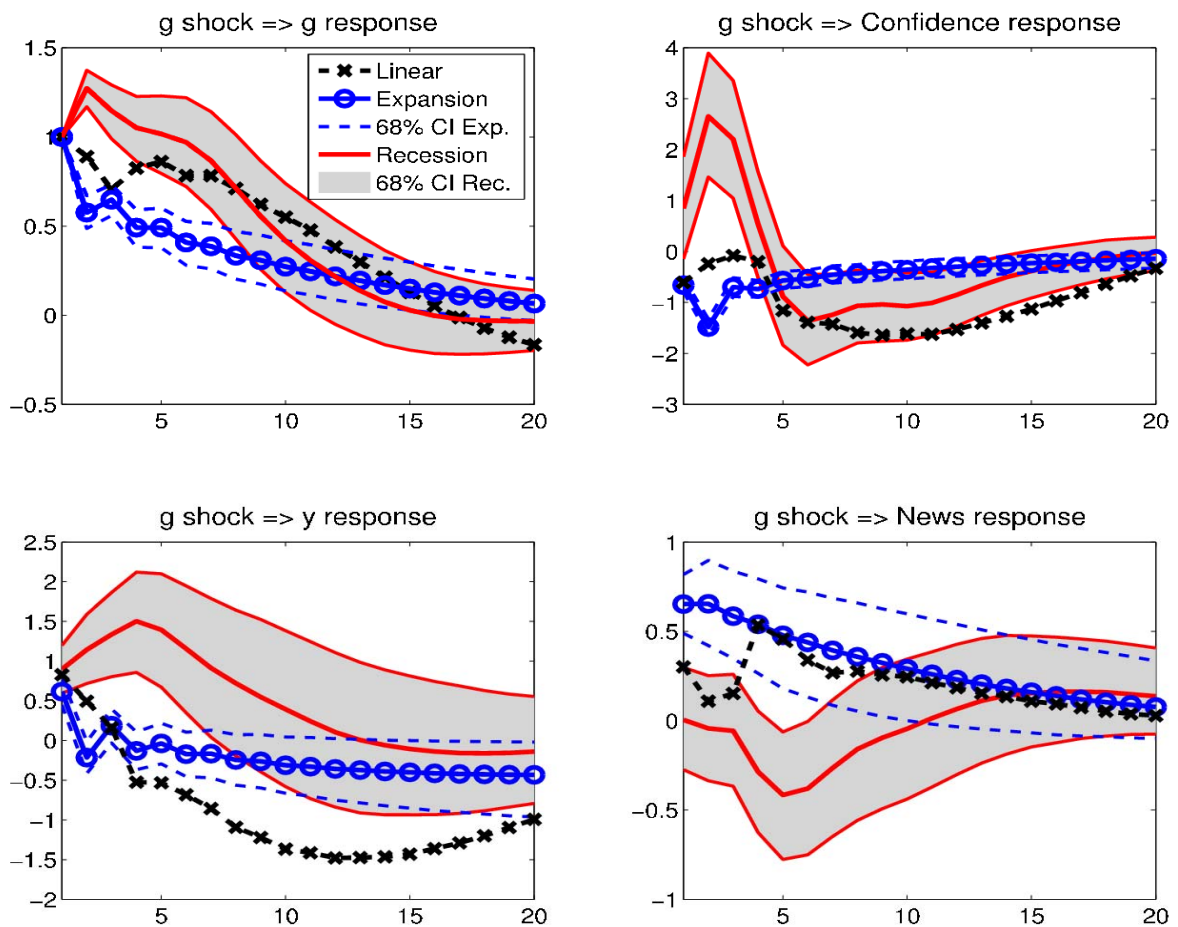


**Figure 5. IRFs to an anticipated (news) government spending shock normalized to one: Recession vs. Expansion.** The blue circled lines draw the median responses of the variables during expansions while the red solid lines depict the median responses during recessions. The black dash-crossed lines indicate the median responses for the linear model. The 68% confidence bands are shown by the blue dashed lines (expansions) and the shaded areas (recessions). The shock is the last of the Cholesky decomposition for the baseline specification including, in the that order, the log of real per capita government spending, the confidence index, the log of real per capita GDP and the News variable. The output responses are scaled by the sample average of  $Y/G$  in order to convert them in the same units than those of government spending, hence both responses are comparable.

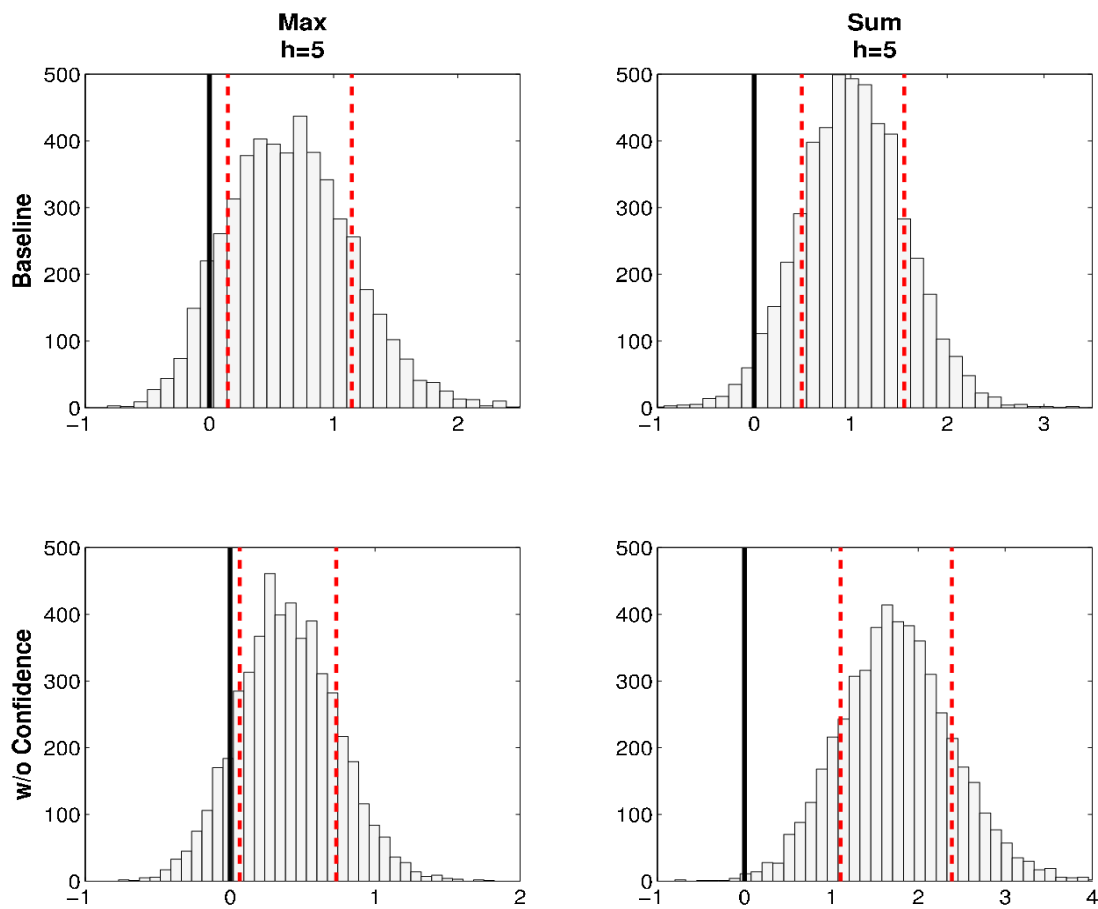


**Figure 6. Difference in multipliers between expansions and recessions : Anticipated (news) government spending shock.** The histograms depict the distribution of the difference in multipliers (max and sum) for the short run of 5 quarters. The top panel shows the distributions for the baseline specification while the bottom panel draws the distributions for the counterfactual specification conditional to a fixed level of confidence. The red dashed lines represent 68% confidence intervals. The empirical densities of the difference in multipliers are obtained by subtracting a realization of the multiplier in expansions from a realization of the multiplier in recessions for a number of times equal to 5,000. Note that when confidence is held constant the difference in multipliers is not statistically different from zero.





**Figure 7. IRFs to an unanticipated government spending shock normalized to one: Recession vs. Expansion.** The blue circled lines draw the median responses of the variables during expansions while the red solid lines depict the median responses during recessions. The black dash-crossed lines indicate the median responses for the linear model. The 68% confidence bands are shown by the blue dashed lines (expansions) and the shaded areas (recessions). The shock is the first of the Cholesky decomposition for the baseline specification including, in the that order, the log of real per capita government spending, the confidence index, the log of real per capita GDP and the News variable. The output responses are scaled by the sample average of  $Y/G$  in order to convert them in the same units than those of government spending, hence both responses are comparable.



**Figure 8. Difference in multipliers between expansions and recessions: Unanticipated government spending shock.** The histograms depict the distribution of the difference in multipliers (max and sum) for the short run of 5 quarters. The top panel shows the distributions for the baseline specification while the bottom panel draws the distributions for the counterfactual specification conditional to a fixed level of confidence. The red dashed lines represent 68% confidence intervals. The empirical densities of the difference in multipliers are obtained by subtracting a realization of the multiplier in expansions from a realization of the multiplier in recessions for a number of times equal to 5,000. Note that now when confidence is held constant the difference in multipliers is still statistically significant, indicating that for an unanticipated government spending shock the confidence reaction does not explain state-dependent fiscal multipliers.

## **Chapter 2**

# **Fiscal-Monetary Policy Mix in Recessions and Expansions**



# Fiscal-Monetary Policy Mix in Recessions and Expansions \*

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## Abstract

This paper studies the role of monetary policy in determining the size of the fiscal spending multiplier in recessions and expansions as for the U.S. economy. To quantify the size of state-dependent fiscal multipliers I estimate a nonlinear VAR model endowed with fiscal and monetary variables. I carefully separate anticipated and unexpected fiscal spending shocks by jointly modeling fiscal spending and the measure of spending news proposed by Ramey (2011). My results indicate that the fiscal multiplier in recessions is larger than one and statistically different from that corresponding to expansions. Importantly, the role of monetary policy during recessions triggers a crowding out effect. In particular, a counterfactual exercise clearly have the role played for the systematic policy to emerge. These findings highlight the importance of jointly consider monetary and fiscal variables when studying the effects of a fiscal stimulus.

*Keywords:* Fiscal spending multiplier, Monetary Policy, Fiscal Policy, Nonlinear models, Smooth Transition Vector AutoRegressions, Generalized Impulse Responses.

*JEL codes:* C32, E32, E50, E62.

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## 1 Introduction

This paper aims to study the role of monetary policy in determining the effectiveness of anticipated fiscal policy shocks in recessions and expansions. The novelty of this study is that jointly considers fiscal stimulus and monetary variables within a nonlinear framework.

The possibility that the effects of government spending shocks may vary across the phases of the business cycle is mentioned by traditional Neo-Keynesian models and New Keynesian models with a binding zero lower bound.<sup>1</sup> Recently, several empirical studies have considered the possibility of spending-based fiscal stimulus having different effects depending on the stage of the business cycle. Among others, Auerbach and Gorodnichenko (2012, 2013a, 2013b), Bachmann and Sims (2012), Mitnik and Semmler (2012), Baum, Poplawski-Ribeiro, and Weber (2012), Caggiano *et al.* (2015), Figueres (2015) find the fiscal multiplier to be significantly larger during recessionary times. On the contrary, Ramey and Zubairy (2014) estimate fiscal multipliers by exploiting historical U.S. data and find no evidence indicating that the size of the multiplier varies depending on the state of the economy.

Furthermore, the importance of monetary policy in determining the effects of fiscal policy is mentioned in the literature by several studies. Rossi and Zubairy (2011) show that failing to recognize that both monetary and fiscal policy simultaneously affect macroeconomic variables might incorrectly attribute the fluctuations to the wrong source. Davig and Leeper (2010) and Leeper, Traum and Walker (2001) find that passive monetary policy produces consistently stronger fiscal multipliers, suggesting that the impact of a fiscal stimulus cannot be understood without studying monetary and fiscal policy jointly. Therefore movements in monetary variables may be key in determining the effects that a government spending shock has on the economic activity during a recessions.

To study the effects of an increase in government spending I quantify the size of state-dependent fiscal multipliers by employing a Smooth Transition Vector Autoregressive (STVAR) model which allows me to consistently estimate the responses to a spending shock in recessions and expansions. The monetary variables included in the model are the Consumer Price index and the corporate bond spread rate defined as the difference between the BAA and the AAA short run Moody's corporate bond rate. My focus on the corporate bond spread is motivated by the relevance that this monetary indicator has in explaining

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<sup>1</sup> An example is the IS-LM-AD-AS model. According to this model the size of the fiscal multiplier reaches large values during times of economic downturn ( as the AS curve is flat, there is a lower crowding out effect harming investment and consumption) and small values when the economy is booming ( as the AS curve is steep, there is a higher crowding out effect). Furthermore, Eggertsson (2009), Christiano, Eichenbaum, and Rebelo (2011) and Woodford (2011) study the effects of a government spending shock within a New Keynesian framework. They show that when the nominal interest rate is binding the zero lower bound, a deficit financed increase in government spending triggers an increase in inflation expectations, which in turn leads to a decrease in real interest rates, triggering in this way an increase in investment and consumption. Therefore, in such cases in absence of a crowding out effect the fiscal multiplier may reach values as high as 3.

movements on private investments and consumption, two important components of total output. Several studies have found movements in corporate bond spreads to convey relevant information on the evolution of the real economic activity that affects the main macroeconomics indicators (see, among others, Gertler and Lown, 1999; Zhang, 2002; Gilchrist, Yankov and Zakrajšek, 2009; Faust *et al.*, 2011; and Gilchrist and Zakrajšek, 2012). Importantly, Zhang (2002) shows that the corporate bond spread features a significant nonlinear dynamic across the business cycle, hence is a more suitable indicator to include in the framework of the present study. In addition, unlike other normally used monetary variables as the three-month Treasury Bill and the Fed Fund rate, the corporate bond spread is likely to be less affected for the zero lower bound episodes present in my sample.

Moreover, to identify the fundamental government spending shocks I enrich the model with the measure of spending news developed by Ramey (2011). This variable captures the expected present value of government expending changes due to foreign political events, being constructed by using the *Business Week* magazine (mainly) and additional newspaper sources. The main advantage of using this variable is that, in contrast to other measures used in the literature, the Ramey variable covers a longer sample including several recessionary episodes and thus allows for a more precise estimation of the fiscal stimulus effects.

Furthermore, to estimate the fiscal multiplier in recessions and in expansions, I compute Generalized Impulse Response Functions (GIRFs) which allow for endogenize the transition from a state to another after that a fiscal spending shock takes place. As Koop, Pesaran and Potter (1996) point out the GIRFs are history-dependent. In a recent paper Caggiano *et al.* (2015) find that the estimated fiscal multiplier is statistically different across states of the economy only for initial histories belonging to deep recessions and strong expansions. Therefore, in the present study I focus my attention on these two extremes events of the economy. Importantly, given that government spending and output enter in the system in logs, to convert elasticities into dollars changes I propose the use of a time-varying ex-post conversion factor that allows me to obtain a more accurate estimation of the fiscal multiplier. Finally, to study the role played by the systematic response of the corporate bond spread in determining the effectiveness of a fiscal stimulus, I compute the fiscal multiplier for the hypothetical situation where the spread bond rate remains constant, i.e., is does not react to movements in the system due to government spending shocks.

My main results are the following. First, the fiscal multiplier during recessions is statistically larger than that corresponding to expansions. Moreover, the fiscal multiplier in recessions is statistically larger than one. Second, corporate bond spread positively reacts to a spending shock during recessions, thus suggesting the existence of a mild crowding out effect of a fiscal stimulus. Third, a counterfactual simulation assuming stagnant corporate bond spread gives as result an even larger fiscal multiplier during recessions. In contrast, the response of corporate bond spread does not turn out to be important for determining the size of the fiscal multiplier during expansions.

The closest paper to mine is Ramey and Zubairy (2014). They investigate whether the government spending multiplier differs when the interest rates are near to the zero lower bound. They find no evidence of elevated fiscal multipliers during the zero lower bound state. With respect to them, I study the role of monetary policy in determining the size of state-dependent multipliers by explicitly insulating the response of monetary variables to a government spending shock. Importantly, I show that movements in monetary variables matter to determine the size of the fiscal multiplier during an economic slack. In particular, a government spending shock is found to trigger an increase in the corporate bond spread, which in turn reduces the effectiveness of a fiscal stimulus during a recession.

The rest of the paper is organized as follows. Section 2 presents the Smooth-Transition VAR (STVAR) model. Section 3 describes the computational details of the estimation of the state-dependent fiscal multipliers and show the main results. The last section concludes.

## 2 Econometric Method

### 2.1 Model Specification

To study the role played by monetary policy in determining the effects that a fiscal stimulus has on the economic activity I employ a two-regime Smooth-Transition VAR (STVAR) model developed by Granger and Teräsvirta (1993). The most relevant advantage of this model is that it allows for estimating responses differentiated across states of the economy while retaining enough information for each state. The model is described below:

$$\mathbf{X}_t = [1 - F(z_{t-1})]\mathbf{\Pi}_R(L)\mathbf{X}_t + F(z_{t-1})\mathbf{\Pi}_E(L)\mathbf{X}_t + \boldsymbol{\varepsilon}_t \quad (1)$$

$$\boldsymbol{\varepsilon}_t \sim N(0, \boldsymbol{\Sigma}), \quad (2)$$

$$F(z_t) = [1 + \exp\{-\gamma(z_t - c)\}]^{-1}, \quad \gamma > 0, \quad z_t \sim N(0, 1). \quad (3)$$

where  $\mathbf{X}_t$  indicates the vector of endogenous variables, while  $\mathbf{\Pi}_R(L)$  and  $\mathbf{\Pi}_E(L)$  are the matrices of coefficients accounting for the dynamic of the variables in  $\mathbf{X}_t$  during recessions and over expansions, respectively. The vector  $\boldsymbol{\varepsilon}_t$  contains the residuals from the reduce form, with zero mean and positive definite variance-covariance matrix  $\boldsymbol{\Sigma}$ . Finally, the crucial feature of the STVAR model is the transition function  $F(z_t)$  which governs the transition from one regime to another.  $F(z_t)$  is increasing in the standardized transition variable  $z_t$ , and it also depends on the parameters  $\gamma$  and  $c$ . The variable  $z_t$  is an indicator of the state of the economy normalized to have zero mean and unit variance. Note that the transition function is bounded between 0 and 1, hence, in the framework of this paper,  $F(z_t)$  indicates the probability of being in an expansion while  $[1 - F(z_t)]$  indicates the probability of being



in a recession.<sup>2</sup> The parameter  $\gamma$  defines the smoothness of the transition when  $z_t$  changes. Lower values of  $\gamma$  determine a smooth transition from expansion to recession regime, implying that more of the observations are considered to contain some information about the behavior of the economy in both regimes. Conversely, when  $\gamma$  is high the transition becomes abrupt, meaning that  $F(z_t)$  spends more time close to the  $\{0,1\}$  bounds. Moreover, notice that when  $\gamma = 0$  the model (1)-(3) falls back to a linear model. The location parameter  $c$  indicates the midpoint of the transition, i.e., it represents the inflection point in which  $F(z_t) = 1/2$  in the sense that in (1) the changing parameter matrix  $\mathbf{\Pi} = (1/2)(\mathbf{\Pi}_R + \mathbf{\Pi}_E)$ . Importantly,  $c$  controls what proportion of the sample the economy spends in each regime. In the present study the baseline specification of vector of endogenous variables is defined as  $\mathbf{X}_t = [Ramey_t \ g_t \ t_t \ y_t \ cpi_t \ spread_t]'$ , where  $Ramey_t$  is the expected present value of government expending changes expressed as a percentage of the previous quarter GDP,  $g_t$  is the log of the real per capita government spending,  $t_t$  is the log of real per capita tax revenues,  $y_t$  is the log of the real per capita GDP,  $cpi_t$  is the log of the Consumer Price Index and  $spread_t$  is the difference between the BAA and the AAA short run Moody's corporate bond rate. The variables are expressed in per capita terms by dividing by the total population. The sample spans the period 1939:Q1-2013:Q4 for U.S. data.<sup>3</sup> Figure 1 depicts the corporate bond spread along with the NBER recession dates. Observe that the corporate bond spread exhibits a nonlinear and countercyclical behavior. Moreover  $\mathbf{\Pi}_R(L)$  and  $\mathbf{\Pi}_E(L)$  are set to be lag polynomials of degree 3. Furthermore, following Auerbach and Gorodnichenko (2012), Bachmann and Sims (2012), Caggiano, Castelnuovo and Groshenny (2014), I define the transition variable  $z_t$  as a standardized moving average of the real per capita GDP quarter-on-quarter percentage growth rate.<sup>4</sup>

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<sup>2</sup> Auerbach and Gorodnichenko (2012) employs a Smooth-Transition Autoregressive model with a transition function defined as  $F(z_t)_{A\&G} = \exp\{-\gamma(z_t - c)\} / [1 + \exp\{-\gamma(z_t - c)\}]$  that indicates the probability of begin in a recession. Note that, in my setting the probability of being in a recession  $1 - F(z_t)$  equals to the transition function used by Auerbach and Gorodnichenko (2012) as  $F(z_t)_{A\&G} = 1 - F(z_t)$ .

<sup>3</sup> The series for all the variables, with the exception of CPI, are drawn from the data set corresponding to Ramey and Zubairy (2014). They provide a detailed description of the all series used in the section "Data Appendix". Notice that the Ramey and Zubairy's (2014) data set extends back until 1889. Nevertheless as the authors point out, due to the interpolation method, the series for government spending and GDP are quite noisy in the pre-1939 period. Therefore to avoid that this behavior affects the final estimates I employ a post-1939 sample. The series for CPI corresponding to the period 1939:Q1-2008:Q4 are the one provided by Ramey (2011), and updated till 2013:Q4 with the series obtained from the Federal Reserve Bank of St. Luis website. For more information about the data set, see <http://econweb.ucsd.edu/~vramey/research.html>.

<sup>4</sup> The transition variable  $z_t$  is computed as the standardized six quarters backward-looking moving average of real per capita GDP growth rate.

## 2.2 Evidence in Favor of Non-linearity

In order to detect nonlinear dynamics at a multivariate level, I carry out the linearity test presented by Teräsvirta and Yang (2014). This analysis consists in testing the null hypothesis of linearity for the dynamics of the system in (1) against the alternative of Smooth Transition Vector Autoregressive (STVAR). The results of test for the baseline specification of the vector of exogenous variables  $X_t$  indicates a clear rejection of the null hypothesis of linearity in favor of the STVAR specification. See Appendix A.

## 2.3 Model Estimation

As suggested by Hurbich and Teräsvirta (2013), the model (1)-(3) can be fully estimated by employing conditional maximum likelihood. Although it is possible to estimate both set of parameters  $\{\gamma, c\}$  and  $\{\Pi_R(L), \Pi_E(L)\}$ , there may exist specific numerical problems with the identification of  $\gamma$  in small samples. Teräsvirta, Tjøstheim and Granger (2010) point out that, when  $\gamma$  is large so that the model converges to a switching regression model, the slope of  $F(z_t)$  at  $c$  is steep and a large amount of observations in the neighborhood of  $c$  would be required to estimate  $\gamma$  accurately, being unlikely to find such a clusters in small samples.<sup>5</sup> Auerbach and Gorodnichenko (2012, 2013a) address this identification problem by imposing fixed values for the pair  $\{\gamma, c\}$ . They calibrate  $\{\gamma, c\}$  to match the observed values of transition function with the post-WWII US recessions frequencies defined by the NBER, that implies values of  $\gamma = 1.5$  and  $c = 0$ . This metric provides a transition as smooth as to allow STVAR model to retain enough information for each regime.

In the present study I proceed with the estimation of the STVAR model (1)-(3) by calibrating the transition function à la Auerbach and Gorodnichenko (2012). Consequently, the location parameter is set  $c = 0$  and the smoothness parameter  $\gamma$  is calibrated to 2.45 to ensure that the economy will be in recessions around 17 percent of the times, a frequency in line with the NBER business cycle dates for my sample. Hence a recession is defined as a period in which  $[1 - F(z_t)] > 0.83$  in such a way that with  $\gamma = 2.45$  the probability to be in recession is  $Pr([1 - F(z_t)] > 0.83) \approx 17\%$ . This calibration implies a threshold value  $\bar{z} \leq -0.65\%$  during the recessionary regime (i.e., when  $z_t \leq -0.65\%$ ,  $[1 - F(z_t)] > 0.83$ ). Figure 2 contrasts the probability of being in a recession  $[1 - F(z_t)]$  with the recessions dated by the NBER.

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<sup>5</sup> As Granger and Teräsvirta (1993) explain, this is because when the true  $\gamma$  is relatively large, then exist a large set of  $\gamma$ -values yielding almost the same  $F(z_t)$ . The transition functions corresponding to these  $\gamma$ -values deviate significantly from each other only in a small neighborhood of the location parameter  $c$ . Thus a large number of observations of the transition variable  $z_t$  would be needed in that neighborhood to accurately estimate  $\gamma$ . See also Bates and Watts (1988) and Seber and Wild (1989).

