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Abstract

We estimate a nonlinear VAR to quantify the impact of economic policy uncertainty shocks originating in the US on the Canadian unemployment rate in booms and busts. We find strong evidence in favor of asymmetric spillover effects. Unemployment in Canada is shown to react to uncertainty shocks in economic busts only. Such shocks explain about 13% of the variance of the 2-year ahead forecast error of the Canadian unemployment rate in periods of slack vs. just 2% during economic booms. Counterfactual simulations lead to the identification of a novel "economic policy uncertainty spillovers channel". According to this channel, jumps in US uncertainty foster economic policy uncertainty in Canada in the first place and, because of the latter, lead to a temporary increase in the Canadian unemployment rate. Evidence of asymmetric spillover effects due to US EPU shocks are also found for the UK economy. This evidence, which refers to a large economy having a low trade intensity with the US, supports our view that a channel other than trade could be behind our empirical results.

Keywords

Economic Policy Uncertainty Shocks, Spillover Effects, Unemployment Dynamics, Smooth Transition Vector Autoregressions, Recessions.

JEL Classification

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 US and a number of other countries. When employing such index in VAR investigations, 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 US and other industrialized countries. This result is important for two reasons. First, because it reaffirms that uncertainty can 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 (2014, 2017)). 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 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 largely affected by shocks coming from neighboring countries and the rest of the world in general. As a matter of fact, however, little is known on the spillover effects related to second moment shocks, and - in particular - economic policy uncertainty shocks.

This paper contributes to the analysis of the transmission of second moment shocks in open economies by focusing on economic policy uncertainty spillovers from the United States to Canada. To this end, we estimate a monthly nonlinear Smooth Transition VAR (STVAR) model for the period 1985-2014.[1] In this framework, economic policy

[1] The measures of EPU we use for Canada and, in a related exercise reported in Section 7, for the UK are based on information contained in newspapers only. An equivalent measure - the historical EPU series - is available for the US up to 2014, which justifies the end date of our sample. A highly correlated, updated series for the US is available starting from 1985. As documented in our Appendix, a robustness check conducted with this updated series, available until 2017, returned results virtually unchanged with respect to those documented in this paper.

uncertainty shocks originating in the US are allowed (but not necessarily required) to act as drivers of real activity in Canada, with possibly asymmetric effects depending on the stance of the Canadian business cycle. We model a number of Canadian macroeconomic variables, including real activity indicators (industrial production, unemployment), inflation, the policy rate, and the US-Canada bilateral real exchange rate. In computing the effects of US EPU shocks on the Canadian economy, we control for the Canadian EPU index, so that uncertainty shocks originating in the US can affect the Canadian economy via uncertainty spillovers. Notice that a jump in policy-related US uncertainty is potentially recessionary, and can in principle lead the Canadian economy to switch from - say - a boom to a bust. As a consequence, modeling absorbing states and estimating conditionally linear impulse responses would likely lead to biased results. Our analysis accounts 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.

We find statistically and economically relevant nonlinear spillover effects. An economic policy uncertainty hike originating in the US is estimated to trigger a strong and persistent downturn in Canada in the 1985-2014 period during bad economic times. An equally-sized shock, when occurring in booms, leads to quantitatively milder and mostly insignificant responses of real activity indicators. Monetary policy reacts by lowering the interest rate much more in bad than in good times. 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 US explain up to 13% of the variance of the 2-year ahead forecast error of the Canadian unemployment rate during slow-growth phases vs. about 2% during economic booms.

One of the variables reacting in a significant and persistent fashion to US EPU shocks is the Canadian EPU index. We then analyze the role played by the evolution of the latter in the transmission of US EPU shocks to the Canadian economy *vis-à-vis* the role played by bilateral trade. We do so by estimating a version of the STVAR model that features Canadian net exports to the United States. Focusing on economic busts, we then conduct two counterfactual exercises to simulate the response of unemployment in Canada to a US EPU shock when i) the Canadian EPU index is not allowed to react to systematic movements in US economic policy uncertainty, and ii) net exports do

not react to US EPU. The response of the Canadian unemployment rate turns out to be dramatically dampened only in the counterfactual scenario where the response of Canada EPU is muted, while it is virtually unchanged relative to the baseline case in the scenario of a muted response of net exports. Such results point to the existence of a novel "economic policy uncertainty spillover channel": hikes in the level of the US economic policy uncertainty foster the build up of EPU in Canada and, consequently, exert a negative effect on the Canadian business cycle. To reinforce the evidence that the main transmission channel of US EPU shocks to other economies is not related to trade, we conduct a similar exercise for the UK, a large open economy with a relatively low degree of trade intensity with the United States.[2] Again, we find that US EPU shocks generate significant real effects that are asymmetric over the business cycle. We interpret this result as pointing to the relevance of channels alternative to the standard trade-related one for the international transmission of US EPU shocks.

We focus on US and Canada to investigate whether economic policy uncertainty originating in a large country can affect, and via which mechanism, business cycle fluctuations in a smaller open economy for three main reasons. First, the degree of interconnection between the US and Canada is high. According to the Observatory of Economic Complexity (OEC), 74% of Canadian total exports were imported by the United States, and 55% of Canadian imports came from the United States in 2014, the end year of our sample.[3] Second, first-moment shocks like technology, monetary policy, and fiscal shocks originating in the United States are typically found to explain a sizeable fraction of the volatility of real activity in Canada (see, among others, Schmitt-Grohe (1998), Justiniano and Preston (2010), Kulish and Rees (2011), Faccini, Mumtaz, and Surico (2016), Ong (2018)). For instance, Justiniano and Preston (2010) document that 52% of 2 year-ahead Canadian output growth volatility is explained by first-moment US shocks. Our paper complements these contributions by focusing on US economic policy uncertainty shocks, which are second-moment shocks. Third, as pointed out by Bloom (2017), small open economies like Canada are likely to be hit by large uncertainty shocks that have a foreign origin, and they can be claimed to be unrelated to the domestic business cycle and, therefore, exogenous. Hence, small open economies like Canada are the ideal laboratory to identify the causal link going from uncertainty to real activity.

Our analysis focuses mainly on the possibly asymmetric response of unemployment.

[2] In 2014, the share of total UK exports imported by the US was 11%., while a share equal to 6.7% of total UK imports was coming from the US. In 1985, the shares were 16% and 11%, respectively.

[3] In 1985, the initial year of our analysis, these ratios were 77% and 72%, respectively. Data available at http://atlas.media.mit.edu/en/ .

Our search for asymmetric responses of unemployment 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 movement 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 the relationship between economic slack and low inflation, movements in aggregate demand caused by spikes in uncertainty may have larger real effects in periods of low growth. Cacciatore and Ravenna (2015) show that deviations from the efficient wage-setting due to matching frictions in the labor market combined with downward wage rigidities generate a strong and state-dependent amplification of uncertainty shocks and contribute to generate a countercyclical aggregate uncertainty. Sichel (1993) proposes a test for deepness and steepness and find empirical support for both when working with the US unemployment rate. Evidence pointing to an asymmetric behavior of the US 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, uncertainty dramatically increases during economic downturns (Jurado, Ludvigson, and Ng (2015), Bloom, Floetotto, Jaimovich, Saporta-Eksten, and Terry (2018)), a countercyclical behavior which is featured also by unemployment. Hence, the effects triggered by uncertainty shocks in recessions are likely to be different than those occurring in expansions. Recent evidence on the US economy along this line is provided by, among others, Nodari (2014), Caggiano, Castelnuovo, and Groshenny (2014), Ferrara and Guérin (2015), Caggiano, Castelnuovo, and Nodari (2017), and Caggiano, Castelnuovo, and Figueres (2017), while Casarin, Foroni, Marcellino, and Ravazzolo (2018) find evidence in favor of state-dependent uncertainty-related coefficients in a panel approach modeling the US, a number of European countries, and Japan.

The structure of the paper is the following. Section 2 makes contacts with the extant literature. Section 3 details our empirical set up. In particular, it discusses the identification of an US EPU-related uncertainty shock and 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. A list of robustness checks, which confirm the baseline results, are documented in Section 5. Section 6 looks at the importance of EPU shocks for the Canadian business cycle by reporting FEVD, and documents the existence of an "economic policy uncertainty spillover" channel. Section 7 extends our analysis to the United Kingdom. Section 8 concludes.