## 2.4 Identification of the Anticipated Government Spending Shock

The identification of the fundamental government spending shocks is a key aspect to be considered for estimating the fiscal multiplier. When working within a VAR framework, an issue that is likely to affect the identification of the spending shocks is the anticipation effect, a phenomenon also known as fiscal foresight. This is because rational agents anticipate future changes (news) in the fiscal policy while VARs only consider the present and past values of the fiscal variables. As it has been shown by several studies, in presence of anticipation effect, standard fiscal VARs may not embed enough information to recover the anticipated government spending shocks (see, among others, Ramey and Shapiro, 1998; Forni and Gambetti, 2010; Ramey, 2011; Forni and Gambetti, 2014; Leeper, Walker, Yang, 2013; Caggiano *et al.*, 2015; Figueres, 2015). Forni and Gambetti (2010) and Ramey (2011) show that government spending shocks estimated by using standard fiscal VARs are predictable, i.e., are non-fundamental. Importantly, Leeper, Walker, Yang (2013) prove that when the econometric analysis fails to account for the anticipation effect, the estimated tax multiplier may exhibit quantitative important bias.

Following Ramey (2011), I identify the anticipated government spending shocks by including the Ramey (news) variable as first in the vector of endogenous variable  $\mathbf{X}_t$  and orthogonalise the reduce-form VAR residuals via a Cholesky decomposition of the estimated covariance matrix. This measure of spending news is computed as the present value of the expected government expending changes due to foreign political events, being constructed by using the *Business Week* magazine and additional newspaper sources. As Ramey (2011) shows, when considering a sample long enough as to include the spending shocks related with WWII and the Korean War, her spending news variable has a significant predictive power about movements in government spending. Note that the sample here used spans back till 1939:Q1, so that containing these two episodes. Figure 3 shows the series for the Ramey variable along with the recessions dates as defined by the NBER. Observe that the Ramey variable exhibits variations during recessions as well as during expansions, thus providing enough information to identify the anticipated (news) government spending shocks during both states of the economy.

### 3 Results

#### 3.1 Generalized Impulse Responses

In order to analyze the effects of fiscal stimulus in recessions and in expansions I estimate impulse responses of the STVAR model (1)-(3) to an anticipated government spending shock as defined in section 2.4. As Koop, Pesaran and Potter (1996) point out, estimating impulse response functions in a nonlinear framework is not as straightforward as it may be in a linear setup. This is because the responses of the endogenous variables to a given shock at time  $t$  may affect the state of the model a time  $t + 1$  and hence the corresponding future responses. Thus nonlinear models generates impulse response functions that are history- and shock-dependent. Therefore Koop, Pesaran and Potter (1996) define a particular type of impulse response functions designed to tackle down these issues, called generalized impulse response functions (GIRFs). The GIRFs allow to take into account the feedback from the evolution of output in the vector  $\mathbf{X}_t$  to the transition variable  $z_t$  and thus to the transition function  $F(z_{t-1})$ . Basically the GIRF, at  $h$  periods ahead, for a given shock of size  $\delta$  hitting the system at time  $t$  and for a given initial history  $\boldsymbol{\omega}_{t-1}$  is defined as:

$$GIRF(h, \delta_t, \boldsymbol{\omega}_{t-1}) = E[\mathbf{X}_{t+h} | \delta_t, \boldsymbol{\omega}_{t-1}] - E[\mathbf{X}_{t+h} | \boldsymbol{\omega}_{t-1}] \quad (4)$$

where  $E[\cdot]$  is the expectation operator and  $\boldsymbol{\omega}_{t-1} = \{\mathbf{X}_{t-1}, \dots, \mathbf{X}_{t-p}; z_{t-1}\}$  contains the starting values for the lags in (1) as well as the transition variable  $z_{t-1}$  that gives the value for transition function  $F(z_{t-1})$  defined in (3). Therefore for a specific history  $\boldsymbol{\omega}_{t-1}$  the GIRFs are computed as the difference between the expectation of  $\mathbf{X}_{t+h}$  conditional on the shock  $\delta_t$  and the expectation of  $\mathbf{X}_{t+h}$  without a shock.

As Koop, Pesaran and Potter (1996) describe, for a given initial history  $\boldsymbol{\omega}_{t-1}$ , the above conditional expectations are estimated by randomizing over the reduce-form residuals of the estimated model (1)-(3) in the following way: First, draw with replacement a sequence of reduce-form residuals  $\boldsymbol{\varepsilon}^{(j)*} = \{\boldsymbol{\varepsilon}_t^*, \boldsymbol{\varepsilon}_{t+1}^*, \dots, \boldsymbol{\varepsilon}_{t+h}^*\}$  from the presumed distribution of  $\hat{\boldsymbol{\varepsilon}}_t$ . Second, recover the structural shocks  $\{\mathbf{e}_t^*, \mathbf{e}_{t+1}^*, \dots, \mathbf{e}_{t+h}^*\}$  by orthogonalizing the reduce-form residuals as  $\mathbf{e}^{(j)*} = \hat{\mathbf{C}}^{-1} \boldsymbol{\varepsilon}^{(j)*}$ , where  $\hat{\mathbf{C}}$  is the Cholesky factor of the corresponding residuals covariance matrix  $\hat{\boldsymbol{\Sigma}}$ . To compute a shock of size  $\delta$  in the *Ramey* variable, form another set of shocks  $\{\mathbf{e}_t^\delta, \mathbf{e}_{t+1}^\delta, \dots, \mathbf{e}_{t+h}^\delta\}$  by replacing the observation corresponding to the *Ramey* shock in  $\mathbf{e}_t^*$  by the perturbed shock  $e_{1t}^\delta = e_{1t}^* + \delta$ .<sup>6</sup> Then transform back to the reduce-form residuals:  $\boldsymbol{\varepsilon}^{(j)\delta} = \hat{\mathbf{C}} \mathbf{e}^{(j)\delta}$  and  $\boldsymbol{\varepsilon}^{(j)\delta} = \hat{\mathbf{C}} \mathbf{e}^{(j)\delta}$ . Third, generate the sequences  $\mathbf{X}_{t+h}^*$  and  $\mathbf{X}_{t+h}^\delta$  from the estimated model (1)-(3) by using the two sequences of residuals and compute the

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<sup>6</sup> In this paper the size of the shock  $\delta$  is calibrated such that the impact response of the *Ramey* variable is equal to one standard deviation of the *Ramey*'s structural shocks.

difference element by element.<sup>7</sup> This gives you one observation of the GIRF in (4) for the horizons  $t, t + 1, \dots, t + h$  when a shock  $\delta$  hits the system at time  $t$  conditional on the initial history  $\omega_{t-1}$ . Then, the conditional distribution of the GIRF are constructed by repeating these three steps for a new draw of reduce-form residuals  $\epsilon^{(j)*}$  and thus generating a new observation for GIRF in (4). Per each horizon  $h$ , median values of the GIRF and the corresponding confidence intervals are computed from the generated distributions.

In order to compute the GIRFs conditional on each regime all the initial histories observed in my sample are separated into recessionary histories and expansionary histories by looking at the values of the transition variable  $z$ .<sup>8</sup> Caggiano *et al.* (2015) show that the fiscal multipliers is statistically different across regimes only for initial histories belonging to deep recessions and strong expansions. Therefore to this study be meaningful, I compute GIRFs for extreme realizations of recessions and expansions present in the sample 1939:Q1-2013:Q4, in the sense that initial histories corresponding to  $z < -1.22\%$  (5<sup>th</sup> percentile) are selected for the recession regime,<sup>9</sup> and initial histories belonging to  $z > 1.33\%$  (95<sup>th</sup> percentile) are chosen for the expansion regime.<sup>10</sup> For more details about the computation of the GIRFs, see Appendix B.

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<sup>7</sup> Following Caggiano *et al.* (2015), each of the generated sequences of  $X_{t+h}$  accounts for the evolution of  $F(z_{t-1})$  by keeping track of the evolution of output, and therefore of  $z_{t-1}$ . Hence, in this way, the probability  $F(z_t)$  is endogenised. Notice that the transition variable is defined as  $z_t \equiv 1/6(\Delta y_t + \Delta y_{t-1} + \Delta y_{t-2} + \Delta y_{t-3} + \Delta y_{t-4} + \Delta y_{t-5})$ , hence the relationship between  $z_t$  and each of its components  $\Delta y_{t-i}, i = 0, 1, \dots, 6$  involves only non-stochastic elements. This stochastic singularity allows me to take into account the interaction between output in the vector  $X_t$  and evolution of  $z_t$  when computing the GIRFs after the estimation of the STVAR model.

<sup>8</sup> Per each given initial history  $z$  I compute 500 different bootstrapped realizations for the GIRFs in (4), and then store median realization. I repeat this step until 500 initial histories (drawn with replacement) corresponding to recessions (expansions) are considered. Finally, the distribution for the GIRFs are constructed by considering 500 median realizations. For a more detailed explanation about the algorithm employed to compute the GIRFs see the Appendix B.

<sup>9</sup> Hence, the GIRFs conditional on recessions are estimated by considering 15 initial histories corresponding to four recessionary events, the post WWII recession in 1945, the late 1950s' recession, the stagflation at the middle of the 1970s, and the 2008' crisis.

<sup>10</sup> One could also make use of the threshold value  $\bar{z} = -0.65\%$  derived from the calibration of  $\{\gamma, c\}$  to discriminate between recessions and expansions. Thus, initial histories with values of  $z$  lower (higher) than the threshold will be classified as a recession (expansion). Then, the set of recessionary (expansionary) initial histories will include recessions (expansions) of all magnitude so that every initial history will be taken into account when computing the GIRFs. As Caggiano *et al.* (2015) show, the U.S. fiscal multiplier estimated conditional on all the histories is not statistically different across regimes, making meaningless any further analysis based on them.

### 3.2 Ex-Post Normalization Factor and Fiscal Multipliers

To quantify the effect that an anticipated (news) government spending shock has on output I construct fiscal multipliers by employing the estimated GIRFs. But first, given that government spending and output enter in the VAR system in logs, the estimated elasticities must be transformed in order to convert percentage changes into dollars changes. The typical approach is to use a *constant* ex-post normalization factor based on the sample average of GDP over government spending,  $Y/G$  (both taken in levels). However, Ramey and Zubairy (2014) show that the  $Y/G$  ratio may exhibit large variations across the time. For example, in my sample 1939-2013 this ratio varies from 2 to 7 with a mean of 4.76. Therefore the use of a constant value of  $Y/G$  may lead to upward biased, or at least distorted, multipliers estimates. Differently, Ramey and Zubairy (2014) estimate fiscal multipliers by employing Jordà's (2005) Local Projection technique that allows to convert GDP and government spending changes to the same units *before* the estimation. In practice, they normalize the variables on the left-hand-side of the model by defining them as  $(Y_t - Y_{t-1})/Y_{t-1}$  and  $(G_t - G_{t-1})/Y_{t-1}$ , thus the coefficients from  $Y$  are in the same units as those from  $G$ . Despite its convenience in the computation of fiscal multipliers, this approach carries other issues that may distort the final estimates. First, given that  $Y_{t-1}$  is used to normalize both GDP and government spending,  $(Y_t - Y_{t-1})/Y_{t-1}$  and  $(G_t - G_{t-1})/Y_{t-1}$  are correlated. Second, the Local Projection method is not system-wide as the SVAR method, implying that when computing impulse responses there is not link between the responses at  $h$  and  $h + 1$ , being the estimates often erratic. Third, this method accounts for the evolution of the state of the economy by estimating an autoregressive process for  $Y_{t+h}$  and  $G_{t+h}$  for each horizon  $h$ . Hence, as the horizon increases, one loses observations from the end of the sample. Moreover the Local Projection method does not allow for endogenize the switch from one regime to other. Furthermore, this method tends to generate serially correlated residuals.

Therefore the present paper develops a new approach that allows for computing a time-varying ex-post normalization factor  $Y/G$  within the robustness of a structural VAR framework. The computation of a time-varying  $Y/G$  is possible thanks to the features of the GIRFs. As explained in the previous section, these nonlinear impulse responses are estimated by randomizing over the bootstrapped reduce-form residuals, and interacting with the structural VAR estimated coefficients and the observed data. Conditional on a shock  $\delta_t$ , an initial history  $\omega_{t-1}$  and a sequence of bootstrapped reduce-form residuals  $\epsilon^{(j)*}$ , the estimates  $\{\hat{\Pi}_R(L), \hat{\Pi}_E(L), \hat{\Sigma}\}$  are employed to generate two different sequences of the vector of endogenous variables  $X_{t+h}$ . Then each observation  $j$  for the GIRFs is computed as:

$$GIRF(h, \delta_t, \omega_{t-1})^{(j)} = X_{t+h}^{(j)}(\omega_{t-1})^\delta - X_{t+h}^{(j)}(\omega_{t-1})^* \quad (5)$$

The equation (5) can be decomposed for each of the endogenous variables included in the vector  $\mathbf{X}_t$ . Therefore, for the observation  $GIRF^{(j)}$ , the responses for government spending and GDP read as follows:

$$GIRF.G_t(h, \delta_t, \boldsymbol{\omega}_{t-1}) = \ln G_{t+h}(\boldsymbol{\omega}_{t-1})^\delta - \ln G_{t+h}(\boldsymbol{\omega}_{t-1})^* \quad (6)$$

$$GIRF.Y_t(h, \delta_t, \boldsymbol{\omega}_{t-1}) = \ln Y_{t+h}(\boldsymbol{\omega}_{t-1})^\delta - \ln Y_{t+h}(\boldsymbol{\omega}_{t-1})^* \quad (7)$$

Notice that, by property of the logarithmic function, the right-hand side of the equation (6) and (7) can be rewritten as:

$$\begin{aligned} \ln G_{t+h}^\delta(\boldsymbol{\omega}_{t-1}) - \ln G_{t+h}^*(\boldsymbol{\omega}_{t-1}) &\approx \frac{G_{t+h}^\delta - G_{t+h}^*}{G_{t+h}^*} \\ \ln Y_{t+h}^\delta(\boldsymbol{\omega}_{t-1}) - \ln Y_{t+h}^*(\boldsymbol{\omega}_{t-1}) &\approx \frac{Y_{t+h}^\delta - Y_{t+h}^*}{Y_{t+h}^*} \end{aligned}$$

Then, the multiplier in dollars changes can be computed in the following way:

$$M(h, \delta_t, \boldsymbol{\omega}_{t-1}) = \frac{Y_{t+h}^\delta - Y_{t+h}^*}{G_{t+h}^\delta - G_{t+h}^*} \approx \frac{GIRF.Y_t(h, \delta_t, \boldsymbol{\omega}_{t-1})}{GIRF.G_t(h, \delta_t, \boldsymbol{\omega}_{t-1})} \times \frac{Y_{t+h}^*}{G_{t+h}^*} \quad (8)$$

where Y/G normalization factor can be computed at each point in time  $t$  and for each horizon  $h$  as:

$$\frac{Y_{t+h}^*}{G_{t+h}^*} = \frac{\exp\{\ln Y_{t+h}^*(\boldsymbol{\omega}_{t-1})\}}{\exp\{\ln G_{t+h}^*(\boldsymbol{\omega}_{t-1})\}} \quad (9)$$

Equally, this approach may be interpreted as converting the response of government spending and GDP into dollars changes at each point in the time *after* the estimation. To the best of my knowledge, this methodological finesse has not been applied in the literature so far. Furthermore, I compute two measures of fiscal multiplier. First, I compute the max multiplier as the maximum response of output divided the maximum response of government spending. Secondly, I calculate the sum multiplier defined as the ratio of the sum of output response (to a spending shock) to the sum of government spending response (to a spending shock).<sup>11</sup> Both measures of fiscal multiplier are computed for five different horizons of 4, 8, 12, 16 and 20 quarters.

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<sup>11</sup> This last definition of the fiscal multiplier is often preferred by many in the literature given that it takes into account both, the persistence of a fiscal shock, and the evolution of the cost of government spending associated with the path of output (see, among others, Mountford and Uhlig, 2009; Uhlig, 2010; Fisher and Peters, 2010; Woodford, 2011; Ramey and Zubairy, 2014; Caggiano *et al.*, 2015).

### 3.3 GIRFs: Anticipated (News) Government Spending shock

This section presents the estimated GIRFs and fiscal multipliers for the baseline vector of endogenous variables  $X_t$  containing, in that order, the Ramey variable, the log of real per capita government spending, the log of real per capita taxes, the log of real per capita GDP, the log of CPI and the spread between the BAA and the AAA Moody's corporate bond rate. But first, I briefly analyze the effects of a government spending shock on a linear SVAR framework. Note from Figure 4 that when a government spending shock hits the system the reaction of the corporate bond spread is negligible and never statistically different from zero. Therefore, realize any further analysis about the role of the monetary policy by employing a linear SVAR would be meaningless. Next, I analyze the generalized impulse responses corresponding to the nonlinear STVAR model (1)-(3). Figure 5 shows the responses of the system to an anticipated government spending (news) shock for initial histories belonging to deep recessions and strong expansions. Observe that the response of output during recessions is so much larger than that corresponding to expansions. Moreover, CPI and the corporate bond spread positively and significantly react to an spending shock during recessions while during expansions the responses are almost never statistically different from zero. Importantly, the positive reaction of the corporate bond spread may suggest the existence of a crowding out effect during recessions. Table 1 contains the estimated fiscal multipliers (max and sum) for the linear case, for expansions and for recessions. Observe that, the fiscal multipliers are so much larger during recessions than over expansions. While the multiplier during recessions reaches values of 1.60 (max) and 1.92 (sum), the multiplier over expansions never takes values larger than one. Moreover the multipliers over recession are statistically larger than one during the first four quarters. Furthermore the linear fiscal multiplier (max and sum) is always larger than those corresponding to expansions but lower than those estimated for recessions. This indicates that the linear SVAR tends to average between the two states of the economy.

The above results suggest that, even though in presence of a possible crowding out effect, the size of fiscal spending multiplier is significantly larger during an economic slack than during a boost. However, it is not clear whether or not the multiplier is statistically different across regimes. Therefore in order to verify if the multiplier is state-dependent I run a test by computing the distribution of the difference between the multiplier estimated for recessions and that estimated for expansions.<sup>12</sup> Then, I plot the estimated distributions along with the corresponding confidence intervals. The aim of this exercise is to test if the

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<sup>12</sup> The empirical density of the difference between multipliers is obtained by subtracting a realization of the multiplier in expansions from a realization of the multiplier in recessions conditional on the same set of draws of reduce-form residuals as well as the same bootstrapped realizations of the matrices of dynamic coefficients and the corresponding covariance matrix. Moreover, the empirical densities are based on 500 realizations of such difference per each horizon.



difference in multipliers between regimes is statistically different from zero. Figure 6 depicts the distribution of the difference in multipliers for five different horizons. Observe that for the most of the cases the zero line lies outside the confidence intervals, therefore providing evidence in favor of state-dependent multipliers from the statistical standpoint. Moreover Figure 7 shows the multipliers for both states of the economy for all the horizons from 1 to 20 along with the corresponding confidence bands. Again, it is easy to note that the multipliers corresponding to recessions are statistically larger than those from expansions.

### 3.4 Systematic Response of Corporate Bond Spread

In order to study the role of the corporate bond spread in determining the effectiveness of a fiscal stimulus I perform a counterfactual exercise by computing the multipliers for the system  $X_t$  conditional to a fixed level of corporate bond spread. In doing so the responses of corporate bond spread to movements in system due to fiscal shocks is switched off by zeroing the coefficients of the corporate bond spread equation in the STVAR model.<sup>13</sup> Table 2 contains the counterfactual fiscal multipliers for recessions and expansions. Observe that now the multipliers during recessions are even larger than those corresponding to baseline case, reaching values of 2.14 (max) and 2.42 (sum). Moreover the estimated counterfactual multipliers during recessions are always statistically larger than one. Figure 8 shows the distribution of the difference between the multiplier estimated during recessions and that estimated over expansions for the counterfactual case. Furthermore Figure 9 depicts the multipliers conditional on a fixed level of corporate bond spread for both states and for all horizons from 1 to 20. Both figures clearly indicate that the counterfactual fiscal multipliers during recessions are always statistically larger than those corresponding to expansions.

Finally, to test whether during recessions the counterfactual fiscal multiplier is statistically larger than that corresponding to the baseline scenario, I compute the distribution of the difference between the counterfactual multiplier and the baseline multiplier during recessions.<sup>14</sup> Figure 10 depicts the distributions along with the corresponding confidence intervals. Observe that, this difference is always positive (about 0.5) and statistically different from zero, thus confirming that the counterfactual multiplier is statistically larger than baseline multiplier. These results indicate that the reaction of the corporate bond spread is

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<sup>13</sup> This approach has been employed by Sims and Zha (2006) and Caggiano, Castelnuovo and Nodari (2014) to study the effectiveness of monetary policy. Alternatively, one could also generate a sequence of hypothetical corporate bond spread shocks enough to keep the corporate bond spread fixed to its pre-shock level. I follow the zeroing coefficient approach to be in line up with the extant empirical literature in monetary policy.

<sup>14</sup> The differences of the multiplier in recessions for the counterfactual scenario *versus* the baseline scenario is computed conditional on the same set of draws of reduce-form residuals as well as the same bootstrapped realizations of the matrices of dynamic coefficients and the corresponding covariance matrix. The empirical densities are based on 500 realizations of such difference per each horizon.

relevant to determine the size of the fiscal multiplier, suggesting that a government spending shock triggers a mild crowding-out effect during recessions.

## 5 Conclusions

The present paper investigates the role of monetary policy in determining the effectiveness of spending-based fiscal stimulus during recessions and over expansions by jointly considering a fiscal stimulus and monetary variables within a nonlinear framework. To do so I quantify the size of the fiscal multiplier by employing a Smooth-Transition VAR (STVAR) model endowed with government spending, taxes, output, CPI and the corporate bond spread. Moreover in order to identify the fundamental government spending shocks I include the measure of spending news developed by Ramey (2011). Furthermore I compute the fiscal multipliers by estimating the generalized impulse response functions.

My results point out to a statistically larger fiscal multiplier during recessions than over expansions. Moreover the estimated fiscal multiplier during recessions is statistically larger than one. Furthermore, corporate bond spread positively reacts to a fiscal stimulus during recessions, therefore suggesting the existence of a mild crowding out effect. Importantly, I show that when the corporate bond spread is not allowed to react to government spending shocks during recessions, the fiscal multiplier is statistically larger than that corresponding to the baseline case. Unlike in recessions, the response of corporate bond spread does not turn out to be important for determining the size of the fiscal multiplier during expansions.

These findings suggest that movements in monetary variables matter to determine the size of the fiscal multiplier during downturns. A government spending shock may trigger an increase in the corporate bond spread, which in turn reduces the effectiveness of a fiscal stimulus during a recession.

## Appendix A- Evidence in Favor of Non-linearity

Following Teräsvirta and Yang (2014) I carry out a linearity test in order to check for nonlinear dynamics at a multivariate level. They develop a test for the null hypothesis of linearity against a Smooth Transition VAR with a single switching variable for the whole system.

Consider the  $p$  –dimensional  $n$ -order Taylor approximation around  $\gamma = 0$  of the STVAR model (1)-(3):

$$\mathbf{X}_t = \Theta_0' \mathbf{Y}_t + \sum_{i=1}^n \Theta_i' \mathbf{Y}_t z_t^i + \varepsilon_t \quad (\text{A1})$$

where  $\mathbf{X}_t = [Ramey_t \ g_t \ t_t \ y_t \ cpi_t \ spread_t]'$  is the  $(p \times 1)$  baseline specification of the vector of endogenous variables,  $\mathbf{Y}_t = [\mathbf{X}_{t-1}, \dots, \mathbf{X}_{t-k}, \boldsymbol{\alpha}]$  is the  $(k \times p + q) \times 1$  vector of exogenous variables including endogenous variables lagged  $k$  time and a column vector of constants  $\boldsymbol{\alpha}$ , and  $z_t$  is the transition variable. Moreover  $\Theta_0$  and  $\Theta_i$  are matrices of parameters. Following Teräsvirta and Yang (2014), the null hypothesis of linearity is  $H_0 : \Theta_i = \mathbf{0} \ \forall i$ . In the present paper I employ a Taylor approximation of order  $n = 1$ . Furthermore the number of exogenous variables is  $q = 1$ , the number of endogenous variables is  $p = 6$  and the number of lags is  $k = 1$  (this choice for the lag order is because the “course of dimensionality”, as indicated in Teräsvirta and Yang, 2014).

The test for linearity against the STVAR model is performed as follows:

- 1- Estimate the model under the null  $H_0 : \Theta_i = \mathbf{0} \ \forall i$  (estimate the linear model) by regressing  $\mathbf{X}_t$  on  $\mathbf{Y}_t$ . Compute the residuals  $\tilde{\mathbf{E}}$  and the matrix residuals sum of squares  $\mathbf{RSS}_0 = \tilde{\mathbf{E}}' \tilde{\mathbf{E}}$ .
- 2- Regress  $\tilde{\mathbf{E}}$  on  $\mathbf{Y}_t$  and  $\mathbf{Z}_n$  where  $\mathbf{Z}_n = [\mathbf{Y}_t' z_t | \mathbf{Y}_t' z_t^2 | \dots | \mathbf{Y}_t' z_t^n]$ . Compute the residuals  $\tilde{\tilde{\mathbf{E}}}$  and the matrix residuals sum of squares  $\mathbf{RSS}_1 = \tilde{\tilde{\mathbf{E}}}' \tilde{\tilde{\mathbf{E}}}$ .
- 3- Compute the test-statistic

$$\begin{aligned} LM_{\chi^2} &= Ttr\{\mathbf{RSS}_0^{-1}(\mathbf{RSS}_0 - \mathbf{RSS}_1)\} \\ &= T(p - tr\{\mathbf{RSS}_0^{-1} - \mathbf{RSS}_1\}) \end{aligned} \quad (\text{A2})$$

where  $tr\{\cdot\}$  indicates the trace of a matrix. Note that under the null hypothesis, the test statistic has an asymptotic  $\chi^2$  distribution with  $np(kp + q)$  degrees of freedom (42 in my case). The value of the test is  $LM = 279$  with a corresponding p-value equal to zero. Therefore, the null hypothesis of linearity is rejected in favour of a STVAR specification of the model. Importantly, the result of this test is robust to a Taylor approximation of order  $n = 2$  and  $n = 3$ .