2 Related literature

Our paper joins three different but related strands of the literature on the role of uncertainty shocks. First, a growing strand of the literature has studied the measurement and the macroeconomic effects of economic policy uncertainty. Proxies for uncertainty have been constructed via measures of forecast disagreement (Bachmann, Elstner, and Sims (2013)), by relating the location of the real GDP forecast errors to the sample distribution of the forecast errors of the same variable (Rossi and Sekhposyan (2015,2016)), by modeling the common component of the volatility of the forecast errors of several macroeconomic and financial indicators (Jurado, Ludvigson, and Ng (2015), Ludvigson, Ma, and Ng (2018), and Carriero, Clark, and Marcellino (2018)), by exploiting Bloomberg forecasts to capture agents' uncertainty surrounding current realizations of real economic activity (Scotti (2016)), focusing on interest rate uncertainty as done by Creal and Wu (2017) and Istrefi and Mouabbi (2017), or working with Google Trends data as Castelnuovo and Tran (2017). The focus of this paper is on economic policy uncertainty. Baker, Bloom, and Davis (2016) have developed 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 US 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 US and Canada to study the spillover effects of hikes in EPU uncertainty from the former country to the latter. Other contributions using these EPU indices to study the macroeconomic impact of policy-related uncertainty shocks have mainly focused on the US taken in isolation. Working with a VAR model, Benati (2013) shows that economic policy uncertainty can

explain about 20-30% of the 1-year ahead forecast error variance of the US industrial production growth rate, and it is 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 American business cycle. Istrefi and Piloiu (2015) document a link between economic policy uncertainty and short- and long-run inflation expectations. Our contributions complement this literature by highlighting an international economic policy uncertainty transmission channel which works asymmetrically along the business cycle in a small-open economy like Canada, as well as in a larger economy like the United Kingdom.

The second strand of the literature focuses on the role of uncertainty in an open economy context. Fernández-Villaverde, Guerrón-Quintana, Rubio-Ramírez, and Uribe (2011) and Born and Pfeifer (2014b) 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 Nisticò (2012) find shocks to the volatility of monetary policy shocks, inflation target shocks, and productivity shocks realizing in the US 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 US real GDP leads to a decline in UK GDP of 1% relative to trend and a 0.7% increase in UK CPI at a 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. Carrière-Swallow and Céspedes (2013) study the impact of uncertainty shocks originating in the US 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. Klößner and Sekkel (2014) study international spillovers of policy uncertainty and find evidence in favor of economic policy uncertainty connectedness for a number of countries, with the US being the main exporter of policy uncertainty. Handley (2014) and Handley and Limão (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, Müller, 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 US 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 US to Canada due to first-moment shocks (Schmitt-Grohe (1998), Justiniano and Preston (2010), Faccini, Mumtaz, and Surico (2016), and Ong (2018)).

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 (2016). They show that a labor market model featuring matching frictions predict 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 imply an even lower number of vacancies posted in equilibrium. Basu and Bundick (2017) 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.[4] Born and Pfeifer (2014a) estimate a new-Keynesian framework featuring policy risk. They find moderately negative output effects after a jump in such risk. Fernández-Villaverde, Guerrón-Quintana, Kuester, and Rubio-Ramírez (2015) also work with a new-Keynesian model and focus on fiscal policy-related uncertainty. They find that unexpected changes in fiscal volatility shocks can have a sizable adverse effect on economic activity. Our results support models predicting a drop in real activity after an uncertainty shock, and stress that this is particularly true when the economy features unused capacity.

3 Modeling asymmetric spillover effects: Shocks and dynamics

The Economic Policy Uncertainty index. As anticipated in the Introduction, we use the index developed by Baker, Bloom, and Davis (2016) for the US and Canada as proxies of economic policy-related uncertainty. This index is based on newspaper coverage frequency. As regards the United States, they 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, The Los Angeles Times, and The Boston Globe. Since 1985, USA Today, The Miami Herald, The Dallas Morning Tribune, and The San Francisco Chronicle have been added to the set. The authors perform within-month searches of all articles, starting in January 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 features terms in all three categories pertaining to uncertainty, the economy and policy. To deal with changing volumes of news articles for a given newspaper over time, Baker, Bloom, and Davis (2016) divide the raw counts of policy uncertainty articles by the total number of news articles containing

[4] de Groot, Richter, and Throckmorton (2017) clarify that Basu and Bundick's (2017) model require the intertemporal elasticity of substitution to be lower than one to generate recessionary effects out of hikes in uncertainty.

terms regarding the economy or business. They then normalize each newspaper's series to unit standard deviation prior to December 2009 and then sum up all series. Details are reported in Baker, Bloom, and Davis (2016).

We now turn to the description of the nonlinear model we employ in our empirical analysis.

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:

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

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

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

where $\boldsymbol{X}_t$ is a set of endogenous variables we aim to model, $\boldsymbol{\Pi}_R$ and $\boldsymbol{\Pi}_E$ 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 variance-covariance matrix $\boldsymbol{\Omega}$, $F(z_{t-1})$ is a logistic transition function which captures the probability of being in a boom and whose smoothness parameter is γ, z_t is a standardized transition indicator, and c is the threshold parameter identifying the two regimes.[5] In brief, this model combines two linear VARs, one capturing the dynamics of the economy during busts and the other one during booms. The transition from a regime to another is regulated by the smoothness parameter γ. Large values of γ imply abrupt switches from a regime to another, while moderate ones point to regimes of longer duration.[6]

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 indexes developed by Baker, Bloom, and Davis (2016) for the US and Canada. We use two lags, as indicated by the AIC. As transition indicator, we employ a standardized moving average of the growth rate of a real activity

[5] Teräsvirta, Tjøstheim, and Granger (2010) point out that γ 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 is a zero-mean, unitary standard deviation variable. This choice makes it easier to guess a good initial condition for the maximization of the likelihood.

[6] Mumtaz and Theodoridis (2016) point to a different way of modelling the possibly evolving role played by uncertainty shocks with an application for the US in which impulse responses are allowed to be time-dependent. A comparison between state- and time-dependent effects of economic policy uncertainty spillovers is material for future research.

indicator, industrial production in our case, in line with other STVAR-based empirical analysis of the US 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)).[7] Conditional on our choice for z_t, we jointly estimate the parameters $\{\mathbf{\Pi}_R, \mathbf{\Pi}_E, \mathbf{\Omega}, \gamma, c\}$ of model (1)-(3) via conditional maximum likelihood as suggested by Teräsvirta, Tjøstheim, and Granger (2010).[8]

Modeled vector. We model the Canadian economy with the following vector of US and Canadian observables: $\boldsymbol{X}_t = [EPU_t^{US}, EPU_t, \overline{\Delta IP}_t, u_t, \pi_t, R_t, \Delta\epsilon_t]'$. The variable EPU_t^{US} is the US EPU uncertainty index constructed by Baker, Bloom, and Davis (2016). All the remaining variables in the vector $\boldsymbol{X}_t$ refer to the Canadian economy. In particular, EPU_t stands for the Canadian uncertainty index, $\overline{\Delta IP}_t$ stands for the eighteen-term moving average of the monthly growth rate of industrial production (percentualized and annualized), u_t is the unemployment rate, π_t stands for CPI inflation (y-o-y percentualized growth rate of the monthly index), R_t is the policy rate, while $\Delta\epsilon_t \equiv \pi_t^{US} + \Delta s_t^{CAN,US} - \pi_t^{CAN}$ is the growth rate of the bilateral real exchange rate between Canada and the US constructed by considering the inflation rates in the two countries and combining it with $\Delta s_t^{CAN,US}$, which is the y-o-y growth rate of the Canada/US nominal exchange rate.[9]

We consider the sample January 1985 - October 2014. The start date 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 US EPU index deliver infor-

[7] We employ a backward-looking moving average of the month-by-month growth rate of industrial production featuring eighteen terms. The number of moving average terms is determined by maximizing the correlation between our transition variable and the dating of Canadian recessions as defined by the Economic Cycle Research Institute (https://www.ecri.org). Our Appendix shows that our results are robust to two alternative transition indicators, a moving average of the real GDP growth rate and the common factor computed via a principal component analysis which considers the growth rates of industrial production and real GDP and the rates of unemployment and employment.