## Appendix B- Generalized Impulse Response Functions

I compute the Generalized Impulse Response Functions as defined in (4) for the nonlinear VAR model (1)-(3) by following the approach proposed by Koop, Pesaran and Potter (1996). The algorithm consists of the following steps:<sup>15</sup>

1. Construct the set of all possible initial histories  $\Omega$  observed in the sample  $t = 1939Q1, \dots, 2013Q4$ :  $\{\omega_{t-1,i} \in \Omega\}$ , where  $\omega_{t-1,i} = \{X_{t-1}, \dots, X_{t-p}; z_{t-1}\}$  contains the lagged endogenous variables and the transition variable  $z_t$  lagged one period at a particular date  $t$ .
2. Separate the set of (deep) *recessionary histories*  $\Omega^R$  from that of (strong) *expansionary histories*  $\Omega^E$  by looking at the value of the transition variable. So, if  $z < -1.22\%$  (5<sup>th</sup> percentile), then  $\omega_{t-1,i} \in \Omega^R$ ; and if  $z > 1.33\%$  (95<sup>th</sup> percentile), then  $\omega_{t-1,i} \in \Omega^E$ .
3. Pick at random one initial history  $\omega_{t-1,i}$  from the set  $\Omega^R$ . Then draw randomly with replacement a sequence of  $h$  six-dimensional residuals  $\boldsymbol{\varepsilon}^{(j)*} = \{\boldsymbol{\varepsilon}_t, \boldsymbol{\varepsilon}_{t+1}, \dots, \boldsymbol{\varepsilon}_{t+h}\}$  from a Gaussian distribution  $N(\mathbf{0}, \hat{\boldsymbol{\Sigma}})$ , where  $\hat{\boldsymbol{\Sigma}}$  is the variance-covariance matrix obtained from the bootstrap distribution for the estimated parameters  $\{\hat{\boldsymbol{\Pi}}_R(L), \hat{\boldsymbol{\Pi}}_E(L)\}$  of the model (1)-(3).<sup>16</sup> Moreover  $h$  indicates the horizon of interest for the GIRF.
4. Orthogonalize the bootstrapped residuals to recover the structural shocks as  $\mathbf{e}^{(j)*} = \hat{\mathbf{C}}^{-1} \boldsymbol{\varepsilon}^{(j)*}$ , where  $\hat{\mathbf{C}}$  is the Cholesky factor of the variance-covariance matrix  $\hat{\boldsymbol{\Sigma}}$ .
5. Form another sequence of bootstrapped structural shocks  $\mathbf{e}^{(j)\delta}$  by replacing the observation corresponding to the *Ramey* shock in  $\mathbf{e}_t^{(j)*}$  with the perturbed shock  $e_{1t}^\delta = e_{1t}^* + \delta$ , with  $\delta > 0$ .

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<sup>15</sup> The algorithm here presented is similar to the one employed by Caggiano, Castelnuovo and Nodari (2014) and Caggiano *et al.* (2015).

<sup>16</sup> In order to account for parameter uncertainty, I construct the generalized impulse response functions for different draws of the coefficients of the vector obtained via bootstrap procedure. Hence, for each selected initial history  $\omega_{t-1,i}$ , a new set of coefficients  $\{\hat{\boldsymbol{\Pi}}_R(L), \hat{\boldsymbol{\Pi}}_E(L), \hat{\boldsymbol{\Sigma}}\}$  is draw at random from the empirical distribution of coefficients reflecting the parameter uncertainty. Moreover this empirical distribution of coefficients is based on 500 bootstrap replications.

6. Then transform back  $\mathbf{e}^{(j)*}$  and  $\mathbf{e}^{(j)\delta}$  to the residuals as  $\boldsymbol{\varepsilon}_t^{(j)*} = \widehat{\mathbf{C}}\mathbf{e}^{(j)*}$  and  $\boldsymbol{\varepsilon}_t^{(j)\delta} = \widehat{\mathbf{C}}\mathbf{e}^{(j)\delta}$ .

7. Conditional on  $\boldsymbol{\omega}_{t-1,i}$ , generate the evolution of  $\mathbf{X}_{t+h}^{(j)*}$  and  $\mathbf{X}_{t+h}^{(j)\delta}$  for the estimated model (1)-(3) by using the sequences of residuals  $\boldsymbol{\varepsilon}^{(j)*}$  and  $\boldsymbol{\varepsilon}^{(j)\delta}$ , respectively. Then compute the GIRF as:

$$GIRF(h, \delta_t, \boldsymbol{\omega}_{t-1,i})^{(j)} = \mathbf{X}_{t+h}^{(j)\delta} - \mathbf{X}_{t+h}^{(j)*} \quad (\text{A3})$$

8. Repeat step 7 for  $j = 1, \dots, J$  vectors of bootstrapped residuals, thus generating  $J$  different observations for  $GIRF(h, \delta_t, \boldsymbol{\omega}_{t-1,i})^{(j)}$ . Set  $J = 500$ .

9. Compute the GIRF by averaging across the different observations  $j$  as:

$$\widehat{GIRF}(h, \delta_t, \boldsymbol{\omega}_{t-1,i}) = J^{-1} \sum_{j=1}^J GIRF(h, \delta_t, \boldsymbol{\omega}_{t-1,i})^{(j)} \quad (\text{A4})$$

10. Repeat steps 3-9 for  $i = 1, \dots, I$  initial histories contained in the set of (deep) *recessionary histories*  $\boldsymbol{\omega}_{t-1,i} \in \boldsymbol{\Omega}^R$ , thus obtaining different observations for  $\widehat{GIRF}(h, \delta_t, \boldsymbol{\omega}_{t-1,i})^R$ , where the subscript  $R$  indicates that the  $\widehat{GIRF}$  are computed conditional upon *recessionary histories*. Set  $I = 500$ .

11. Compute the median GIRF under (deep) recessions  $\widehat{GIRF}(h, \delta_t, \boldsymbol{\Omega}^R)$  by taking the average across  $i$  as:

$$\widehat{GIRF}(h, \delta_t, \boldsymbol{\Omega}^R) = I^{-1} \sum_{i=1}^I \widehat{GIRF}(h, \delta_t, \boldsymbol{\omega}_{t-1,i}) \quad (\text{A5})$$

12. In order to compute the GIRF conditional upon expansions, repeat the previous steps 3-11 for 500 initial histories belonging to the set of (strong) *expansionary histories*  $\boldsymbol{\omega}_{t-1,i} \in \boldsymbol{\Omega}^E$ , and obtain  $\widehat{GIRF}(h, \delta_t, \boldsymbol{\Omega}^E)$ .

13. The 68% confidence bands are computed by taken the 14<sup>th</sup> and the 86<sup>th</sup> percentiles of the generated densities  $\widehat{GIRF}(h, \delta_t, \boldsymbol{\omega}_{t-1,[1:500]})^R$  and  $\widehat{GIRF}(h, \delta_t, \boldsymbol{\omega}_{t-1,[1:500]})^E$ .

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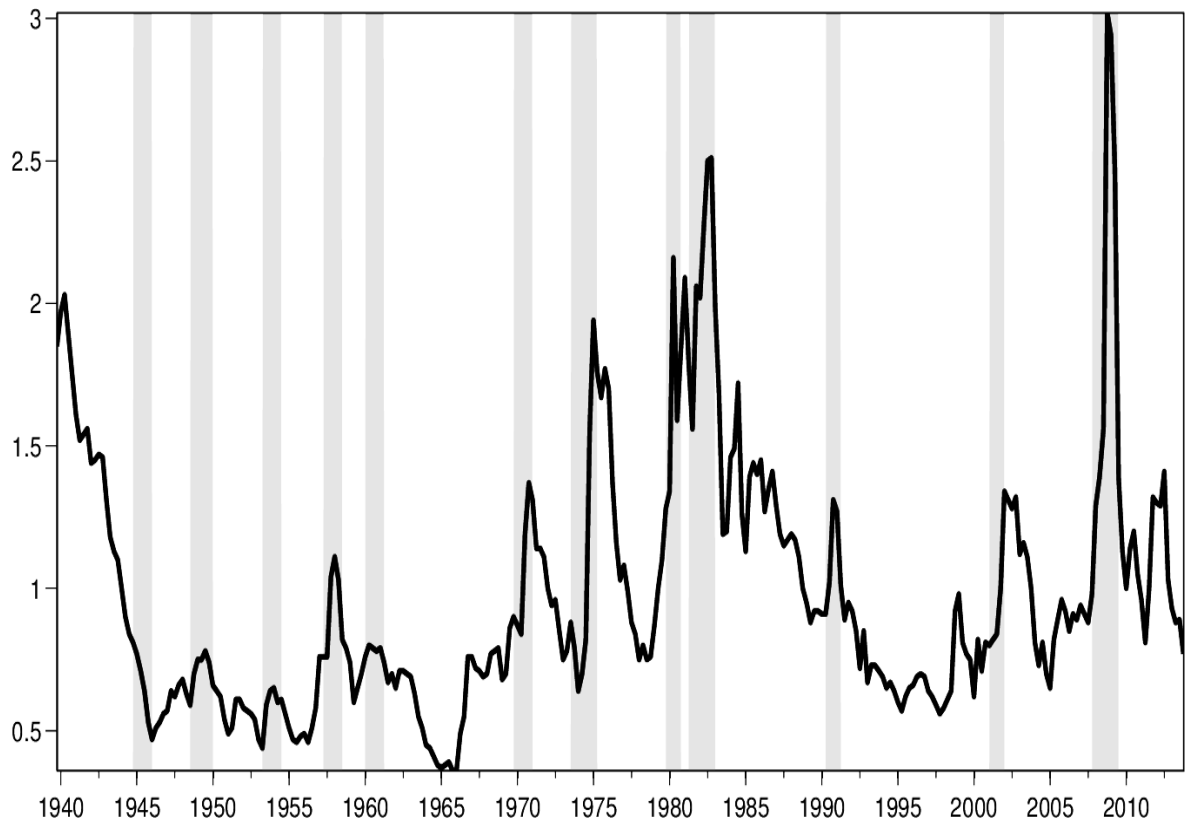


<i>Multiplier</i>	<i>Regime /horizon</i>	<i>Linear</i>	<i>Expansion</i>	<i>Recession</i>
<i>Max</i>	<i>h = 4</i>	0.79 [0.51 1.05]	0.26 [0.07 0.58]	1.60 [1.28 2.07]
	<i>h = 8</i>	0.79 [0.57 1.01]	0.43 [0.17 0.72]	1.26 [0.93 2.13]
	<i>h = 12</i>	0.80 [0.58 1.01]	0.47 [0.21 0.76]	1.24 [0.88 2.20]
	<i>h = 16</i>	0.80 [0.59 1.01]	0.48 [0.21 0.78]	1.22 [0.86 2.11]
	<i>h = 20</i>	0.80 [0.59 1.01]	0.49 [0.22 0.80]	1.20 [0.85 2.09]
	<i>Sum</i>	<i>h = 4</i>	1.40 [0.85 1.80]	0.03 [-1.15 1.07]
<i>Sum</i>	<i>h = 8</i>	0.91 [0.62 1.18]	0.32 [-0.17 0.76]	1.21 [0.89 2.16]
	<i>h = 12</i>	0.90 [0.60 1.18]	0.41 [-0.06 0.95]	1.01 [0.71 2.28]
	<i>h = 16</i>	0.93 [0.57 1.30]	0.49 [-0.05 1.19]	1.04 [0.70 2.26]
	<i>h = 20</i>	0.81 [0.51 1.47]	0.51 [-0.05 1.18]	1.11 [0.72 2.10]

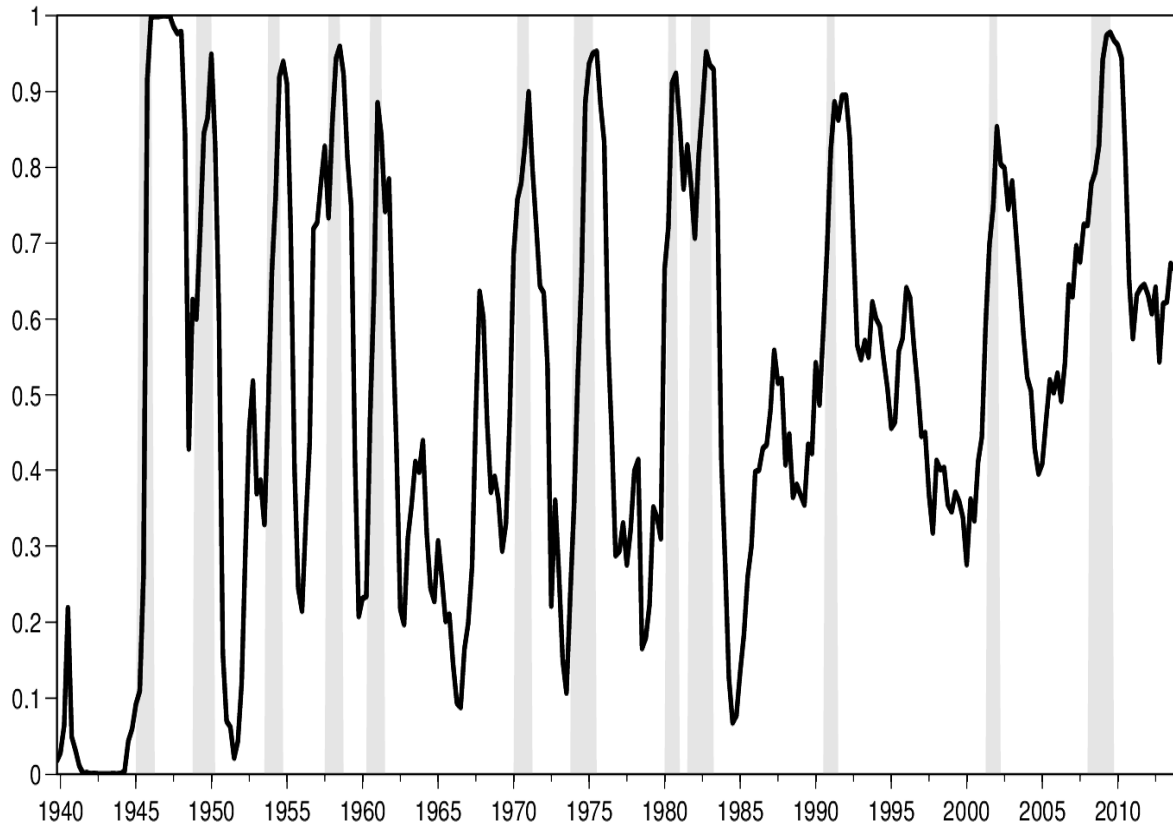
**Table 1. Fiscal Multiplier: Baseline.** Fiscal multipliers for the specification containing, in that order, the Ramey variable, the log of real per capita government spending, the log of real per capita taxes, the log of real per capita GDP, the log of CPI and the spread between the BAA and the AAA Moody's corporate bond rate. The shock is the first of the Cholesky decomposition. The third column shows the fiscal multipliers for a linear VAR, while the last two columns present the fiscal multipliers for the nonlinear STVAR model with GIRFs conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample. The estimated multipliers are scaled by the time-varying factor  $Y/G$  in order to transform percentage changes into dollars changes. The  $Y/G$  factor is computed at each point in the time (initial history) and for each horizon (quarters ahead after the shock). The numbers in brackets indicate the 68% confidence intervals.

<i>Multiplier</i>	<i>Regime /horizon</i>	<i>Expansion</i>	<i>Expansion w/o spread</i>	<i>Recession</i>	<i>Recession w/o spread</i>
<i>Max</i>	<i>h = 4</i>	0.26 [0.07 0.58]	0.23 [0.07 0.50]	1.60 [1.28 2.07]	2.14 [1.58 3.08]
	<i>h = 8</i>	0.43 [0.17 0.72]	0.44 [0.16 0.72]	1.26 [0.93 2.13]	1.77 [1.38 3.22]
	<i>h = 12</i>	0.47 [0.21 0.76]	0.50 [0.20 0.78]	1.24 [0.88 2.20]	1.75 [1.38 3.18]
	<i>h = 16</i>	0.48 [0.21 0.78]	0.51 [0.22 0.80]	1.22 [0.86 2.11]	1.75 [1.39 2.85]
	<i>h = 20</i>	0.49 [0.22 0.80]	0.53 [0.22 0.82]	1.20 [0.85 2.09]	1.76 [1.40 2.51]
<i>Sum</i>	<i>h = 4</i>	0.03 [-1.15 1.07]	-0.08 [-1.46 0.98]	1.92 [1.57 2.43]	2.42 [1.87 3.22]
	<i>h = 8</i>	0.32 [-0.17 0.76]	0.20 [-0.39 0.59]	1.21 [0.89 2.16]	1.82 [1.27 3.34]
	<i>h = 12</i>	0.41 [-0.06 0.95]	0.25 [-0.41 0.64]	1.01 [0.71 2.28]	1.72 [1.14 3.52]
	<i>h = 16</i>	0.49 [-0.05 1.19]	0.24 [-0.52 0.74]	1.04 [0.70 2.26]	1.77 [1.23 3.10]
	<i>h = 20</i>	0.51 [-0.05 1.18]	0.28 [-0.52 0.88]	1.11 [0.72 2.10]	1.80 [1.28 2.83]

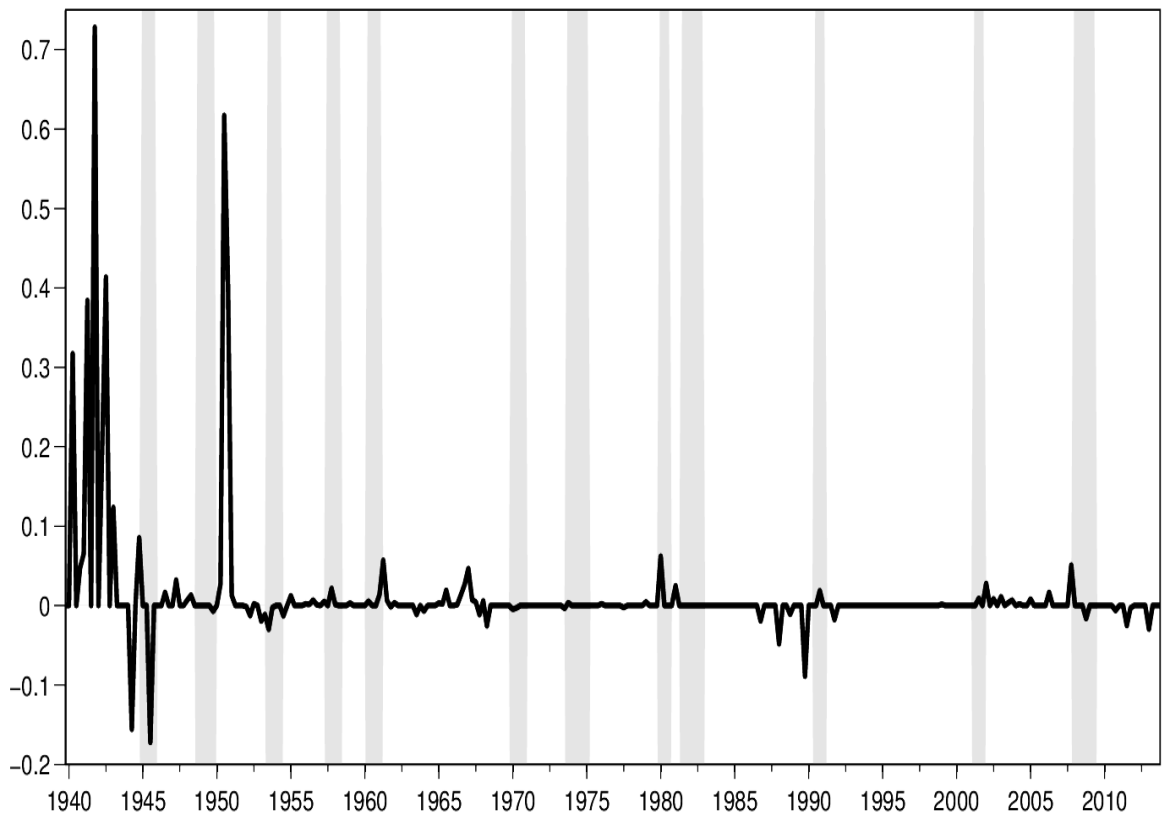
**Table 2. Fiscal Multiplier: Baseline vs. Counterfactual w/o spread.** Fiscal multipliers for the specification containing, in that order, the Ramey variable, the log of real per capita government spending, the log of real per capita taxes, the log of real per capita GDP, the log of CPI and the spread between the BAA and the AAA Moody’s corporate bond rate. The shock is the first of the Cholesky decomposition. Values for the nonlinear STVAR model with GIRFs conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample. The label “w/o spread” stands for the fiscal multipliers computed conditional on a constant level of spread bond rate, i.e., the responses of spread Baa-Aaa to movements in system due to fiscal shocks is switched off by zeroing the coefficients of the spread bond rate equation in the STVAR model. Moreover, in all the cases, the estimated multipliers are scaled by the time-varying factor Y/G in order to transform percentage changes into dollars changes. The Y/G factor is computed at each point in the time (initial history) and for each horizon (quarters ahead after the shock). The numbers in brackets indicate the 68% confidence intervals.



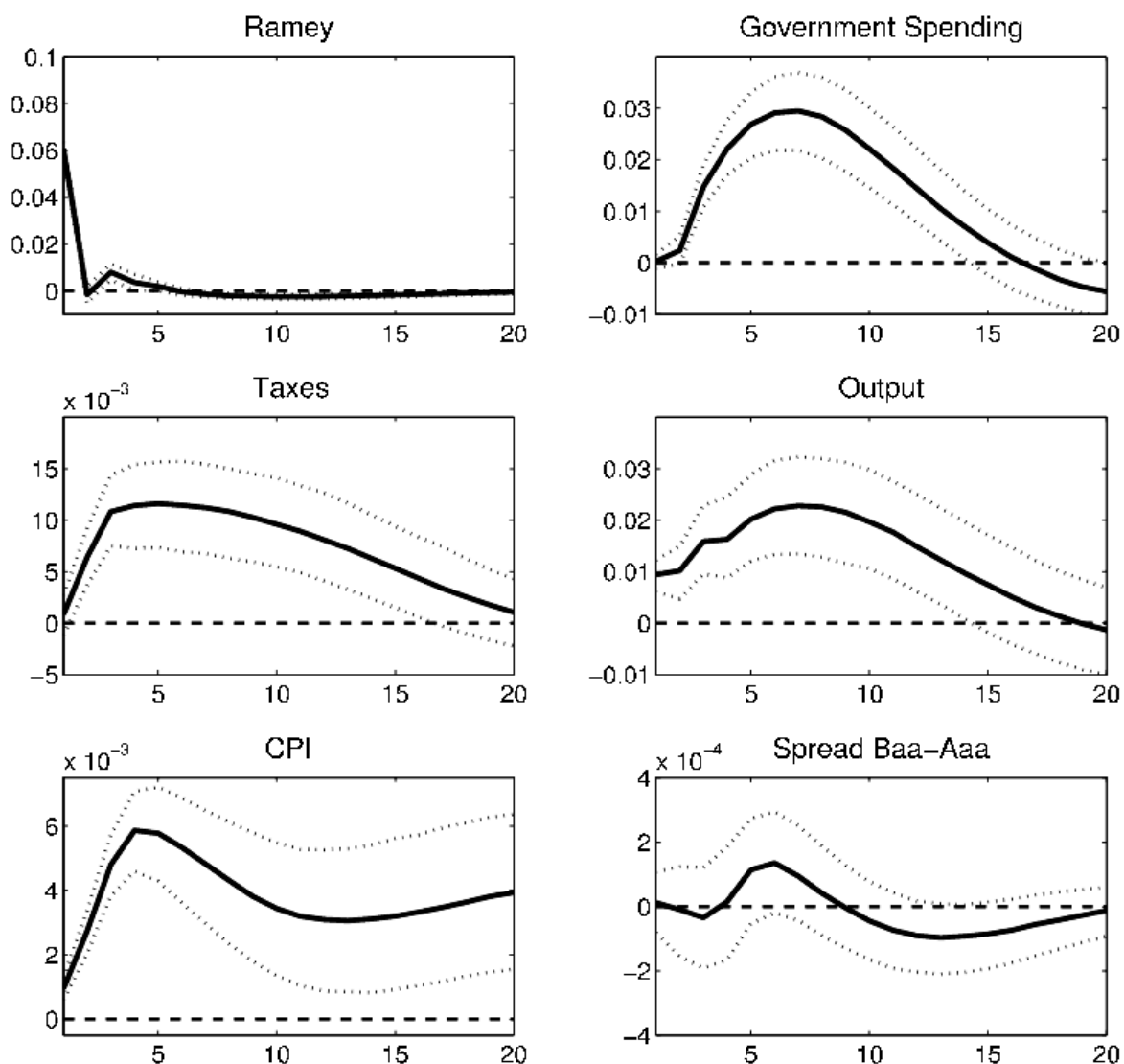
**Figure 1. Corporate Bond Spread and NBER Recessions.** The spread between the BAA and the AAA short run Moody’s corporate bond rate and the NBER recession dates. The black solid line draws the corporate bond spread while the shaded regions indicates the recessions defined by the NBER. Note that the corporate bond spread exhibits a countercyclical behavior.



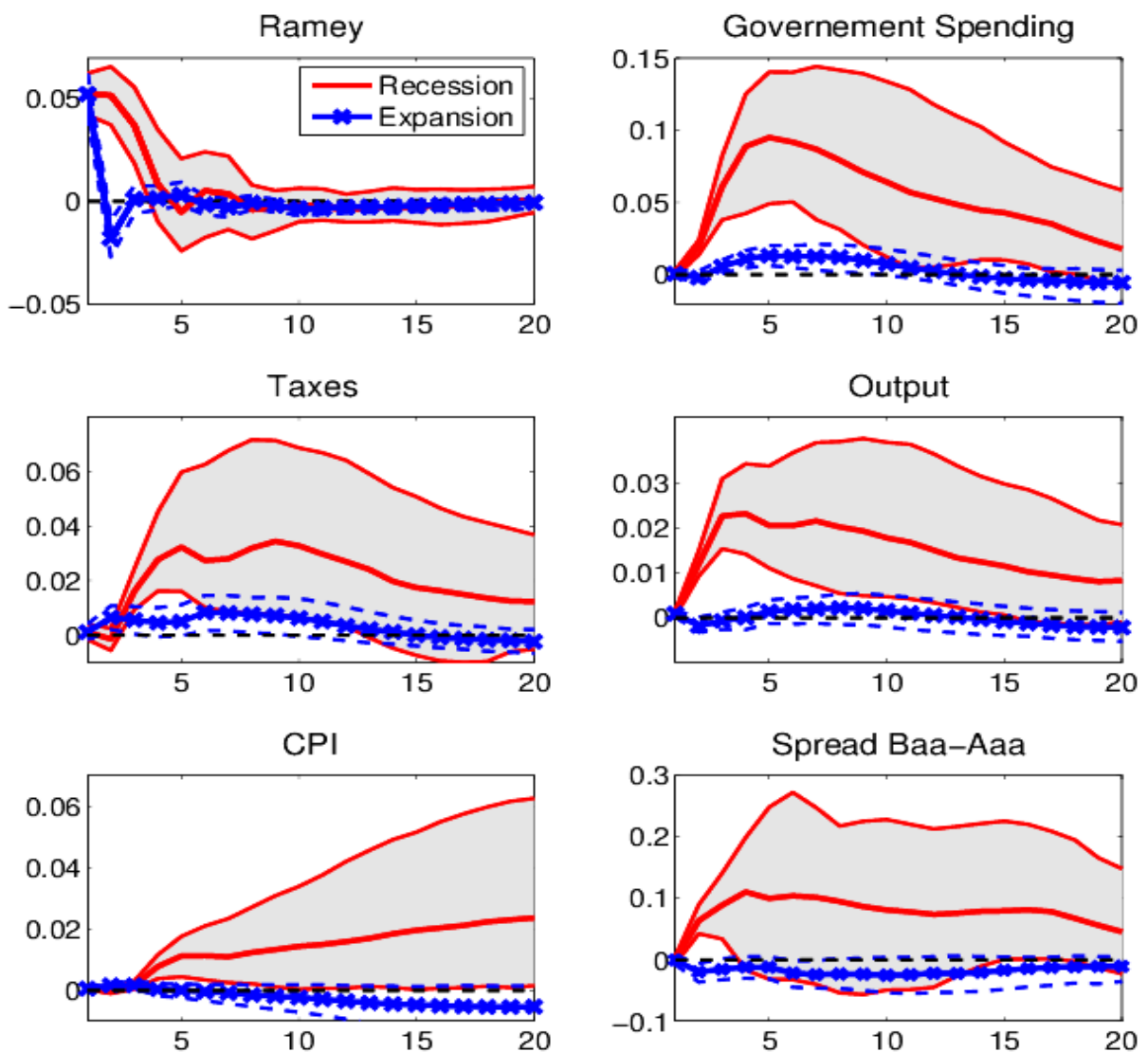
**Figure 2. Probability of being in a recession.**  $[1 - F(z_{t-1})]$  and the NBER recession dates. Note that the positive picks of the black solid line indicating the probability of being in the recessionary regime  $[1 - F(z_{t-1})]$  coincide with the shaded regions indicating the recessions defined by the NBER. The transition variable is computed as the standardized six quarters backward-looking moving average of real per capita GDP growth rate.



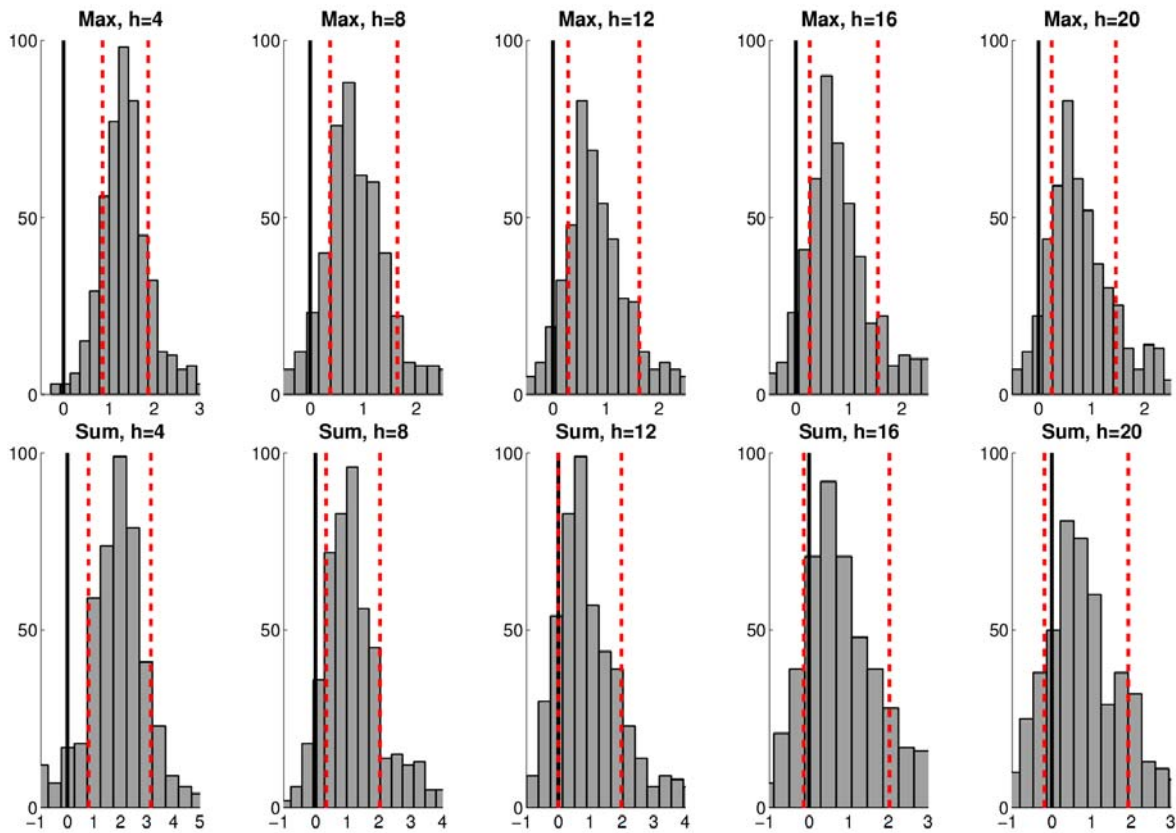
**Figure 3. Ramey Variable.** The black solid line draws the series for the Ramey variable while the shaded region indicate the recessions as defined by the NBER. The Ramey's variable is computed as the present value of the expected government expending changes due to foreign political events expressed as a percentage of the previous quarter GDP.



**Figure 4. IRFs to one standard deviation Anticipated (News) Government Spending Shock.** The black solid lines draw the median responses while the dotted lines depict the 68% confidence intervals. The shock is the first of the Cholesky decomposition for the specification including, in that order, the Ramey variable, the log of real per capita government spending, the log of real per capita taxes, the log of real per capita GDP, the log of CPI and the spread between the BAA and the AAA Moody's corporate bond rate. Moreover, the confidence intervals are computed as the 16th and 84th percentiles of the IRFs distributions obtained by bootstrap procedure with 1000 iterations.

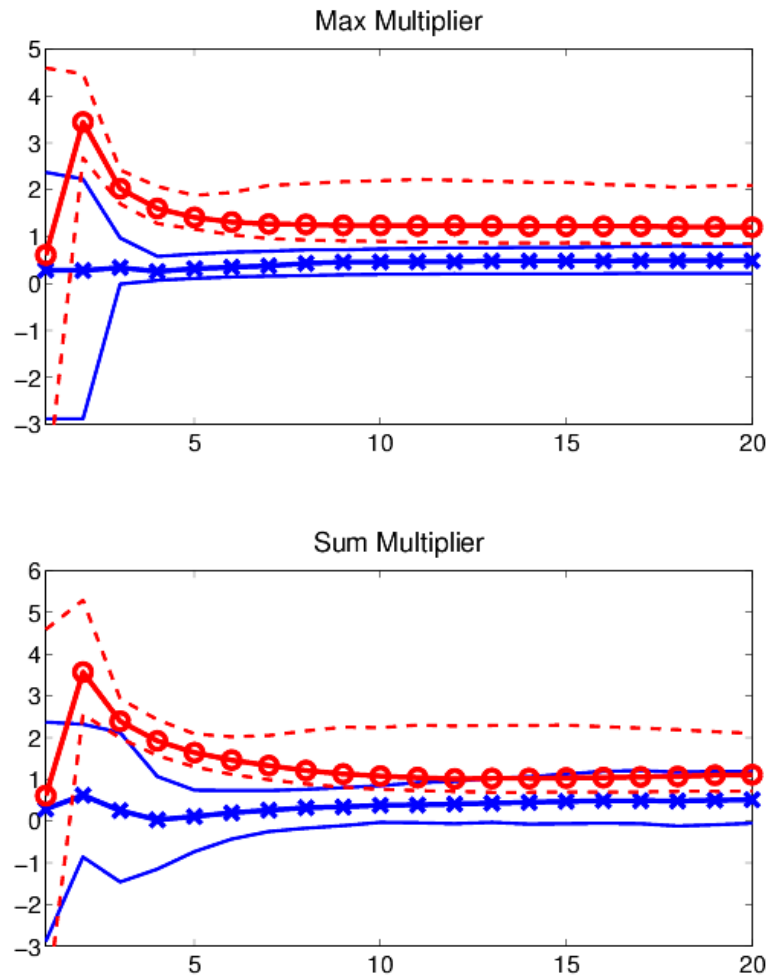


**Figure 5. GIRFs to an Anticipated (News) Government Spending Shock: Recession vs. Expansion.** The blue crossed lines draw the median responses of the variables during expansions while the red solid lines depict the median responses during recessions. The 68% confidence bands are shown by the blue dashed lines (expansions) and the shaded areas (recessions). The shock is the first of the Cholesky decomposition for the specification including, in that order, the Ramey variable, the log of real per capita government spending, the log of real per capita taxes, the log of real per capita GDP, the log of CPI and the spread between the BAA and the AAA Moody’s corporate bond rate. The GIRFs are computed conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample.

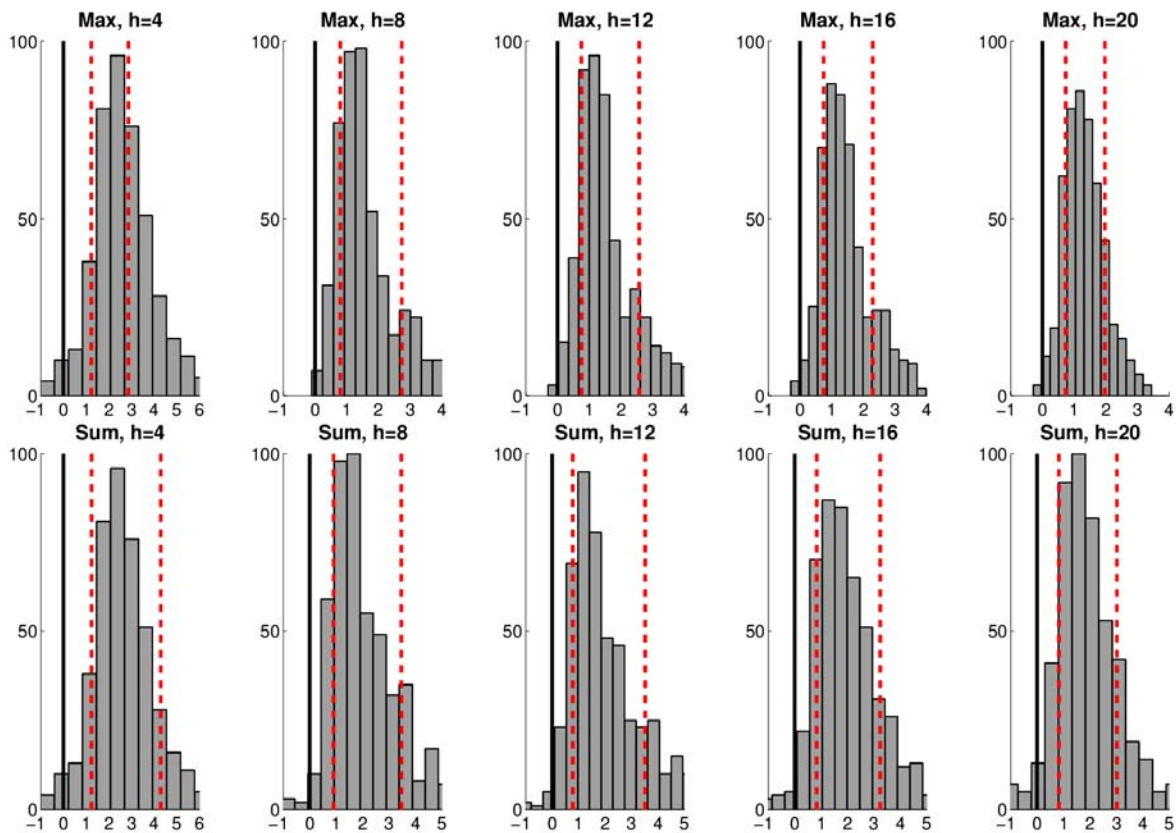


**Figure 6. Difference in Multipliers Between Expansions and Recessions.** The histograms depict the empirical densities of the difference in multipliers computed as multipliers in recession minus multipliers in expansions. The densities are constructed conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample. The multipliers are computed for the same draw of the stochastic elements of the STVAR model as well as the same draw of coefficients of the vector. The empirical densities are based on 500 realizations of such difference per each horizon. The red dashed lines represent 68% confidence intervals.

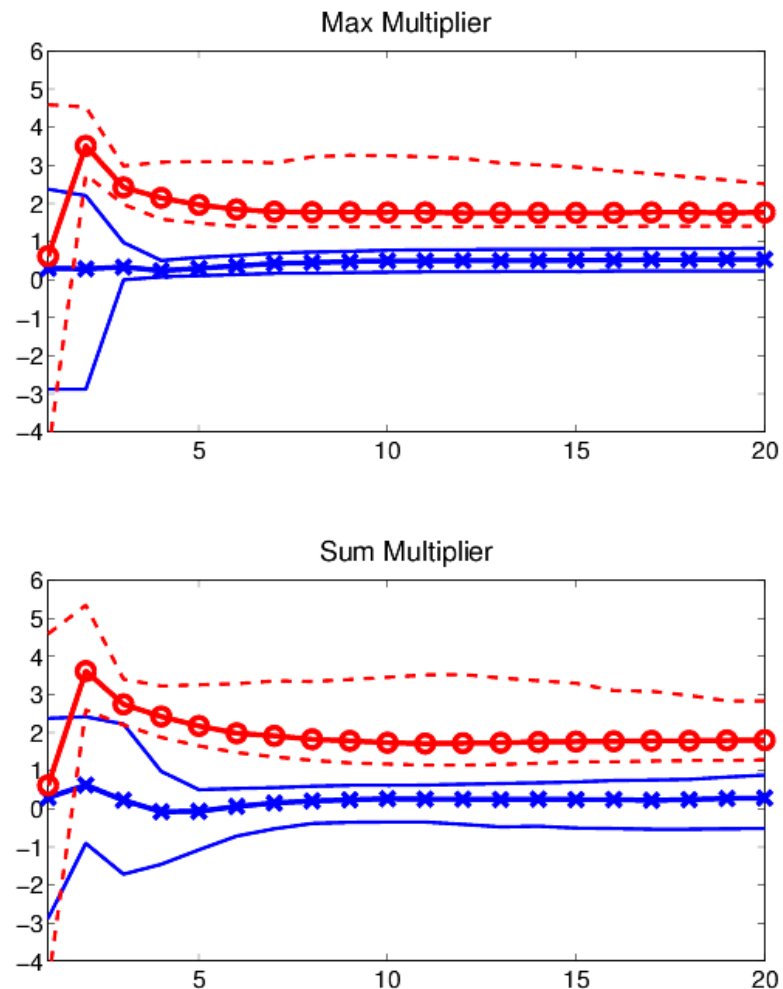




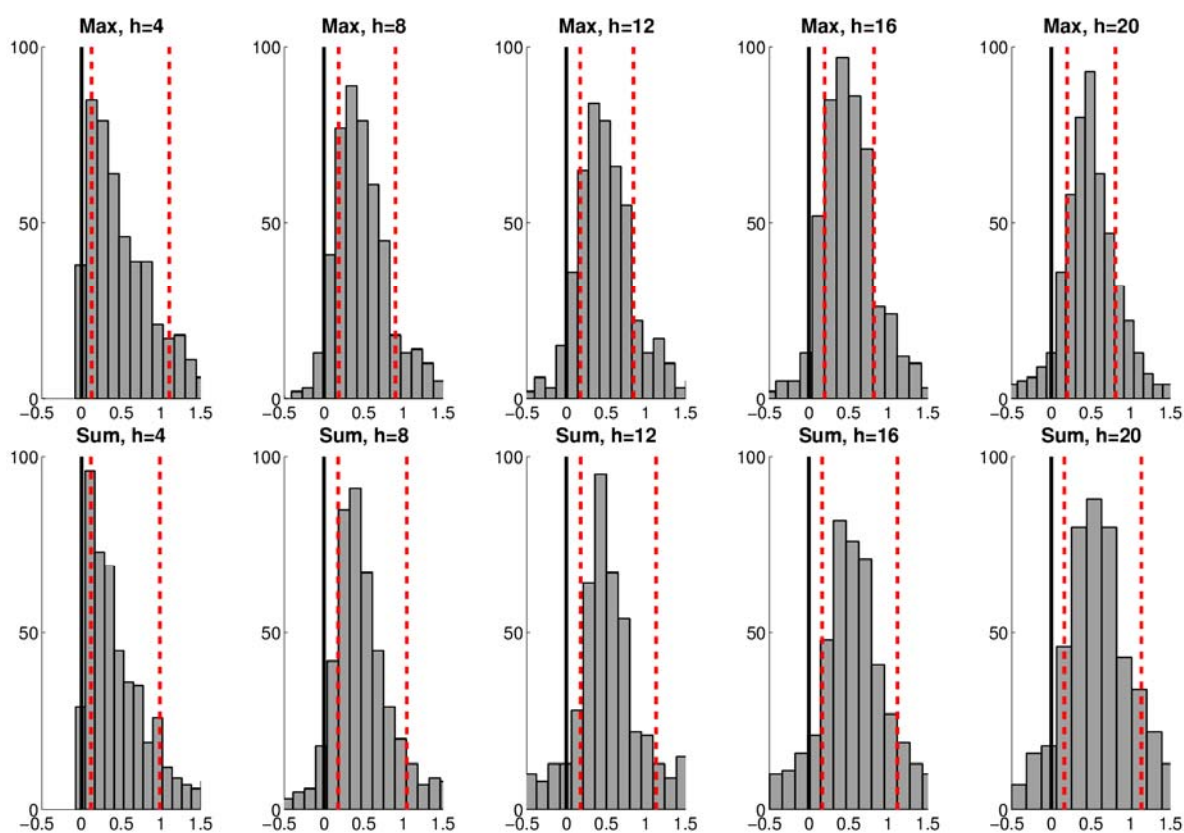
**Figure 7. Nonlinear Multipliers for All the Horizons.** The blue crossed lines draw the median multipliers for expansions while the red circled lines depict the median multipliers corresponding to recessions. The 68% confidence bands are shown by the blue solid lines (expansions) and the red dashed lines (recessions). Values for the nonlinear STVAR model with GIRFs conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample. The estimated multipliers are scaled by the time-varying factor  $Y/G$  in order to transform percentage changes into dollars changes. The  $Y/G$  factor is computed at each point in the time (initial history) and for each horizon (quarters ahead after the shock). The numbers in brackets indicate the 68% confidence intervals.



**Figure 8. Difference in Multipliers Between Expansions and Recessions: Counterfactual Case.** The histograms depict the empirical densities of the difference in counterfactual multipliers (i.e., conditional on a constant level of spread bond rate) computed as counterfactual multipliers in recession minus multipliers in expansions. The densities are constructed conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample. The multipliers are computed for the same draw of the stochastic elements of the STVAR model as well as the same draw of coefficients of the vector. The empirical densities are based on 500 realizations of such difference per each horizon. The red dashed lines represent 68% confidence intervals.



**Figure 9. Nonlinear Multipliers for All the Horizons: Counterfactual case.** Multipliers computed conditional on a constant level of corporate bond spread. The blue crossed lines draw the median multipliers for expansions while the red circled lines depict the median multipliers corresponding to recessions. The 68% confidence bands are shown by the blue solid lines (expansions) and the red dashed lines (recessions). Values for the nonlinear STVAR model with GIRFs conditional on initial histories belonging to extreme events (strong expansions and deep recessions) present in the sample. The estimated multipliers are scaled by the time-varying factor  $Y/G$  in order to transform percentage changes into dollars changes. The  $Y/G$  factor is computed at each point in the time (initial history) and for each horizon (quarters ahead after the shock). The numbers in brackets indicate the 68% confidence intervals.



**Figure 10. Difference in Multipliers Between Recession w/o Corporate Bond Spread and Recessions.** The histograms depict the empirical densities of the difference in multipliers computed as counterfactual multipliers in recessions minus baseline multipliers in recessions. The densities are constructed conditional on initial histories belonging to deep recessions present in the sample. The multipliers are computed for the same draw of the stochastic elements of the STVAR model as well as the same draw of coefficients of the vector. For the counterfactual case the responses of spread Baa-Aaa to movements in system due to fiscal shocks is switched off by zeroing the coefficients of the spread bond rate equation in the STVAR. The empirical densities are based on 500 realizations of such difference per each horizon. The red dashed lines represent 68% confidence intervals.

## **Chapter 3**

# **Economic Policy Uncertainty Spillovers in Booms and Busts**



# Economic Policy Uncertainty Spillovers in Booms and Busts \*

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## Abstract

We estimate a nonlinear VAR to quantify the impact of economic policy uncertainty shocks originating in the U.S. on the Canadian business cycle in booms and busts. We find strong evidence in favor of asymmetric spillover effects. Uncertainty shocks originating in the U.S. explain about 27% of the variance of the 2-year ahead forecast error of the Canadian unemployment rate in periods of slack vs. 8% during economic booms. Counterfactual simulations lead to the identification of a novel "economic policy uncertainty spillovers channel". According to this channel, spikes in U.S. economic policy uncertainty foster economic policy uncertainty in Canada in first place and, because of the latter, lead to a temporary increase in the Canadian unemployment rate. This channel is shown to work only in periods of slack.

*Keywords* : Economic Policy Uncertainty Shocks, Spillover Effects, Unemployment Dynamics, Smooth Transition Vector AutoRegressions, Recessions.

*JEL Codes* : C32, E32, E52.

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## 1 Introduction

*My view is that much of Canada's current economic policy uncertainty is due to contagion from the US. [...] Given the integrated and interdependent nature of the US and Canadian economies, this US-based economic policy uncertainty will continue to impede and adversely affect Canadian economic growth.*

Nicholas Bloom, Fraser Alert, February 2013, p. 2.

Is economic policy uncertainty a driver of the business cycle? Baker, Bloom, and Davis (2016) address this question by constructing a novel index of economic policy uncertainty for the U.S. and a number of other countries. When employing such index in carefully designed VAR-based analysis, they find that increases in the level of uncertainty associated to policy decisions can explain a non-negligible share of the business cycle in the U.S. and other industrialized countries. This result is important for two reasons. First, because it reaffirms that uncertainty can very well be one of the drivers of fluctuations in real activity in the United States, a result previously found by a number of authors (for recent surveys, see Bloom, Fernandez-Villaverde, and Schneider (2013) and Bloom (2014)).<sup>1</sup> Second, because it points to a particular type of uncertainty - the one connected to policy decisions - as an independent source of fluctuations in real activity.

Most of the theoretical and empirical literature on uncertainty has focused on autarkic frameworks to identify the effects of an uncertainty shock. While being a natural first-step to understand the macroeconomic effects of movements in uncertainty, this assumption appears to be questionable for small open-economies, which are naturally affected by shocks coming from neighboring countries and the rest of the world in general. A textbook example is Canada. It is well known that first-moment shocks - say, technology, monetary policy, or fiscal shocks - originating in the United States are able to explain a large fraction of the volatility of real activity in Canada (see, for instance, Schmitt-Grohe (1998), Justiniano and Preston (2010), and Faccini, Mumtaz, and Surico (2016)). However, to our knowledge, little is known on the spillover effects related to second moment shocks, and - in particular - economic policy uncertainty shocks.