[8] Notice that an alternative to modeling a smooth transition from a regime to another with an observable indicator would be to model the switch by estimating a latent factor process as in the context of a different - although related - regime switching model (for an extensive presentation, see Hamilton (2016)). We prefer using an observable to determine the regimes in our model (conditional on the estimated values of the logistic function $F(z_t)$) to ease the comparison of our estimated impulse responses to the theoretical and empirical literature cited in the Introduction and in Section 2, which motivates the asymmetric dynamics of the unemployment rate over the business cycle.

[9] 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 different newspapers: The Gazette, The Globe and Mail, Canadian Newswire, The Ottawa Citizen, The Toronto Star, and The Vancouver Sun. The Canadian policy rate is the immediate rates: Less than 24 hours: Central Bank rates for Canada.

mation over and above the one delivered by abrupt changes in the Canadian one.[10] The end date is justified by the availability of the EPU historical index for the United States. The EPU indices were downloaded from the http://www.policyuncertainty.com/ website. All remaining data were downloaded from the Federal Reserve Bank of St. Louis' website.

Linearity test. We test if a nonlinear framework provides us with a statistically better representation of the covariance structure of the data $\boldsymbol{X}_t$ relative to a standard linear multivariate framework. Teräsvirta and Yang (2014) propose a Lagrange Multiplier test of the null hypothesis of linearity vs. a specified non-linear alternative that is exactly the logistic STVAR framework with a single transition variable. The Lagrange Multiplier statistic is 241.12, with a p-value equal to 0.00, which clearly points to the rejection of the null hypothesis of linearity of the model. Details on this test are reported in our Appendix.

4 EPU spillovers: Empirical evidence

This Section reports our main empirical findings. We begin by describing the identified US EPU shocks. We then show the estimated probability of slack for Canada according to our model. Finally, we report the GIRFs of the Canadian macroeconomic indicators to an uncertainty shock coming from the United States.

EPU shocks. We identify US EPU shocks by orthogonalizing the residuals $\boldsymbol{\varepsilon}_t$ in (1) via a Cholesky-decomposition of the variance-covariance matrix $\boldsymbol{\Omega}$. Since we place the US EPU index first in the vector $\boldsymbol{X}_t$, this identification scheme implies that Canadian variables cannot exert a contemporaneous impact on the US uncertainty index. This assumption is weaker than the block-exogeneity assumption usually entertained when working with a small-open economy model for Canada, and let the data free to speak as regards possible feedbacks from Canada to the United States.

Figure 1 plots the estimated series of the US EPU shocks. Vertical lines identify upward spikes in this series that can be interpreted as "large" uncertainty shocks.[11] We give all these spikes an interpretation based on historical facts, which we report in Table 1. Some spikes regard monetary or fiscal policy related events, like the large

[10] In February 1991, the Bank of Canada officially adopted an inflation target. Our results are robust to the employment of the sample 1991:M2-2014:M10 (evidence available upon request).

[11] These large shocks are identified as positive realizations of the US EPU shocks estimated with our baseline model (and displayed in Figure 1) which exceed the value of two standard deviations of the shock.

interest rates cuts in early 2008 and the discussions on the budget and the fiscal cliff in 2011 and 2012. These are shocks which we associate to domestic (US) economic conditions, and are likely to be exogenous to the Canadian business cycle. All these events can potentially increase the uncertainty on how economic policy will operate in the future in the US and, as such, represent important drivers behind firms' and households' economic decisions, that eventually affect real activity, both domestically and in countries which are strictly interconnected to the United States, Canada in first place. A few spikes relate to events like the Gulf War I in 1991, the invasion of Iraq in 2003, and the acceleration of the Global Financial Crises in 2008, which can be classified as "global" (i.e., non-US only) events. Our Section on robustness checks document that our results are robust to shocks identified with a US EPU dummy which considers only US related events.

Probability of being in a slack period. Figure 2 plots the estimated probability $[1 - F(z)]$ of being in a negative phase of the business cycle for Canada and contrasts it with the 1990-92 and 2008-09 recessions as dated by the Economic Cycle Research Institute (ECRI).[12] Our estimated logistic function for Canada detects both recessions. The delay with which these two deep downturns are tracked is due to the backward-looking nature of the transition indicator we use. Conditional on our estimated threshold $\widehat{c}$, our model classifies as recession any date t for which $z_t < \widehat{c}$, which in turn implies $[1 - F(z_t)] > 0.5$.[13] This leads us to classify about 18% 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, which is not classified as a recession by the ECRI. As explained in detail in our Appendix, the reason for this discrepancy is the following. The early 2000s saw Canada experience a drop in industrial production as large as the one experienced during the two ECRI recessions in our sample. However, labor market

[12] 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 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, which regards the fact that we will later use the same source for dating the UK business cycle.

[13] The estimated value of $\widehat{c}$ is consistent with a threshold value for the non-standardized transition variable of -1.34. This implies that our model defines as recession any period t in which the eighteen-term moving average of the monthly growth rate of industrial production has been less than -1.34%.

indicators pointed to a strong downturn, but not to a clear recession.[14] 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 estimated threshold, more than on official "expansions" and "recessions".

GIRFs. Figure 3 plots the nonlinear impulse responses of a selected subset of Canadian macroeconomic variables to a one-standard deviation shock to the US EPU shock, along with 68% confidence bands computed with the bootstrap-after-bootstrap methodology proposed by Kilian (1998).[15] We focus on unemployment and 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 US to Canada during phases of slack. An unexpected hike in the US 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 response of unemployment and industrial production is clearly asymmetric. In particular, the response of real activity is strong and statistically relevant during busts, while it is economically modest and statistically insignificant in booms. Third, differently from unemployment, industrial production displays an abrupt drop, a quick rebound, and a prolonged (but temporary) overshoot when the shock hits in a phase of slack. This pattern is in line with the one predicted, for real activity indicators, by Bloom's (2009) partial equilibrium model featuring labor and investment non-convex 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. Vice-versa, a positive short run reaction is detected

[14] The Canadian 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%).

[15] Our GIRFs are computed by considering all realizations (and the corresponding initial conditions) of our transition indicator below/above the estimated threshold as busts/booms. Our Appendix shows that our results are robust to selecting initial conditions corresponding to more "extreme" realizations of the business cycle which, without doubt, can be classified as belonging to the "busts"/"booms" regimes. For an example in the literature of this "extreme events" analysis, see Caggiano, Castelnuovo, Colombo, and Nodari (2015).

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. As documented by Figure 4, these responses are statistically different between states.[16]

5 Robustness checks

We check the robustness of our baseline results along four different dimensions: i) the identification of US-related EPU shocks; ii) the inclusion of proxies of US financial and economic uncertainty; iii) the control for US first moment shocks; iv) the control for commodity and oil price fluctuations.

US EPU dummy. The results shown in Section 4 rely on the use of the EPU index for the US as an observable in the VAR, whose orthogonalized residuals are interpreted as US shocks external to Canada. However, some of the spikes of the EPU index can be attributed to events connected to global pressure. Davis (2016) proposes a Global Economic Policy Uncertainty (GEPU) index constructed by considering a GDP-weighted average of national EPU indices for 16 countries that account for two-thirds of global output. The national EPU indices are constructed following Baker et al.'s (2016) newspaper-based approach. In our sample, the GEPU index rises sharply in correspondence of the Asian Financial Crisis, the 9/11 terrorist attacks, the US-led invasion of Iraq in 2003, the Global Financial Crisis in 2008-09, and the European immigration crisis.[17] Some global events are indeed picked up by the estimated US EPU shock plotted in Figure 1. In particular, the Gulf War I spike, the one corresponding to the Iraq invasion, and the one identifying the peak of uncertainty due to the Global Financial Crisis can be considered as global shocks, more than domestic (US) shocks.

How relevant are these global shocks for our result? We address this question by

[16] To account for correlation between the impulse responses in the two states, the differences plotted in Figure 4 are computed conditional on the same set of bootstrapped residuals. The empirical density of the difference is estimated using 500 realizations per each horizon of interest.

[17] Perhaps not surprisingly, the correlation between the GEPU index and the US EPU one is very high and equal to 0.84. We avoid jointly modeling them in our vector to avoid issues related to multicollinearity.

constructing a US EPU dummy which takes value 1 only for those events that can be safely classified as US-specific. Table 1 lists the US-specific EPU shocks we consider. This dummy is then included in our STVAR *in lieu* of the US EPU index. In this way, we check the solidity of our baseline results to the employment of US-specific shocks only.