This paper studies economic policy uncertainty spillovers. It does so by estimating a nonlinear Smooth-Transition VAR (STVAR) model in which economic policy uncertainty shocks originating in the U.S. are allowed (but not necessarily required) to act as drivers of real activity in Canada. The STVAR set up allows us to study the potentially asymmetric effects of external uncertainty shocks during phases of booms and busts of the Canadian

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<sup>1</sup> Uncertainty may very well be in part endogenous and due to a number of mechanisms (Bachmann and Moscarini (2012), Bachmann and Bayer (2013)). We discuss the endogeneity issue and the way in which we tackle it in the next Section.



business cycle. We model the effects of a spike in the U.S. EPU index on a number of Canadian macroeconomic variables, including real activity indicators (industrial production, unemployment), inflation, a short-term interest rate, and the bilateral real exchange rate connecting the U.S. and Canada. Importantly, we account for the possible transition from a state of the economy to another by computing Generalized Impulse Response Functions (GIRFs) à la Koop, Pesaran, and Potter (1996). This modeling choice implies that the probability of being in a given state of the business cycle is a fully endogenous object in our framework. This is important for our analysis, because a priori we would expect a potentially recessionary shock like a spike in uncertainty to force the Canadian economy to switch from, say, a boom to a bust. Our empirical model enables us to assess to what extent this is true as regards an EPU shocks realizing in the U.S. and spilling over in Canada.

We find statistically and economically relevant nonlinear spillover effects. An equally-sized economic policy uncertainty hike originating in the U.S. is estimated to trigger a strong and persistent downturn in Canada in the 1985-2014 period. The same shock, when occurring in booms, leads to quantitatively milder and mostly insignificant responses of real activity indicators. A forecast error variance decomposition exercise confirms that contagion via uncertainty shocks is a quantitatively more relevant phenomenon when Canada's growth rate is below trend. In particular, uncertainty shocks originating in the U.S. explain up to 27% of the variance of the 2-year ahead forecast error of the Canadian unemployment rate during slow-growth phases vs. about 8% during economic booms.

One of the variables reacting in a significant and persistent fashion to U.S. EPU shocks is the Canadian EPU index. We then analyze the role played by the evolution of the latter in the transmission of the external EPU shocks to the Canadian economy. We do so by conducting a counterfactual simulation which shuts down the response of the Canadian EPU index. The responses of the Canadian macroeconomic indicators turn out to be dramatically dampened, above all when the economy is slack. This result points to the existence of a novel "economic policy uncertainty spillover channel". Our reading of the transmission mechanism is that hikes in the level of the U.S. economic policy uncertainty foster the build up of EPU in Canada and, consequently, exert a negative effect on the Canadian business cycle.

Our search for asymmetric responses of real activity indicators, and unemployment in particular, is driven by a well-established theoretical and empirical literature. Chetty and Heckman (1986) show that exit costs lower than entry costs in a given industry may lead to fast drops in production and slow recoveries. Mortensen and Pissarides (1993) build up a model featuring job creation slower than job destruction due to search-related costs. This model delivers faster upward movements in unemployment than downward ones. Benigno and Ricci (2011) analytically show that downward wage rigidities imply a nonlinear aggregate supply curve which is vertical in presence of high inflation but flattens when inflation is low. Given that relationship between slack and low inflation, movements in aggregate demand caused by spikes in uncertainty may have larger real effects in periods of low growth.

Sichel (1993) proposes a test for deepness and steepness and find empirical support for both when working with the U.S. unemployment rate. Evidence pointing to an asymmetric behavior of the U.S. unemployment rate is also provided by, among others, Koop and Potter (1999), van Dijk, Teräsvirta,, and Franses (2002), Morley and Piger (2012), and Morley, Piger, and Tien (2013). Dibooglu and Enders (2001) find the Canadian unemployment rate to adjust nonlinearly to its long-run equilibrium. Moreover, unemployment tends to increase during economic downturns, which are phases in which uncertainty is typically found to substantially increase (Jurado, Ludvigson, and Ng, 2015; Bloom, Floetotto, Jaimovich, Saporta-Eksten, and Terry, 2014). Hence, the effects triggered by uncertainty shocks in recessions are likely to be different than those occurring in expansions. Recent evidence along this line is provided by, among others, Nodari (2014), Caggiano, Castelnuovo, and Groshenny (2014), and Caggiano, Castelnuovo, and Nodari (2015). Alessandri and Mumtaz (2014) show that the real effects of uncertainty shocks are larger in periods of financial distress. Our paper makes a further step along this path by showing economic policy uncertainty spillovers contribute to the asymmetric behavior of the Canadian unemployment rate.

The structure of the paper is the following. Section 2 makes contacts with other related literature. Section 3 presents our empirical set up in detail. In particular, it explains the construction of an U.S. EPU-related dummy which we use to identify spikes in economic policy uncertainty in the United States to facilitate the identification of EPU shocks in our empirical exercise. Then, it presents the Smooth-Transition VAR model we employ in our analysis. Section 4 presents the estimated dynamics responses of the Canadian economy to economic policy uncertainty spillovers coming from the United States. It also documents a list of robustness checks which confirm our baseline result. Section 5 analyzes further our empirical results and proposes empirical support in favor of an "international economic policy uncertainty spillover" channel. Section 6 concludes.

## **2 Other Related Literature**

Our paper joins three different but related strands of the literature on the role of uncertainty shocks. First, several authors have already studied the effects of economic policy uncertainty shocks. Baker, Bloom, and Davis (2016) develop country-specific indices of economic policy uncertainty. These indices are based on newspaper coverage frequency, and are shown by the authors to be closely related to movements in policy related economic uncertainty. In particular, the U.S. index is documented to peak near events like tight presidential elections, wars, 9/11, the failure of Lehman Brothers, and a number of battles over fiscal policy. The authors find that an upward movement in economic policy uncertainty leads to an increase in stock price volatility and a reduction in investment, output, and employment in the United States. A panel VAR modeling 12 major economies largely

confirms this result. Our paper builds on Baker, Bloom, and Davis' (2016) and employs their EPU indices for the U.S. and Canada to study the spillover effects of hikes in EPU uncertainty from the former country to the latter. As anticipated, we find evidence of stronger spillover effects when the Canadian economy is slack, particularly as regards the response of the unemployment rate. Working with a VAR model, Benati (2013) shows that economic policy uncertainty to be able to explain a fraction of the 1-year ahead forecast error variance of the U.S. industrial production growth rate of about 20-30%, and to be an important driver of real activity also for the Euro area, the United Kingdom, and Canada. Mumtaz and Surico (2013) use a VAR to model a number of indicators of fiscal stance and find fiscal policy uncertainty to be a relevant driver of the U.S. business cycle. Istrefi and Piloiu (2015) document a link between economic policy uncertainty and short- and long-run inflation expectations. Mumtaz and Theodoridis (2016) employ a flexible Factor Augmented VAR model with which they jointly estimate a measure of uncertainty and its time-varying impact on a number of variables. They find the relevance of uncertainty shocks in the United States to have declined over time as regards real and financial indicators, but not as regards inflation and a short-term interest rate. They interpret these findings through the lens of a nonlinear DSGE model which replicates their stylized facts via an increase in the anti-inflationary monetary policy stance and a flatter supply curve. Our contributions complement this literature by highlighting an international transmission channel which works asymmetrically along the business cycle in a small-open economy like Canada.

The second strand of the literature focuses on the role of uncertainty in an open economy context. Fernandez-Villaverde, Guerrón-Quintana, Rubio-Ramirez, and Uribe (2011) and Born and Pfeifer (2014) find changes in the volatility of the real interest rate at which small open emerging economies borrow to exert effects on real activity in open economies such as Argentina, Ecuador, Venezuela, and Brazil. Benigno, Benigno, and Nisticó (2012) find shocks to the volatility of monetary policy shocks, inflation target shocks, and productivity shocks realizing in the U.S. to be important drivers of a number of nominal and real indicators in the G7. They propose a general-equilibrium theory of exchange rate determination based on the interaction between monetary policy and uncertainty, and show that their theoretical model is able to replicate the stylized facts identified with their VARs. Working with a VAR framework, Mumtaz and Theodoridis (2015) estimate that a one standard deviation increase in the volatility of the shock to U.S. real GDP leads to a decline in U.K. GDP of 1% relative to trend and a 0.7% increase in U.K. CPI at the two-year horizon. They propose a model featuring sticky prices and wages delivering predictions in line with their stylized facts. Colombo (2013) studies the spillover effects of an economic policy uncertainty shock originating in the United States for the Euro area. She finds such shocks to be an important driver of the European policy rate. Carriere-Swallow and Cespedes (2013) study the impact of uncertainty shocks originating in the U.S. on a number of developed and developing countries. They find substantial heterogeneity in the response of investment and

consumption across countries. In particular, the response is more accentuated in developing countries, a stylized fact which the authors interpret in light of the different credit frictions affecting the functioning of financial markets in the countries under scrutiny. Gourio, Siemer, and Verdelhan (2013) build up a two-country RBC model in which aggregate uncertainty is time-varying and countries have heterogeneous exposures to a world aggregate shock. To test the empirical predictions of their framework, they construct a measure of international uncertainty by averaging up the volatility of equity returns of the G7 countries. They show that a shock to this measure of international uncertainty triggers a drop, rebound, and overshoot-type of response of industrial production in all these countries. Moreover, unemployment is also shown to respond to such shock. Cesa-Bianchi, Rebucci, and Pesaran (2014) employ a Global-VAR approach to study the effects of hikes in volatility on real activity for a number of industrialized and developing countries. They find the role of uncertainty shocks to be modest. Handley (2014) and Handley and Limao (2014, 2015) study the interconnections between policy uncertainty, trade, and real activity in a number of countries. They find policy uncertainty to be a key factor affecting trade and investment decisions. Similar conclusions are reached by Born, Muller, and Pfeifer (2013), who find that terms of trade uncertainty may be a relevant driver of real GDP in Chile. Our paper adds to this literature by unveiling the effects that economic policy uncertainty shocks originating in the U.S. exert as regards the Canadian business cycle. This result, which points to the relevance of external second moment shocks for a small open economy like Canada, complements previous contributions focusing on spillover effects from the U.S. to Canada due to first-moment shocks (see, for instance, Schmitt-Grohe, 1998; Justiniano and Preston 2010; and Faccini, Mumtaz, and Surico, 2016).

The third strand of the literature regards the effects of uncertainty shocks on real activity as predicted by micro-founded DSGE models. Gilchrist and Williams (2005) work with a standard real business cycle model featuring a Walrasian labor market. They show that uncertainty shocks are expansionary because, in their model, they exert a negative effect on households' wealth, increase the marginal utility of consumption and, therefore, labor supply, which eventually increases output. A different perspective is offered by Leduc and Liu (2015). They show that a labor market model featuring matching frictions predicts a negative impact on output by uncertainty shocks. This negative effect is related to an optimal "wait-and-see" strategy implemented by firms because of the lower expected value of filled vacancies in presence of uncertainty. This leads firms to post a lower number of vacancies, which leads to a lower number of matches on the labor market in equilibrium. Sticky prices are shown to magnify this effect due to the negative impact of uncertainty on aggregate demand and, consequently, on firms' relative prices, whose fall implies an even lower number of vacancies posted in equilibrium. Basu and Bundick (2014) also work with a model featuring sticky prices and show that their framework is able to replicate the conditional (on an uncertainty shock) comovements often found in the data. Back to RBC models, Bloom

(2009) show that a partial equilibrium framework modeling firms' decisions over labor and investment in presence of non-convex adjustment costs imply an optimal "wait-and- see" strategy which implies a drop in real activity after an uncertainty shock. When estimating his model with micro-data, he finds such costs to be empirically relevant, above all those related to changes in investment. Bloom, Floetotto, Jaimovich, Saporta-Eksten, and Terry (2014) augment Bloom's (2009) framework by modeling households' consumption-savings decisions. They show that the negative real activity effects to an uncertainty shock are still present when allowing households to pursue consumption smoothing. Our results support models predicting a drop in real activity after an uncertainty shock, and stress that this is particularly true when the economy is affected by an increase in uncertainty features unused capacity.

### **3 Modeling Asymmetric Spillover Effects: Shocks and Dynamics**

**U.S. EPU index and spikes in uncertainty.** As anticipated in the Introduction, Baker, Bloom, and Davis (2016) construct an index of economic policy uncertainty for the U.S. and a number of other industrialized countries. This index is based on newspaper coverage frequency. As regards the United States, Baker, Bloom, and Davis use two overlapping sets of newspapers. The first spans the 1900-1985 period and comprises the Wall Street Journal, the New York Times, the Washington Post, the Chicago Tribune, Los Angeles Times, and the Boston Globe. From 1985 until 2014, USA Today, the Miami Herald, the Dallas Morning Tribune, and the San Francisco Chronicle are added to the set. The authors perform month-by-month searches of each paper, starting in January of 1900, for terms related to economic and policy uncertainty. In particular, they search for articles containing the term "uncertainty" or "uncertain", the terms "economic", "economy", "business", "commerce", "industry", and "industrial", and the terms: "congress", "legislation", "white house", "regulation", "federal reserve", "deficit", "tariff", or "war". The article is included in the count if it includes terms in all three categories pertaining to uncertainty, the economy and policy. To deal with changing volumes of news articles for a given paper over time, Baker, Bloom, and Davis (2016) divide the raw counts of policy uncertainty articles by the total number of news articles containing terms regarding the economy or business in the paper. They then normalize each paper's series to unit standard deviation prior to December 2009 and sum each paper's series. Details are reported in Baker, Bloom, and Davis (2016).

We are interested in selecting realizations which are extreme and, therefore, likely to be informative as regards possible movements in the EPU index. We isolate spikes in uncertainty by selecting realizations of the Hodrick-Prescott filtered EPU index larger than 1.65 times its standard deviation. The smoothing weight of the Hodrick-Prescott filtered is set to 129,600 as suggested by Ravn and Uhlig (2002). This "event-study" approach follows the one adopted by Bloom (2009) to identify spikes in the U.S. stock market volatility.

Figure 1 plots the EPU index for the United States, along with the identified spikes in economic policy uncertainty. We give all these spikes an interpretation based on historical facts, which we report in Table 1. Some of these spikes relate to wars, the dissolution of the Soviet Union, and 9/11, which can be seen as external shocks. Some other spikes regard fiscal-or monetary-policy related events like discussions on the budget, the fiscal cliff, and huge monetary policy adjustments. These are shocks which we associate to domestic (U.S.) economic conditions. All these events have the potential to increase the uncertainty on how economic policy will operate in the future. Hence, at least in theory, they are all potentially important drivers behind decisions by agents in the economic system (firms and households) that eventually affect real activity, both domestically and in countries which are strictly interconnected to the United States, Canada in first place. We now turn to the description of the nonlinear framework we employ to achieve this purpose.

**STVAR model.** We allow for asymmetric spillover effects by modeling Canadian macroeconomic indicators with a Smooth-Transition VAR framework (for a reference textbook, see Teräsvirta, Tjøstheim, and Granger, 2010). Formally, our STVAR model reads as follows:

$$\mathbf{X}_t = [1 - F(z_{t-1})]\mathbf{\Pi}_R(L)\mathbf{X}_t + F(z_{t-1})\mathbf{\Pi}_E(L)\mathbf{X}_t + \boldsymbol{\varepsilon}_t \quad (1)$$

$$\boldsymbol{\varepsilon}_t \sim N(0, \boldsymbol{\Omega}), \quad (2)$$

$$F(z_t) = [1 + \exp\{-\gamma(z_t - c)\}]^{-1}, \gamma > 0, z_t \sim N(0, 1). \quad (3)$$

where  $\mathbf{X}_t$  is a set of endogenous variables which we aim to model,  $\mathbf{\Pi}_R$  and  $\mathbf{\Pi}_E$  are the VAR coefficients capturing the dynamics of the system during phases of slack and booms (respectively),  $\boldsymbol{\varepsilon}_t$  is the vector of reduced-form residuals having zero-mean and whose variance-covariance matrix is  $\boldsymbol{\Omega}$ ,  $F(z_{t-1})$  is a logistic transition function which captures the probability of being in a boom and whose smoothness parameter is  $\gamma$ ,  $z_t$  is a transition indicator, and  $c$  is the threshold parameter identifying the two regimes.

In brief, this model combines two linear VARs, one capturing the dynamics of the economy during busts and the other one during booms. Conditional on the transition indicator  $z_t$ , the logistic function  $F(z_t)$  indicates the probability of being in a boom. The transition from a regime to another is regulated by the smoothness parameter  $\gamma$ . Large values of this parameter imply abrupt switches from a regime to another, while moderate ones point to regimes of longer duration.<sup>2</sup>

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<sup>2</sup> Mumtaz and Theodoridis (2016) point to a different way of modelling the possibly evolving role played by uncertainty shocks with an application for the U.S. in which time-dependent responses are allowed to occur. A comparison between state- and time-dependent effects of economic policy uncertainty spillovers appear to be a promising avenue for future research.

A key choice for our empirical exercise is that of the transition indicator  $z_t$ . A standard choice in the literature is to consider a moving average of the growth rate of real GDP, which offers a good approximation of the ups and downs experienced by the U.S. business cycle (see, among others, Auerbach and Gorodnichenko, 2012; Bachmann and Sims, 2012; Caggiano, Castelnuovo, and Groshenny, 2014; Berger and Vavra, 2014; Nodari, 2014; Caggiano, Castelnuovo, Colombo, and Nodari, 2015; and Figueres, 2015). Our empirical exercise deals with monthly data to maximize the number of observations for the countries we study while retaining the possibility of studying the impact of EPU uncertainty shocks via the indices developed by Baker, Bloom, and Davis (2016) for the U.S. and Canada, and we employ a moving average of the growth rate of industrial production.<sup>3</sup> Conditional on our choice for  $z_t$ , we jointly estimate the parameters  $\{\mathbf{\Pi}_R(L), \mathbf{\Pi}_E(L), \mathbf{\Omega}, \gamma, c\}$  of model (1)-(3) with conditional maximum likelihood as suggested by Teräsvirta, Tjøstheim, and Granger (2010).<sup>4</sup>

**Modeled vector.** We model the Canadian economy with the following vector:  $\mathbf{X}_t = [EPU D_t^{US}, EPU_t, \overline{\Delta IP}_t, u_t, \pi_t, R_t, \Delta \epsilon_t]'$ . The variable  $EPU D_t^{US}$  is the dummy obtained by considering the spikes in the U.S. EPU uncertainty as described above. We anticipate here that our results are robust (and, in fact, reinforced) when the original EPU index is used *in lieu* of the dummy. All the remaining variables in the vector  $\mathbf{X}_t$  refer to the Canadian economy. In particular,  $\overline{\Delta IP}_t$  stands for the eighteen-term moving average of the monthly growth rate of industrial production (percentualized and annualized),  $u_t$  is the unemployment rate,  $\pi_t$  stands for CPI inflation (y-o-y percentualized growth rate of the monthly index),  $R_t$  is the monetary policy rate, while  $\Delta \epsilon_t \equiv \pi_t^{US} + \Delta s_t^{CANUS} - \pi_t^{CAN}$  is the growth rate of the bilateral real exchange rate between Canada and the U.S. constructed by considering the inflation rates in the two countries and combining it  $\Delta s_t^{CANUS}$ , which is the y-o-y growth rate of the Canada/US nominal exchange rate. All data were downloaded from the Federal Reserve Bank of St. Louis' website, with the exception of the EPU index, which was downloaded from the website <http://www.policyuncertainty.com/>. The Canadian EPU index is constructed by Baker, Bloom, and Davis (2016) by searching keyword terms such as "spending", "policy", "deficit", "budget", "tax", "regulation", and "central bank" in six

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<sup>3</sup> We employ a moving average of the month-by-month growth rate of industrial production involving eighteen terms. This moving average returns a correlation of -0.52 with the official dating of the Canadian recessions by the Economic Cycle Research Institute ( <https://www.ecri.org> ). Such correlation is higher (in absolute value) than the one found when considering the month-on-month growth rate of the Canadian industrial production (-0.23), a moving average involving six terms of the growth rate of industrial production (-0.47), and a moving average involving twelve terms (-0.50).

<sup>4</sup> Teräsvirta, Tjøstheim, and Granger (2010) point out that  $\gamma$  is not a scale-free parameter. To make it scale free, we follow their suggestion (p. 381 of their book) and standardize the transition indicator so that  $z_t$  takes a unitary standard deviation. This operation, along with the fact that we demean such indicator, makes our estimates more easily comparable with those present in the extant literature.

different newspapers, which are "the Gazette", "Globe and Mail", "Canadian Newswire", "Ottawa Citizen", "Toronto Star", and "Vancouver Sun".

We consider the sample 1985M1-2014M10. The beginning of the sample is dictated by the availability of the Canadian EPU index produced by Baker, Bloom, and Davis (2016), which we use here to make sure that spikes in the U.S. EPU index deliver information over and above the one delivered by abrupt changes in the Canadian one.<sup>5</sup> The end of the sample is justified by the end of the availability of the EPU historical index for the United States.<sup>6</sup>

**Test of linearity of the model.** We conduct a test in order to understand if a nonlinear framework provides us with a statistically better representation of the covariance structure of the data  $X_t$  than a standard linear multivariate framework. Teräsvirta and Yang (2014) propose a Lagrange Multiplier test of the null hypothesis of linearity vs. a specified nonlinear alternative that is exactly the logistic STVAR framework with a single transition variable. The Lagrange Multiplier statistic is 94.545, and the computed p-value approximately equal to zero clearly points to the rejection of the null hypothesis of linearity of the model. Details on this test are reported in our Appendix.

## EPU Spillovers: Empirical Evidence

We document our empirical findings starting with the estimated probability of slack according to our model. Then, we document the GIRFs of the Canadian macroeconomic indicators to an uncertainty shock coming from the United States. Finally, we document the robustness of our results to a variety of perturbations of the baseline framework.<sup>7</sup>

**Probability of being in a slack period.** Figure 2 plots the probability of being in a negative phase of the business cycle for Canada and contrasts it with the official 1990-92 and 2008-09 recessions as dated by the Economic Cycle Research Institute (ECRI).<sup>8</sup> The estimated logistic

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<sup>5</sup> In February 1991, the Bank of Canada officially adopted an inflation target. Our results are robust to the employment of the sample 1991M2-2014M10.

<sup>6</sup> To be precise, there are two U.S. EPU indices available at <http://www.policyuncertainty.com/>. One is the historical version of the EPU index, which is the one we use in our analysis. The other one is an index available since 1985 and constantly updated by the researchers behind the EPU project. This latter index measures policy-related economic uncertainty on the basis of three components, i.e., uncertainty as present in selected newspapers, federal tax code provisions set to expire in future years, and disagreement among economic forecasters as a proxy for uncertainty. Differently, the historical EPU index is constructed on the newspaper-component only. To preserve homogeneity and, at the same time, maximize the degrees of freedom of our exercise, we focus on the historical version of the EPU index.

<sup>7</sup> Estimates for the U.S. case point to asymmetric effects of EPU shocks for the U.S. unemployment rate. While representing a novel set of results, we decided to focus on the spillover effects from the U.S. to Canada. Our results for the U.S. case are documented in an Appendix available upon request.

<sup>8</sup> We are aware of two official datings of the business cycle for Canada. The first one is the one provided by the ECRI, and it is available at <https://www.businesscycle.com/ecri-business-cycles/international-business-cycle-dates-chronologies>. The second one is provided by the C.D. Howe Institute, and it is available



function for Canada turns out to be able to detect these recessions. The delay via which these two deep downturns are tracked is due to the backward-looking nature of the transition indicator we use. Conditional on our estimated  $\hat{c} = -0.72$ , our model classifies about 20% of the observations in the sample as recessions, a larger fraction than the 12% the ECRI classification suggests. This is mainly due to the fact that our logistic function also points to a deep downturn in the early 2000s, but this downturn is not an official recession.

The reason why our estimated logistic function indicates a high probability of slack in the early 2000s is the evolution of our transition indicator, i.e., the (standardized) 18-month growth rate of industrial production. The growth rate of industrial production experienced a dramatic fall between January 2000 and December 2001. In non-standardized terms, the 18-month growth rate fell from 13.6% to -8.3%. The magnitude of this fall is similar to the one recorded in correspondence of the two official recessions in our sample. This indicator of real activity fell from 12.5% to -7.1% in the May 1988-March 1991 period, and from 0.3% to -15.6% during the July 2008-May 2009 Great Recession phase. As shown in Figure 3, the evolution of the growth rate of industrial production in this sample mimics the one of the growth rate of the real GDP. Then, why were the early 2000s not officially classified as "recession"? The answer is that not all indicators of the business cycle pointed to a recession. A look at the Canadian unemployment rate helps us make this point. The unemployment rate went up from 6.8% to 8.1% from January 2000 to the end of 2001. The variation (difference between these two rates) reads 1.3%. Differently, the unemployment rate jumped from 7.8% to 10.5% in the 1988-1991 period (difference: 2.7%) and from 6.1% to 8.6% during the Global Financial Crisis (difference: 2.5%). Hence, while the early 1990s and the 2008-09 periods clearly featured strong and converging signals in favor of a recession, the early 2000s looked more like a severe downturn. In light of this evidence, our analysis should be interpreted as focusing on phases of growth of industrial production above vs. below the sample average, more than on official "expansions" and "recessions". However, it is of interest to count the number of U.S. EPU shocks hitting the Canadian economy in recessions and expansions. The number of U.S. EPU shocks hitting during official recessions in Canada is 9, while the number of shocks hitting during official expansions is 27. These figures are close to those related to the number of U.S. EPU shocks realizing in booms and busts according to our model, which is, 13 in busts and 23 in booms. Hence, the impact of the different classification of the early 2000s discussed above is likely to be moderate.

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here: <https://www.cdhowe.org/council/business-cycle-council>. While following slightly different procedures for the dating of the business cycle, these Institutes point to very similar datings of the Canadian business cycle. Our choice of the ECRI dating is due to internal consistency, in that such dating for the U.S. basically corresponds to the one provided by the NBER.

**GIRFs.** Figure 4 plots the impulse responses of a selected subset of Canadian macroeconomic variables to a one-standard deviation shock to the U.S. EPU dummy, as well as 68% confidence bands computed with the bootstrap-after-bootstrap strategy proposed by Kilian (1998). We focus in particular on unemployment and the growth rate of industrial production as real activity indicators, and inflation and the policy rate because of their policy-relevance. Several comments are worth making. First, there is significant evidence of a spillover effect going from the U.S. to Canada during busts. An unexpected hike in the U.S. economic policy uncertainty index triggers an increase in the Canadian unemployment rate, a decrease in industrial production, and a significant response of inflation and the policy rate. Second, the shape of the response of unemployment is similar in the two phases, but the quantitative response is very different, with unemployment responding more abruptly in recessions and remaining persistently high after the shock. Third, and differently from unemployment, industrial production displays an abrupt drop, and quick rebound, and a prolonged (but temporary) overshoot when the shock hit in recessions. This pattern is in line the one predicted, for real activity indicators, by Bloom's (2009) partial equilibrium model featuring non-convex adjustments labor and investment adjustment costs. Differently, the reaction of industrial production is insignificant when the shock hits in expansions. Fourth, the response of inflation is found to be different in the two states not only quantitatively but also qualitatively. The response of the growth rate of domestic CPI is negative, and persistently so, in periods of slack, a behavior consistent with a demand-driven interpretation of price formation. Viceversa, a positive short run reaction is detected when uncertainty hits during booms. This result may find its rationale in the behavior of firms operating in an environment facing price and wage stickiness. As pointed out by Mumtaz and Theodoridis (2015), firms in this environment may optimally decide to increase their prices to avoid getting stuck with "too costly" contracts, i.e., sub-optimally high real wages. Most likely, the different response of the inflation rate in the two states is the reason why the policy rate suggests a prolonged easing in recessions and a short-lived tightening in expansions. Importantly, as shown in Figure 5, most of these responses are also significantly different between states.

**Robustness checks.** We check the robustness of our baseline results along different dimensions.

*Alternative definitions of the U.S. uncertainty.* The results shown before rely on the use the dummy we constructed by isolating spikes in the U.S. EPU index. It is of interest to check if our baseline result is robust to the employment of two alternative indicators. First, we replace our dummy with the original U.S. EPU index by Baker, Bloom, and Davis (2016). This exercise is conducted to check if our dummy is driving our results. This exercise has an interesting by-product, which is, it allows us to understand if movements in uncertainty following the abrupt increases in the EPU index documented in Table 1 play a role in making the effects of uncertainty shocks more persistent. Second, we check the robustness of our

baseline findings by employing a different dummy which considers U.S.-related events only. As anticipated in the previous Section, some of the historical events associated to the peaks in uncertainty captured by the baseline version of our dummy are actually world-level shocks which are likely to have an influence also on the Canadian uncertainty index. Hence, we re-run our exercise by using an alternative dummy which excludes all the events which are obviously related to external elements (the most prominent example being wars). The selection of the dates is reported in Table 1, where we indicate those employed to construct this dummy as "U.S.-related".

Our results are plotted in Figure 6. Two results stand out. First, our qualitative and quantitative baseline result are clearly robust to the employment of the EPU index *per se* in our empirical model. Second, a comparison between our baseline GIRFs and those obtained with these two dummies points to a minor role of second round effects related to the evolution of the EPU index after an uncertainty shock. Indeed, the reaction of almost all the variables remain largely unchanged.

*Real GDP growth as transition indicator.* Our results are driven by our modeling choices, the one of the transition indicator included. While being a plausible indicator of the business cycle, the moving average of industrial production is clearly not the only indicator one may consider. In particular, a measure of real GDP at a monthly frequency is actually available for Canada.<sup>9</sup> We then use a moving average of the real GDP growth rate to replace industrial production in our VAR and conduct our empirical exercise. Figure 7 reports the comparison between our baseline impulse responses and those obtained with the real GDP growth rate. Our main results are clearly unchanged.

*Initial conditions to identify booms and busts.* Our baseline results are obtained by separating initial conditions (historical realizations of the lags of the variables we model with our nonlinear VAR) in two different groups, i.e., those indicating that the economy is in a boom and those that indicate that it is in a bust. Considering the logistic function (3), these initial conditions are technically associated to the transition indicator  $z_{t-1}$ , which per each given  $t$  is compared with the estimated threshold  $\hat{c}$ . In particular, values of  $z_{t-1} > \hat{c}$  ( $z_{t-1} < \hat{c}$ ) indicate that the economy is in a boom (bust). As in all nonlinear analysis of this kind, the risk of incorrectly classifying booms and busts is present, above all when initial conditions are associated to values of  $z_{t-1}$  close to the threshold. We then check the robustness of our results by dropping initial conditions associated to values of  $z_{t-1}$  which are "too close" to the threshold. Given that the transition indicator  $z_{t-1}$  is a standardized variable with unitary variance, we conduct two robustness checks so that initial conditions are considered only if  $|z_{t-1} - \hat{c}| > 1/\delta$ , with  $\delta \in \{1, 2\}$ . In line with Caggiano, Castelnuovo, Colombo, and Nodari (2015), who use this strategy to study the asymmetric effects of fiscal shocks, these robustness checks are basically based on the selection of "extreme"

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<sup>9</sup> See <http://www.statcan.gc.ca/start-debut-eng.html>. We used the Real GDP (2002 constant prices) series.

realizations of the business cycle (say, deep downturns or solid booms). When  $\delta = 2$ , about 10% (63%) of the observations in the sample are classified as recessions (expansions) according to our model, while when  $\delta = 1$ , our model classifies on about 5% (42%) of observations as recessionary (expansionary). Figure 8 depicts the impulse responses conditional on these sub-sets of initial conditions. Our GIRFs turn out to be robust to the exclusion of initial conditions related to more "tranquil" times. This suggests that, even in the baseline scenario in which no observation is discarded, the information discriminating between dynamics in booms vs. busts is actually the one related to the most extreme events.

*Financial market volatility.* The EPU index constructed by Baker, Bloom, and Davis (2016) captures economic policy-related spikes in uncertainty. Obviously, one concern related to our analysis is to what extent we are capturing effects coming from spikes in economic policy uncertainty as opposed to overall economic uncertainty. We then run an exercise by adding the S&P 100 Volatility index computed by the Chicago Board Options Exchange - known as the VXO index - at the top of our baseline VAR. The VXO index captures the evolution of the volatility of expected stock market returns, and it has been used since Bloom's (2009) contribution as a measure of broad economic uncertainty in applied macroeconomic investigations.<sup>10</sup> This exercise is conducted to control for a broader measure of economic activity, therefore isolating the contribution of the EPU shocks per se. Alternative measures of uncertainty are currently available, e.g., the one recently proposed by Jurado, Ludvigson, and Ng (2015) and based on a combination of real activity and financial indicators. Importantly, in a following paper, Ludvigson, Ma, and Ng (2016) employ the methodology proposed by Jurado, Ludvigson, and Ng (2015) to compute financial market uncertainty and real economic activity uncertainty separately. They find only the former to be a driver of the U.S. business cycle. Their estimate of the financial market uncertainty index conditional on a one-month horizon is highly correlated (0.84) with the VXO in our sample. We see this empirical fact as a validation of our choice to use the VXO as a proxy of a broader measure of uncertainty. As stressed by Stock and Watson (2012), uncertainty shocks and liquidity/financial risk shocks are highly correlated, which makes their separate interpretation problematic. The employment of the VXO is also an attempt to isolate the contribution of EPU with respect to financial shocks.<sup>11</sup> Figure 9 displays a comparison between the GIRFs computed with our baseline seven-variate nonlinear VAR and the eight-variate VAR featuring the VXO as first variable in the vector. In phases of slack, the response of unemployment is

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<sup>10</sup> A close measure is the S&P 500 Volatility index computed by the Chicago Board Options Exchange, which is known as the VIX. The correlation between the VIX and the VXO at a monthly frequency in the sample January 1990 (first month of availability of the VIX)-October 2014 is 0.99. We prefer to work with the VXO because it goes back in time to January 1986.

<sup>11</sup> For contributions aiming at separating uncertainty and financial shocks, see Christiano, Motto, and Rostagno (2014), Furlanetto, Ravazzolo, and Sarferaz (2014), and Caldara, Fuentes-Albero, Gilchrist, and Zakrajsek (2016).

now dampened, with a peak of about 0.5%, which is about half the value suggested by the baseline case. Still, this response is statistically significant and different with respect to the one we get in economic booms (confidence intervals not shown here for the sake of clarity of the figure, but available upon request). Going back to busts, the response of industrial production is shorter-lived, and the evidence of overshoot is now milder. In spite of a barely changed response of inflation, the reaction of the short-term interest rate is dampened too. Interestingly, the GIRFs related to booms seem to be unaffected by the introduction of the VXO, something which is consistent with the idea that the effects of financial market-related uncertainty shocks are particularly strong during downturns (Caggiano, Castelnuovo, and Groshenny (2014), Caggiano, Castelnuovo, and Nodari (2015)).

## 5 EPU Shocks: Contribution and Transmission Mechanism

The results documented so far speak in favor of the fact that variations in the U.S. EPU index can be associated to fluctuations in real activity, inflation, a short-term interest rate, and the real exchange rate in Canada. But how strong is this relationship? And what is the transmission mechanism, really? We answer these questions by considering, in turn, the results coming from a forecast error variance decomposition (FEVD) analysis and from a counterfactual exercise aiming at isolating the role of the Canadian EPU for the transmission of U.S. EPU shocks to the Canadian economy.

**FEVD.** We conduct the forecast error variance decomposition analysis by implementing the algorithm by Lanne and Nyberg (2016), who propose a generalized version of the forecast error variance decomposition for multivariate nonlinear models. Table 2 collects the figures related to the forecast error variance decomposition analysis conditional to a 24-month horizon.<sup>12</sup> A number of considerations are in order. First, as shown by the first row of the Table, jumps in the U.S. EPU shocks explain up to 27% of the volatility of the Canadian unemployment rate in the short-run. This figure points to EPU spillovers as being as important as domestic EPU shocks, the latter explaining about 24% of the Canadian unemployment rate. Hence, EPU spillover effects are actually quantitatively important if one aims at understanding the dynamics of a key labor market variable such as the unemployment rate. Moreover, uncertainty is important in general, given that it is responsible for about 51% of the variation in unemployment at a 2-year horizon. Second, the role of uncertainty is relevant in recessions only. Indeed, these figures dramatically drop to 8% (U.S. EPU shocks) and 2% (Canadian EPU shocks) when it comes to explaining unemployment during booming phases of the Canadian business cycle. A similar result holds

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<sup>12</sup> A FEVD analysis focusing on a 12-month horizon delivers very similar results, which are available upon request. Our FEVD analysis is conducted by considering the U.S. EPU index produced by Bloom, Baker, and Davis (2016) instead of our dummy. This is done to maximize the comparability between the figures related to the U.S. EPU shock and those related to the Canadian one.

true as regards industrial production, with uncertainty shocks explaining about 8% (U.S. EPU) and 14% (Canadian EPU) in busts, and about 1% and 3% in booms. The contribution of external economic policy uncertainty shocks to the volatility of inflation, the short-term interest rate, and the bilateral real exchange rate reads, respectively, 12%, 15%, and 13% in busts while it ranges from 3% to 5% in booms. Again, independently of the state of the economy, these figures are found to be fairly in line with the contribution of the Canadian EPU shocks.

Another interesting result of our FEVD analysis regards the drivers of the EPU indices employed in our analysis. As reported in Table 2, about 65% of the volatility of the U.S. EPU index in busts is driven by its own innovation, while the contribution of the innovation to the Canadian counterpart of this index is about 7%. This latter innovation explains an even lower share of the U.S. EPU index in booms (about 2%), which are phases in which about 74% of the volatility U.S. EPU index is driven by its own shock.<sup>13</sup> Differently, the contribution of U.S. EPU innovations to the volatility of the Canadian EPU index is 34% in busts and 32% in booms. This information is consistent with Granger causality tests conducted with a linear bivariate framework modeling the two EPU indices.<sup>14</sup> These tests point to the rejection of the null of causality running from the Canadian EPU index to the U.S. one (p-value: 0.00) and to the impossibility of rejecting, at standard confidence levels, the causality running from the U.S. EPU index to the Canadian one (p-value: 0.36). This result supports a novel reading of the role of big countries like the U.S. as regards the dynamics of small neighboring countries like Canada. Small open economies like Canada can be affected not only via the already well-known effects related to first-moment shocks like variations in technology or changes in macroeconomic policies, but also via a novel contagion channel which hinges upon second moments.

It is of interest to compare the contribution of uncertainty shocks to those of monetary policy shocks. Table 2 clearly point to a much smaller role played by monetary policy shocks as regards unemployment, with a contribution of about 5% during downturns (one fifth of external uncertainty shocks') and about 2% in booms (vs. 8% by U.S. EPU shocks'). The contribution of monetary policy shocks to the volatility of inflation reads 12% (no matter what the state of the business cycle is), and it is larger than that of uncertainty shocks, above all during expansions. Interestingly, the overall contribution of uncertainty shocks to the dynamics of the real exchange rate in busts is about 33%, much larger than the 4% due to monetary policy shocks. This gap is much smaller in booms, with the former shocks being responsible for about 7% of the variance of the real exchange rate against a contribution of

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<sup>13</sup> Notice that here we are referring to the volatility of the EPU indices, not to that of the innovations to such indices. Such innovations, which are those we use to compute the GIRFs documented in the previous Section and the FEVD reported in this Section, are exogenous under the assumption of our VAR being rich enough from an informational standpoint.

<sup>14</sup> We model a linear VAR(6) as suggested by the Akaike lag-length criterion.

about 2% by monetary policy shocks. Finally, and not surprisingly, the main driver of the short-term interest rate is monetary policy shocks. All in all, our results clearly point to uncertainty shocks (both external and domestic) as relevant drivers of the Canadian business cycle, at least when compared to monetary policy disturbances.

**Transmission mechanism.** The results of our FEVD analysis point to the possibility of an "international EPU spillover channel" linking the United States and Canada. In particular, one can conjecture the former country to be a big player whose economic policy uncertainty may lead neighboring countries like Canada to record subsequent increases in domestic uncertainty.<sup>15</sup> Interestingly, a simple regression modeling the Canadian uncertainty index with a constant and only one lag of the U.S. EPU index points to a far from negligible ability by the latter to predict the former, with an adjusted R<sup>2</sup> reading 0.33. Hence, it may very well be that fluctuations in uncertainty occurring in the U.S. foster uncertainty in Canada, at least the one perceived by readers of Canadian newspapers. This conjecture is confirmed by the impulse response of the Canadian economic policy uncertainty to a shock to the U.S. EPU index. Figure 10 plots the response of the baseline case, which features the U.S. EPU dummy, and the one in which our empirical model embeds the U.S. EPU index a la Baker, Bloom, and Davis (2016). The Canadian index significantly responds to external shocks in both states, and it does so in a persistent fashion. When the U.S. EPU index is modeled, the response of the Canadian one is even larger in magnitude and more persistent, possibly because of second-round effects going from the endogenous component of the measure of economic policy uncertainty in the U.S. to the Canadian one. These responses point to a direct "economic policy uncertainty spillover" channel linking the source of the shock, i.e., the United States, to the country receiving it, i.e., Canada. As documented above, another fact is that, after a shock to the level of U.S. economic policy uncertainty, Canada experiences temporary negative realizations of real activity. One possible way to interpret these facts is that spikes in uncertainty in the U.S. exert a contemporaneous impact on a number of variables in Canada, Canadian uncertainty included. Another interpretation is that spikes in the U.S. level of uncertainty affect the level of Canadian economic policy uncertainty in first place and, because of that, they affect real activity.

We shed light on the role played by the Canadian EPU index *per se* by conducting a counterfactual scenario in which the Canadian EPU index does not react to U.S. EPU shocks.<sup>16</sup> If the economic policy uncertainty actually perceived and considered by the Canadian households and firms is the Canadian one, and not the U.S. one *per se*, what this counterfactual should produce is more moderate responses of the Canadian macroeconomic

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<sup>15</sup> Given its interconnections with the United States, a country which would offer relevant information to validate this hypothesis is Mexico. Unfortunately, no EPU index for Mexico has been produced to date.

<sup>16</sup> The Canadian EPU index is maintained fixed by creating fictitious shocks to this index which offset the effects due to the U.S. EPU shocks on such index. Similar results were obtained by setting to zero the VAR coefficients in the equation that describes the Canadian EPU index.

indicators to an U.S. EPU shocks with respect to the baseline ones.<sup>17</sup>

Figure 11 shows that this is indeed the case, first and foremost for the unemployment rate. In recessions, this variable displays a quantitatively negligible (and statistically insignificant) response in the counterfactual scenario in which the Canadian EPU index remains still. Industrial production drops quickly, it immediately rebounds after the drop, and it displays an insignificant overshoot. The nominal interest rate displays an insignificant response, signalling a short-run negative reaction. The behavior of inflation is only mildly affected by the muted response of the Canadian uncertainty. Intriguingly, a different picture emerges in expansions, where no major differences emerge in terms of impulse responses when the Canadian EPU is not allowed to react to the external uncertainty shock. This is consistent with our FEVD analysis, which shows that EPU shocks in general, and the Canadian one in particular, are drivers of second-order importance for the Canadian business cycle when the economy experiences periods of sustained growth.

In order to test if the responses depicted in Figure 11 are statistically different we compute the distribution of the difference between the baseline responses and those of our counterfactual scenario. Figure 12 plots such distributions along with 68% confidence intervals. In line with what commented above, almost all these differences are significant, being much larger during recessions.

Finally, for completeness, Figure 13 depicts the responses for the baseline and counterfactual scenario along with 68% confidence bands. As reported above, the counterfactual responses are hardly significant.

## 6 Conclusions

We investigate the spillover effects of a jump in U.S. economic policy uncertainty as regards the Canadian macroeconomic environment. Using a nonlinear (Smooth-Transition) VAR, we find that such effects are present, significant, and asymmetric over the Canadian business cycle. In particular, our empirical model points to a strong evidence of spillovers in recessions, and a much more moderate one in expansions. The macroeconomic responses in these two states are found to be different from a statistical and economic standpoint. Counterfactual simulations conducted by freezing the response of the Canadian economic policy uncertainty index lead to results pointing to the existence of an "economic policy uncertainty spillover channel", i.e., spikes in U.S. economic policy uncertainty are likely to foster uncertainty in Canada and, therefore, lead to a temporary slowdown of the latter country's real activity.

Our results suggest that much of the Canadian economic policy uncertainty is actually due to contagion from the United States, as recently conjectured by Bloom (2013). On top of

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<sup>17</sup> To be consistent with our FEVD analysis the counterfactual exercise is conducted by employing the U.S. EPU index instead of our dummy. Similar results were obtained when considering the EPU dummy.



unveiling a novel transmission channel at an international level, our empirical findings support the conclusions of previous empirical studies documenting the asymmetric behavior of the unemployment rate along the business cycle (Koop and Potter, 1999; Dibooglu and Enders, 2001; van Dijk, Teräsvirta, and Franses, 2002; Morley and Piger, 2012; and Morley, Piger, and Tien, 2013), and to contributions pointing to the asymmetric business cycle effects of uncertainty shocks (Alessandri and Mumtaz, 2014; Nodari, 2014; Caggiano, Castelnuovo, and Groshenny, 2014; and Caggiano, Castelnuovo, and Nodari, 2015).

From a policy perspective, our evidence suggests that the uncertain policy actions in influential countries like the U.S. may not only be costly for such countries but also negatively affect neighboring small-open economies like Canada. As discussed by Davis (2015), the large increase in the number of norms and regulations that the U.S. economy has experienced for several years now is likely to have increased the level of policy-related uncertainty. Davis (2015) and Baker, Bloom, and Davis (2016) call for a clear, simple, and easy to administer regulatory system, a simple tax system, and predictable, timely, and clearly communicated policies. Thinking of the advantages of having economically sound commercial partners, our results suggest that the pay-off for the U.S. of implementing the policies suggested by Davis (2015) and Baker, Bloom, and Davis (2016) may be larger than those typically estimated when considering the U.S. case in isolation.

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## Appendix- Test of linearity of the model

In order to detect for nonlinear dynamics at a multivariate level, we apply the linearity test proposed by Teräsvirta and Yang (2014). This analysis consists in testing the null hypothesis of linearity against a Smooth Transition VAR with a single transition variable.

Consider the  $p$  –dimensional  $n$ -order Taylor approximation around  $\gamma = 0$  of the STVAR model (1)-(3):

$$\mathbf{X}_t = \boldsymbol{\Theta}'_0 \mathbf{Y}_t + \sum_{i=1}^n \boldsymbol{\Theta}'_i \mathbf{Y}_t z_t^i + \boldsymbol{\varepsilon}_t \quad (4)$$

where  $\mathbf{X}_t = [EPU D_t^{US}, EPU_t, \overline{\Delta IP}_t, u_t, \pi_t, R_t, \Delta \varepsilon_t]'$  is the  $(p \times 1)$  baseline specification of the vector of endogenous variables,  $\mathbf{Y}_t = [\mathbf{X}_{t-1}, \dots, \mathbf{X}_{t-k}, \boldsymbol{\alpha}]$  is the  $(k \times p + q) \times 1$  vector of exogenous variables including endogenous variables lagged  $k$  time and a column vector of constants  $\boldsymbol{\alpha}$ , and  $z_t$  is the transition indicator. Moreover  $\boldsymbol{\Theta}_0$  and  $\boldsymbol{\Theta}_i$  are matrices of parameters. Under the null hypothesis of linearity  $\boldsymbol{\Theta}_i = \mathbf{0} \forall i$ . The number of exogenous variables is  $q = 1$ , the number of endogenous variables is  $p = 7$  and the number of lags is  $k = 1$  (this choice for the lag order is because the “course of dimensionality”, as indicated in Teräsvirta and Yang, 2014). Furthermore we fix the value of the order of the Taylor approximation to  $n=1$ .

The test for linearity versus the STVAR model is performed as follows:

- 1- Estimate the model under the null  $H_0 : \boldsymbol{\Theta}_i = \mathbf{0} \forall i$  (estimate the linear model) by regressing  $\mathbf{X}_t$  on  $\mathbf{Y}_t$ . Compute the residuals  $\tilde{\mathbf{E}}$  and the matrix residuals sum of squares  $\mathbf{RSS}_0 = \tilde{\mathbf{E}}' \tilde{\mathbf{E}}$ .
- 2- Regress  $\tilde{\mathbf{E}}$  on  $\mathbf{Y}_t$  and  $\mathbf{Z}_n$  where  $\mathbf{Z}_n = [\mathbf{Y}'_t z_t | \mathbf{Y}'_t z_t^2 | \dots | \mathbf{Y}'_t z_t^n]$ .<sup>18</sup> Compute the residuals  $\tilde{\tilde{\mathbf{E}}}$  and the matrix residuals sum of squares  $\mathbf{RSS}_1 = \tilde{\tilde{\mathbf{E}}}' \tilde{\tilde{\mathbf{E}}}$ .
- 3- Compute the test-statistic

$$\begin{aligned} LM_{\chi^2} &= T \text{tr}\{\mathbf{RSS}_0^{-1}(\mathbf{RSS}_0 - \mathbf{RSS}_1)\} \\ &= T(p - \text{tr}\{\mathbf{RSS}_0^{-1} - \mathbf{RSS}_1\}) \end{aligned} \quad (5)$$

where  $\text{tr}\{\cdot\}$  indicates the trace of a matrix. Note that under the null hypothesis, the test statistic has an asymptotic  $\chi^2$  distribution with a number of degrees of freedom equal to  $p$  multiplied by the column dimension of  $\mathbf{Z}_n$  (49 in my case). The value of the test is  $LM = 94.545$  with a corresponding p-value approximately equal to zero. Therefore, the null hypothesis of linearity is rejected in favour of a STVAR specification of the model.

<sup>18</sup> Given that  $\mathbf{Y}_t$  contains a vector of constants  $\boldsymbol{\alpha}$  and our transition indicator  $z_t$  is the standardized  $\overline{\Delta IP}_t$ , we exclude the vector  $\boldsymbol{\alpha}$  from  $\mathbf{Z}_n$  in order to avoid perfect collinearity.

<b>Date</b>	<b>Event</b>	<b>Classification</b>
Jan. 1986	Balance budget act	U.S.-related
Oct. 1987	Black Monday	U.S.-related
Sep. 1990	Pres. Bush's speech on the military intervention in Kuwait	External
Jan. 1991	Gulf War I	External
Dec. 1991	Dissolution of the Soviet Union	External
Feb. 1992	Pres. Bush meets with Russian Pres. Yeltsin at Camp David	External
Dec. 1992	Clinton election	U.S.-related
Sep. 1998	Russian, LTCM default	External
Nov. 2000	Bush election	U.S.-related
Sep. 2011	9/11	U.S.-related
Jan. 2003	Gulf War II	External
Mar. 2003	Iraq invasion	External
Jan. 2008	Large interest rate cuts	U.S.-related
Sep. 2008	Lehman Brothers' bankruptcy	U.S.-related
Jan. 2009	Banking crisis	U.S.-related
July 2010	Mid-term elections	U.S.-related
Sep. 2010	Mid-term elections	U.S.-related
July 2011	Debt Ceiling	U.S.-related
Dec. 2011	Debt Ceiling	U.S.-related
Nov. 2012	Fiscal cliff	U.S.-related
Oct. 2013	Government shutdown	U.S.-related

**Table 1. Major U.S. Economic Policy Uncertainty Realizations.** Spikes identified as realizations exceeding the value 1.65 times the standard deviation of the Hodrick- Prescott filtered version of the U.S. Economic Policy Uncertainty index developed by Baker, Bloom, and Davis (2015). Smoothing weight of the Hodrick-Prescott filter set to 129,600. The label "External" refers to shocks whose origin can be assigned to an event external to both Canada and the U.S., and which is therefore in common. The label "U.S.-related" refers to shocks whose origin can be evidently referred to the U.S. economy.

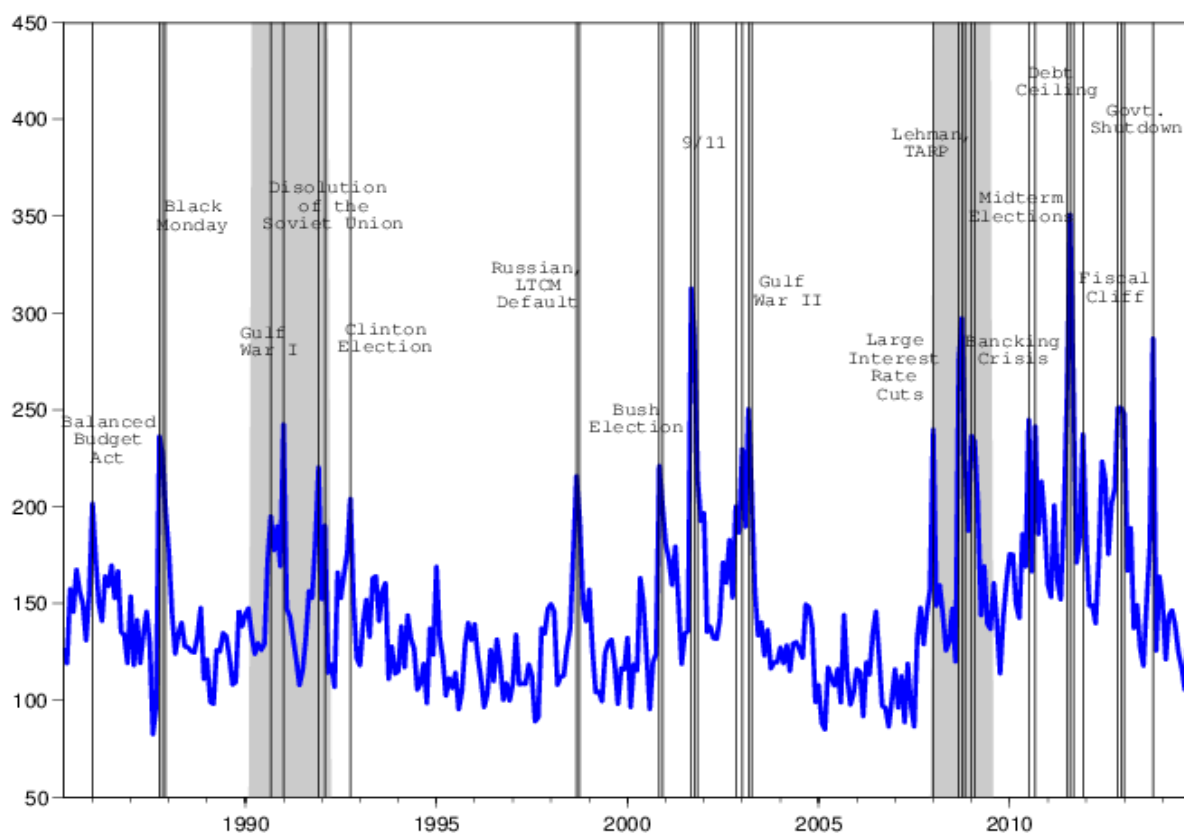


<b>Busts</b>							
Shock/Variable	$EPU_t^{US}$	$EPU_t$	$\overline{\Delta IP}_t$	$u_t$	$\pi_t$	$R_t$	$\Delta \epsilon_t$
$\tilde{\epsilon}^{EPU_t^{US}}$	0.65	<b>0.34</b>	0.08	<b>0.27</b>	0.12	0.15	0.13
$\tilde{\epsilon}^{EPU_t}$	<b>0.07</b>	0.39	0.14	<b>0.24</b>	0.10	0.17	0.20
$\tilde{\epsilon}^{R_t}$	0.04	0.06	0.07	0.05	0.12	0.51	0.04

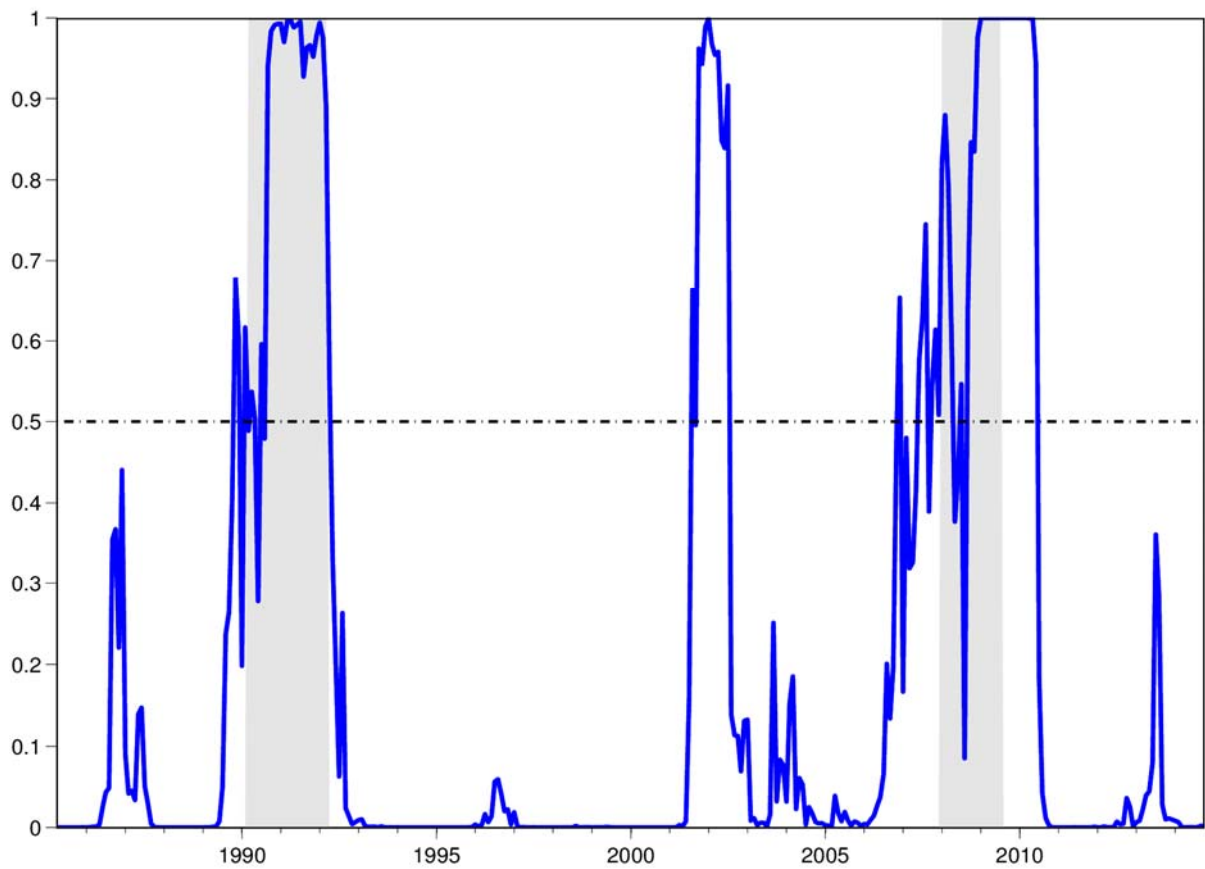
  

<b>Booms</b>							
Shock/Variable	$EPU_t^{US}$	$EPU_t$	$\overline{\Delta IP}_t$	$u_t$	$\pi_t$	$R_t$	$\Delta \epsilon_t$
$\tilde{\epsilon}^{EPU_t^{US}}$	0.74	<b>0.32</b>	0.01	<b>0.08</b>	0.03	0.05	0.03
$\tilde{\epsilon}^{EPU_t}$	<b>0.02</b>	0.45	0.03	<b>0.02</b>	0.04	0.09	0.04
$\tilde{\epsilon}^{R_t}$	0.02	0.05	0.10	0.02	0.12	0.65	0.02

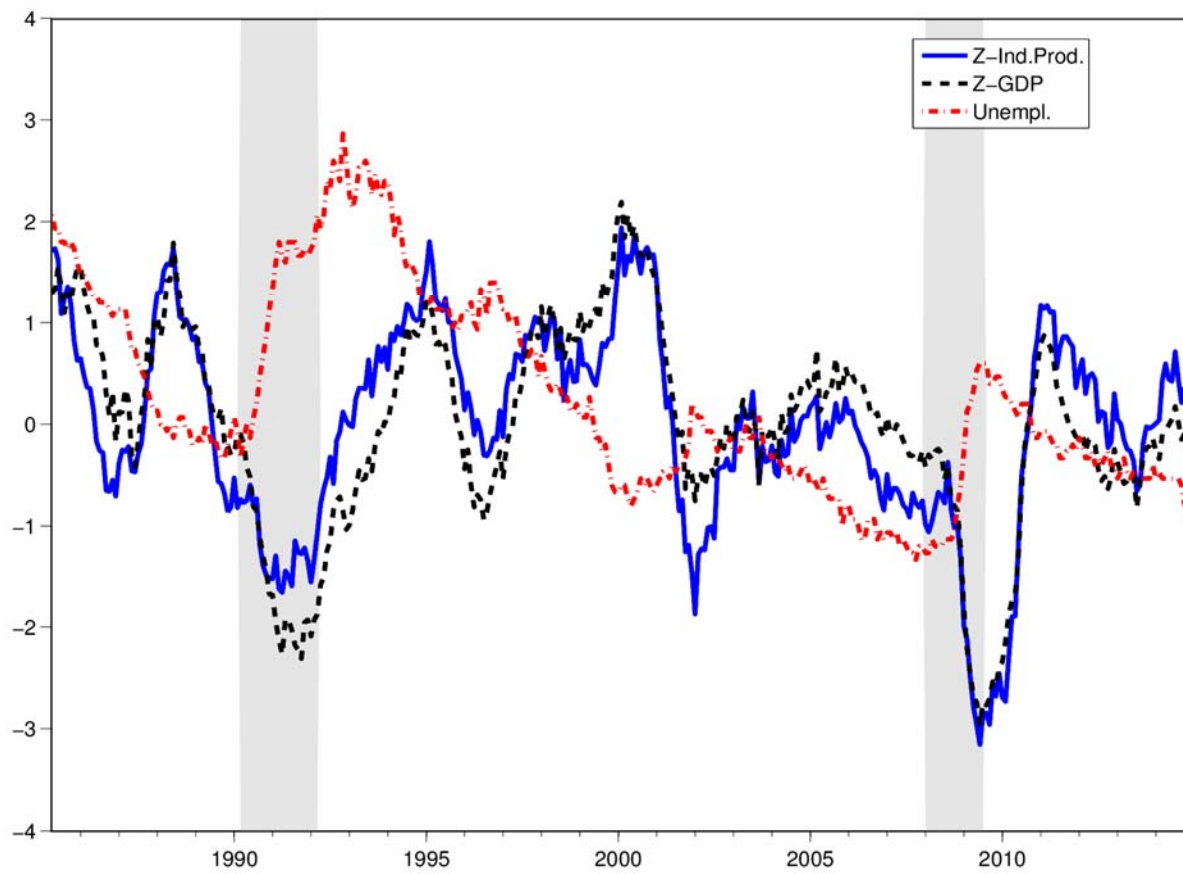
**Table 2. Forecast Error Variance Decomposition: U.S. vs. Canadian EPU Shocks.** 2 year-ahead forecast error variance decomposition. The figures reported in the table refer to the point estimates of the baseline model.



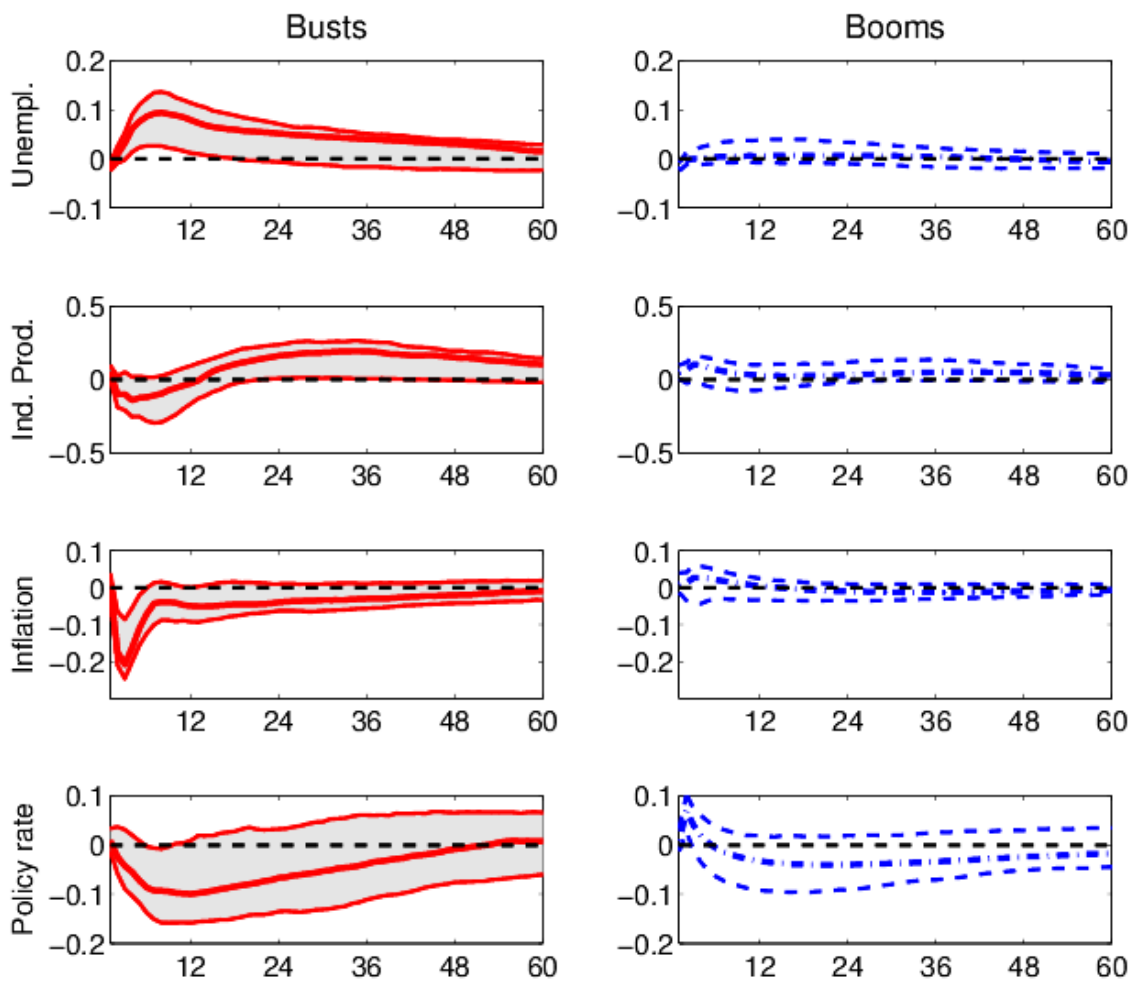
**Figure 1. U.S. EPU Dummy.** Blue line: Historical EPU index for the United States as in Baker, Bloom, and Davis (2016). Black vertical lines: Realizations of the cyclical component of the EPU index (computed via the Hodrick-Prescott filter, smoothing weight: 129,600) whose value is larger than 1.65 times the standard deviation of the EPU index cyclical component. Grey vertical bars: ECRI recessions.



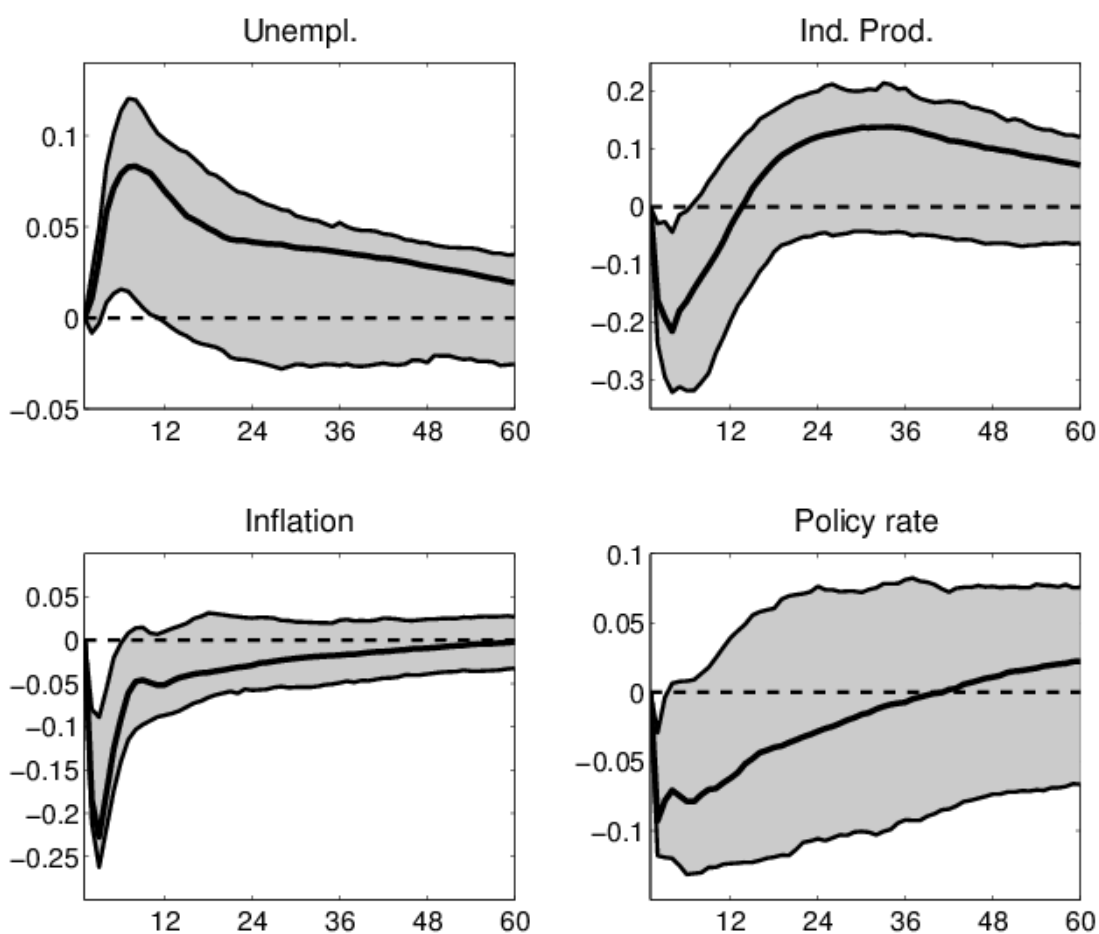
**Figure 2. Probabilities of Economic Booms for Canada as Estimated by the STVAR model.** Function  $[1 - F(z)]$  estimated jointly with the STVAR, baseline version with the U.S. EPU dummy. Transition indicator  $z$ : Moving average of the month-to-month growth rate of the Canadian industrial production comprising eighteen terms. The point estimate for the slope parameter is  $\hat{\gamma}=6.36$  and for the threshold value is  $\hat{c}=-0.72$ .



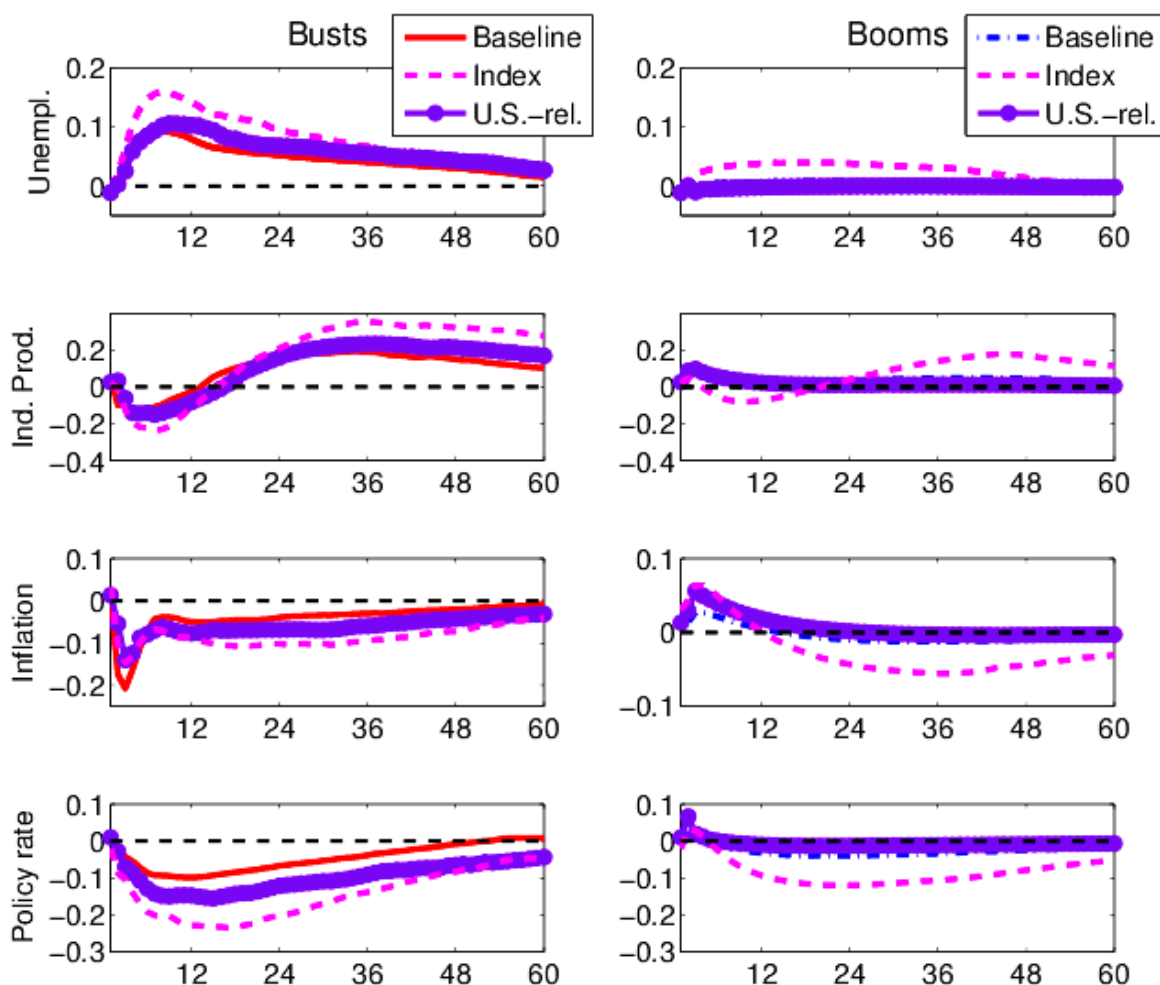
**Figure 3: Canada. Different Real Activity Indicators.** Moving averages of the monthly growth rates of industrial production and real GDP consider eighteen terms. All the activity indicators are normalized to have unit variance and zero mean.



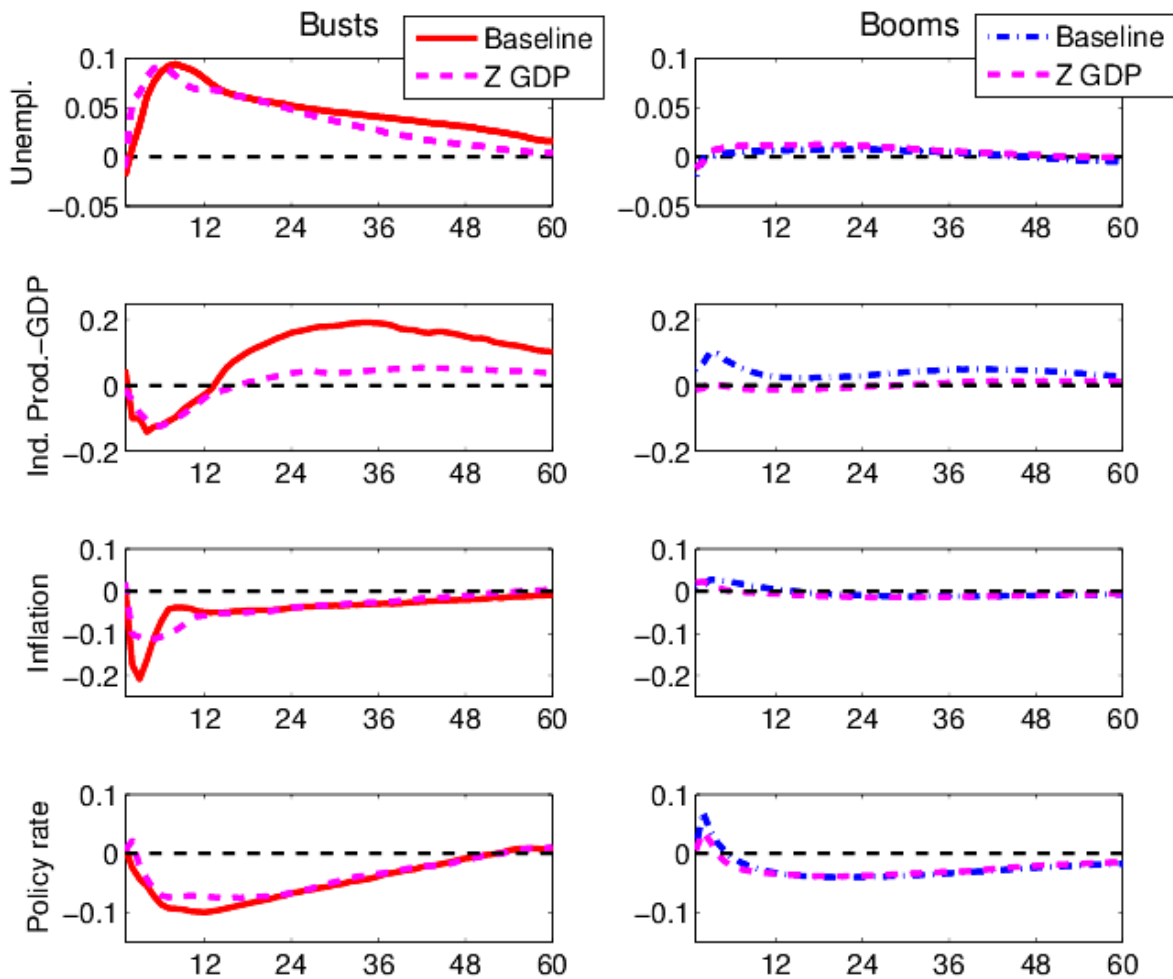
**Figure 4. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy.** Sample: 1985:M1-2014:M10. Generalized median impulse responses to a one-standard deviation shock to the U.S. EPU dummy hitting the Canadian economy in busts (red solid line) and booms (blue dash-dot line). 68% confidence intervals identified via shaded areas (busts) and dashed blue lines (booms). Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.



**Figure 5. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Difference between states.** Sample: 1985:M1-2014:M10. Differences between generalized median impulse responses in busts and booms to a one-standard deviation shock to the U.S. EPU dummy. Median realizations identified via black lines, 68% confidence intervals identified via shaded areas. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.

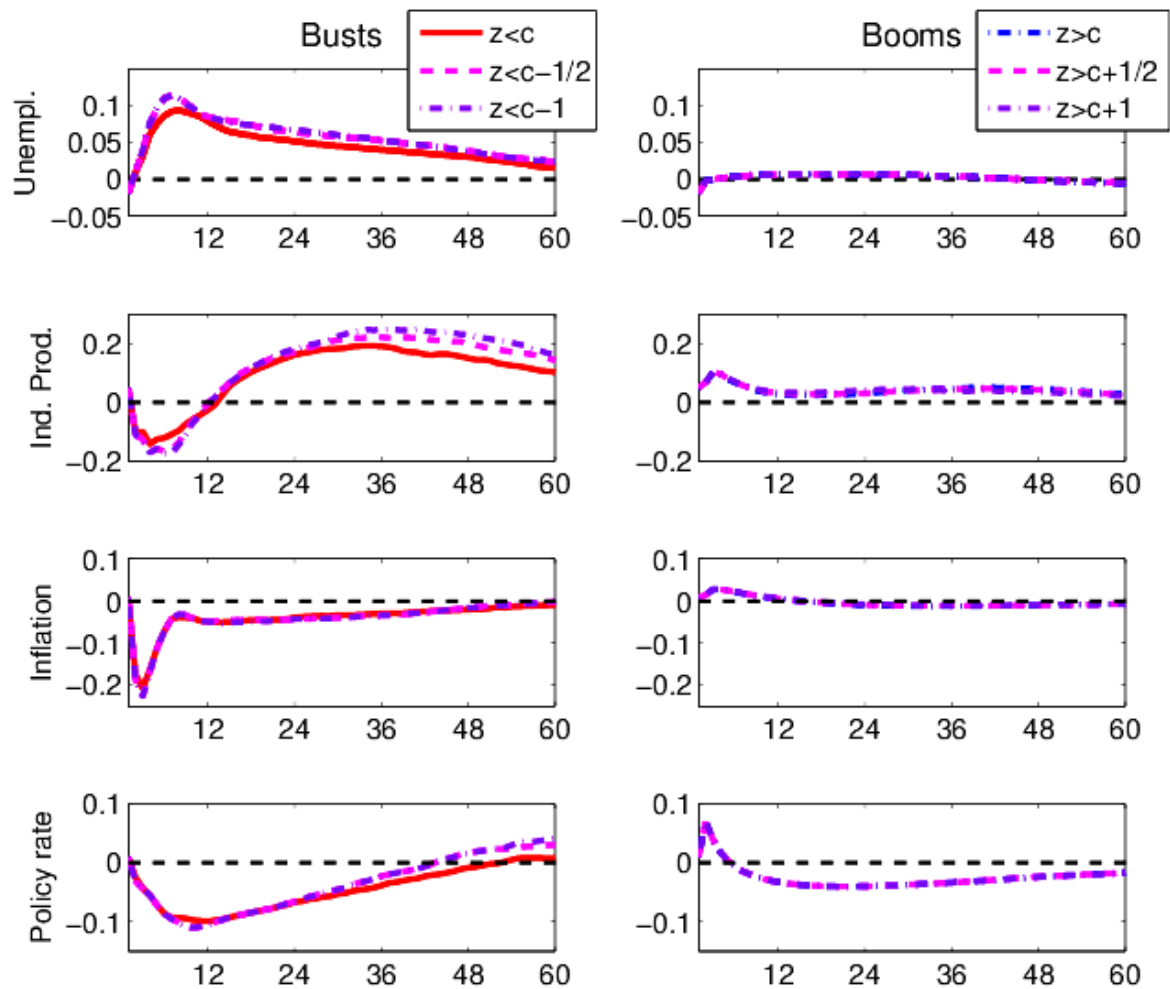


**Figure 6. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Robustness to different proxies for uncertainty in the United States.** Generalized impulse responses to a one-standard deviation shock to a proxy for the U.S. EPU. Baseline/Index/U.S.-rel. indicates exercises conducted with the U.S. EPU dummy as in the baseline case, the U.S. EPU index, and the dummy constructed by selecting only U.S.-related episodes, respectively. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.

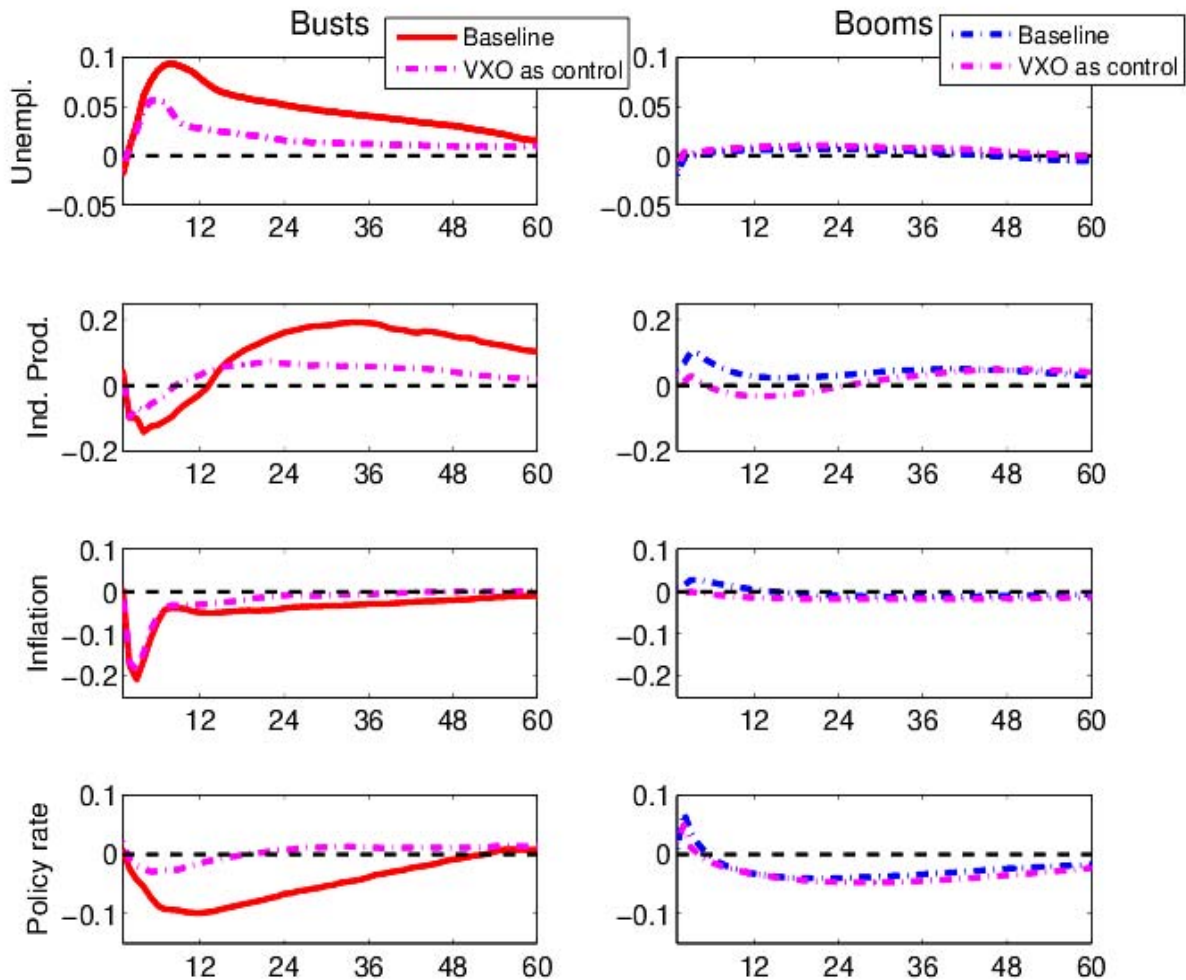


**Figure 7. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Robustness to different transition indicators.** Generalized median impulse responses to a one-standard deviation shock to the U.S. EPU dummy hitting the Canadian economy in busts and booms. Transition indicators for Canada: Baseline/Z GDP, which refer to a 18-term moving average of the monthly growth rate of the Canadian industrial production and a 18-term moving average of the monthly growth rate or real GDP, respectively.

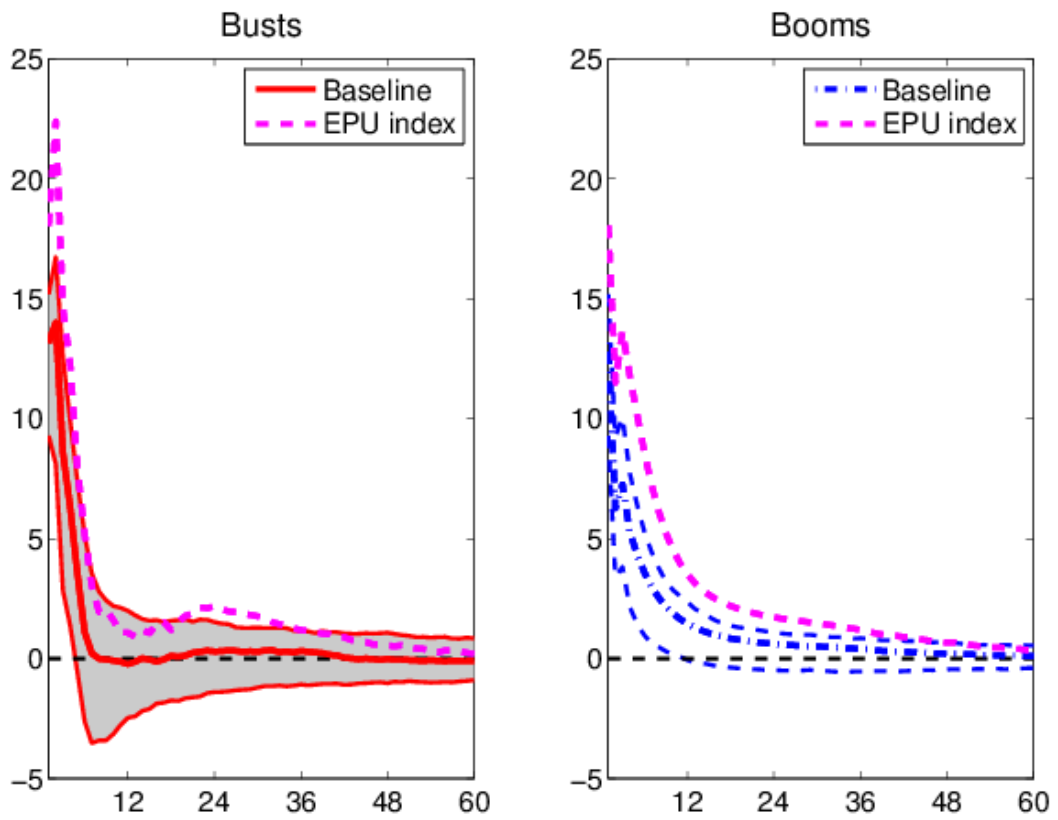




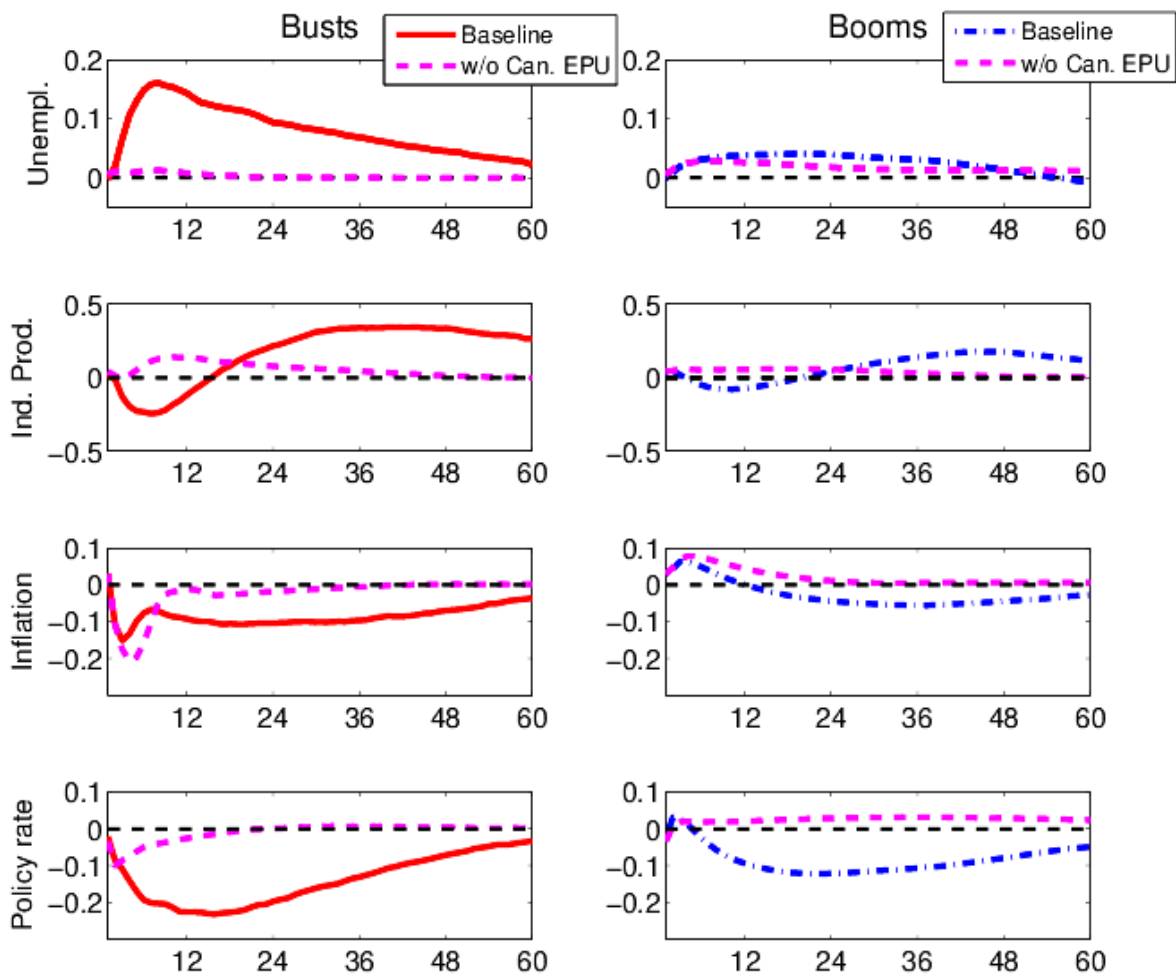
**Figure 8. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Robustness to different sets of initial conditions.** Generalized median impulse responses to a one-standard deviation shock to the U.S. EPU dummy hitting the Canadian economy in busts and booms. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.



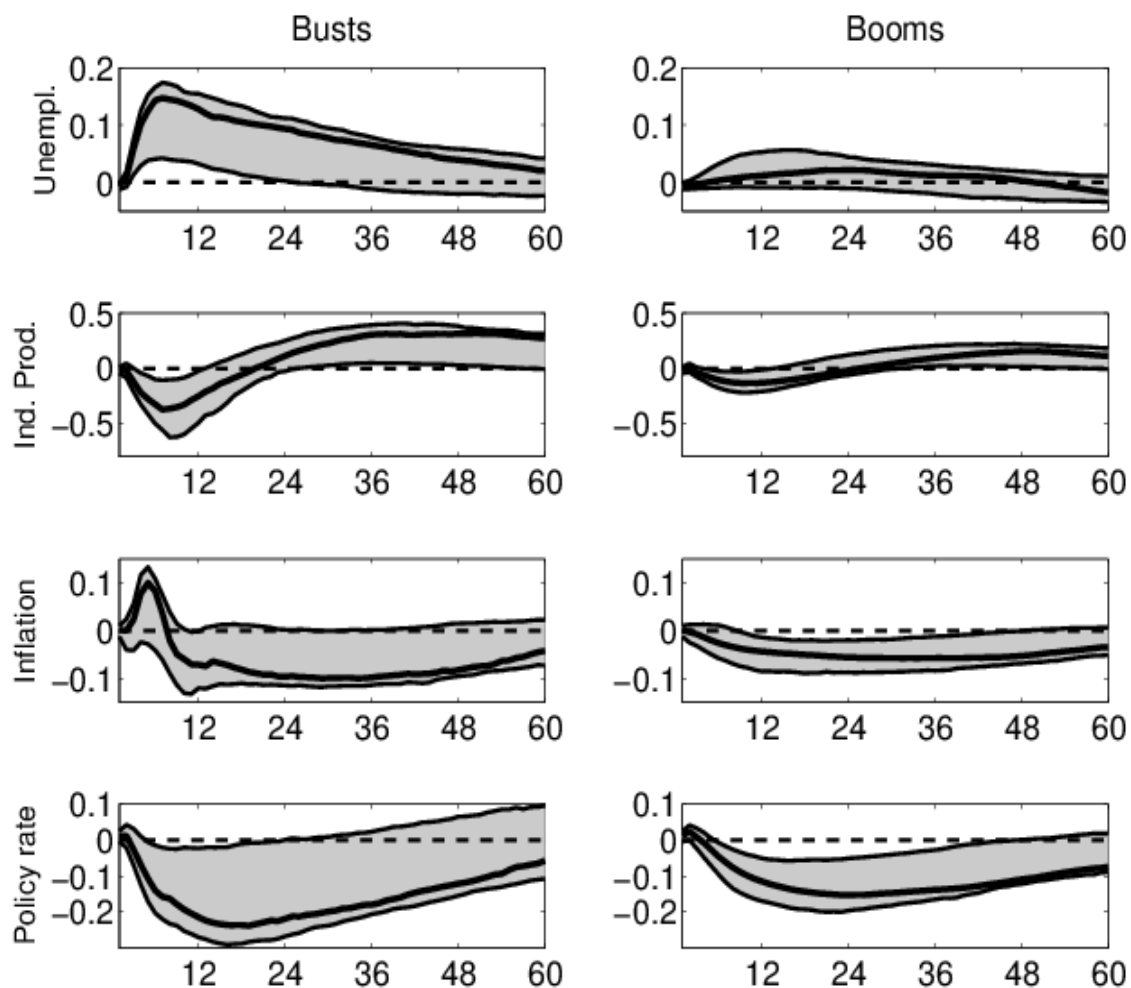
**Figure 9. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Robustness to the inclusion of broader measures of U.S. uncertainty.** Generalized median impulse responses to a one-standard deviation shock to the U.S. EPU index hitting the Canadian economy in busts and booms. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.



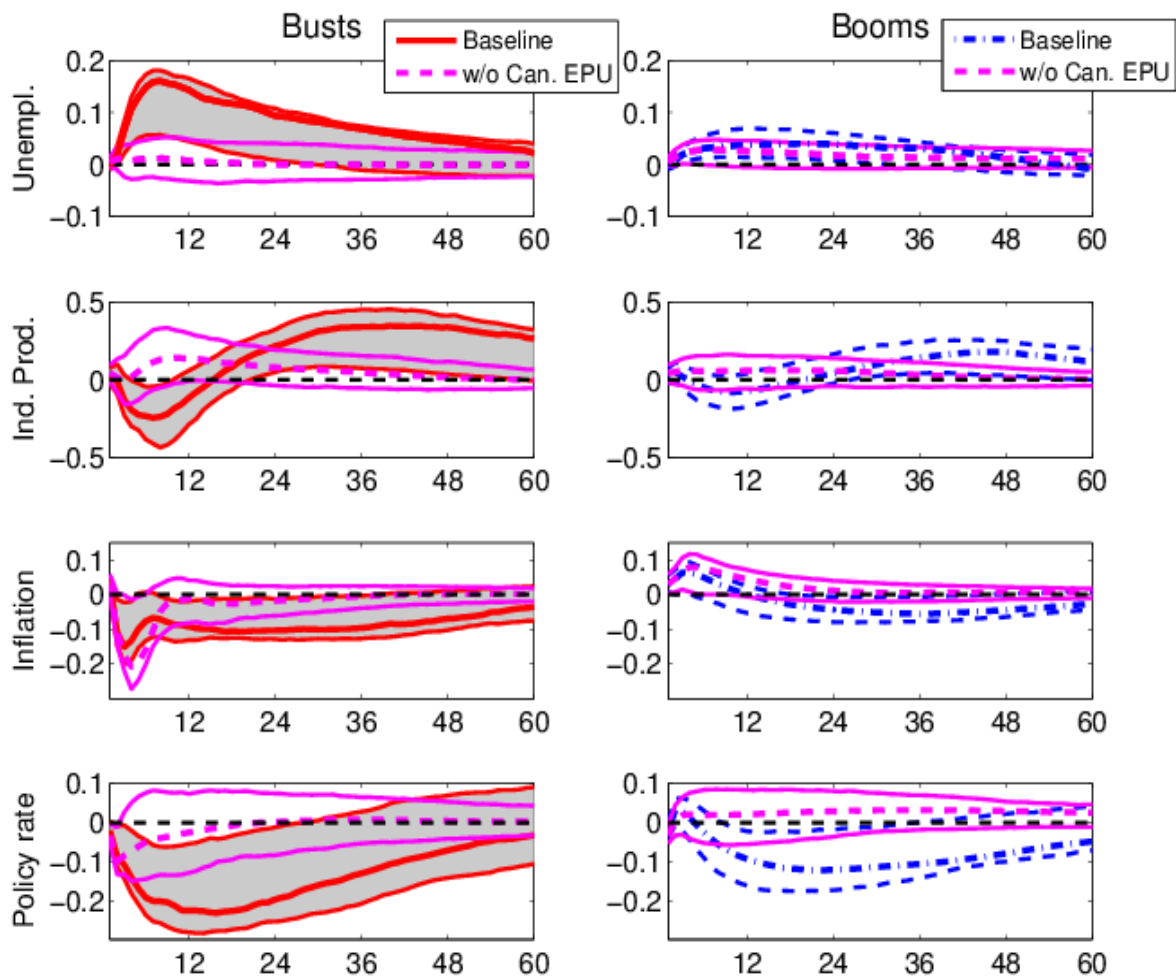
**Figure 10. Response of Canadian EPU to U.S. EPU shocks.** Baseline case: Model with our U.S. EPU dummy. EPU index: Model with U.S. EPU index a la Baker, Bloom, and Davis (2016).



**Figure 11. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Role of Domestic Uncertainty.** Generalized median impulse responses to a one-standard deviation shock to the U.S. EPU index hitting the Canadian economy in busts and booms. Counterfactual simulations conducted by working with fictitious shocks to the Canadian EPU index to keep it fixed. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.



**Figure 12. Role of Domestic Uncertainty: Statistical Difference.** Differences between “baseline” minus “muted Canadian EPU” impulse responses to a one-standard deviation shock to the U.S. EPU index hitting the Canadian economy in busts and booms. Median realizations identified via black lines, 68% confidence intervals identified via shaded areas. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.



**Figure 13. Effects of a Shock to the U.S. EPU Dummy on the Canadian economy: Role of Domestic Uncertainty. Girls with confidence intervals.** Generalized median impulse responses to a one-standard deviation shock to the U.S. EPU index hitting the Canadian economy in busts and booms. Counterfactual simulations conducted by working with fictitious shocks to the Canadian EPU index to keep it fixed. 68% confidence intervals identified via shaded areas (busts), dashed blue lines (booms) and solid lines in magenta (counterfactual). Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.