Uncertainty ordered last. Our baseline model assumes that both the US and the Canadian EPU indicators are not contemporaneously influenced by any of the shocks originating in Canada. This assumption appears to be plausible and fully consistent with the block-exogeneity approach typically employed when it comes to modeling the interaction between a large economy like the United States and a small-open economy like Canada (see, e.g., Justiniano and Preston (2010)). However, while this assumption is typically entertained for aggregates like inflation and output, less is known on the interconnections between economic policy uncertainty in neighboring countries strictly related by intense trading flows. To understand how relevant this assumption is for our results, we then run a check in which we order US and Canada EPU last, i.e. after the block of Canadian macro variables. This allows both the US and the Canadian uncertainty indices to react on impact to Canadian first moment shocks.

VXO. The EPU index constructed by Baker, Bloom, and Davis (2016) is meant to capture economic policy-related spikes in uncertainty. One concern with our analysis is to what extent we are capturing effects coming from spikes in economic policy uncertainty as opposed to different aspects of economic uncertainty. We tackle this issue by augmenting our VAR with the VXO, which is the S&P 100 implied volatility index computed by the Chicago Board Options Exchange. The VXO index captures the evolution of the volatility of expected stock market returns, and has been used since Bloom (2009) as a proxy of financial uncertainty in applied macroeconomic investigations.[18] Adding the VXO to our VAR allows us to control for movements in a financial measure of uncertainty, which also spikes up in correspondence of events like, e.g., 9/11 which drove the US EPU index up.[19] The idea of our robustness check is then to obtain a

[18] 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.

[19] As stressed by Stock and Watson (2012), uncertainty shocks and liquidity/financial risk shocks are highly correlated, which makes their separate interpretation problematic. For contributions aiming at separating uncertainty and financial shocks, see Christiano, Motto, and Rostagno (2014), Caldara, Fuentes-Albero, Gilchrist, and Zakrajšek (2016), Furlanetto, Ravazzolo, and Sarferaz (2017).

purged measure of the US EPU shocks, which is not driven by financial uncertainty.[20] Following Baker, Bloom, and Davis (2016), we order the VXO after the US EPU index and before the Canadian block of variables in our vector.

Economic Uncertainty. Another possibility is that of confounding economic policy uncertainty with the broader concept of economic uncertainty. Baker, Bloom, and Davis (2016) construct an overall Economic Uncertainty (EU) index by dropping all terms related to policy in the keyword-based search they conduct.[21] We then add the overall EU index to our VAR to isolate the policy component of the US EPU shocks over and above a more general measure of economic uncertainty. As before, we follow Baker, Bloom, and Davis (2016) and order the EU index just after the US EPU index.

Excess bond premium. Recent contributions, e.g. Caldara, Fuentes-Albero, Gilchrist, and Zakrajšek (2016) and Alessandri and Mumtaz (2018), show that the effects of uncertainty shocks are amplified when financial stress is high. Gilchrist and Zakrajšek (2012) propose a micro-founded measure of excess bond premium (EBP). Such measure of credit spread is constructed by controlling for the systematic movements in default risk on individual firms. Consequently, the EBP isolates the cyclical changes in the relationship between measured default risk and credit spreads. Gilchrist and Zakrajšek (2012) show that the EBP has predictive power for a number of US real activity indicators. Moreover, when embedded in a VAR to quantify the effects of a credit shock, unexpected jumps in EBP are found to be associated to temporary but economically significant recessions and deflations. To capture the possibility that movements in EBP caused some of the movements in the US EPU index, we add EBP on top of our baseline vector to control for credit shocks.

Factor-Augmented STVAR. Our baseline model indicates that a substantial chunk of the volatility of the Canadian unemployment rate could be due to a second

[20] Notice that alternative measures of uncertainty are currently available, e.g., the one recently proposed by Ludvigson, Ma, and Ng (2018). Such measure is constructed following the data-rich approach modeling the expected volatility of a large number of financial series proposed by Jurado, Ludvigson, and Ng (2015), who model a broader macroeconomic uncertainty measure combining financial and real economic indicators. Ludvigson, Ma, and Ng (2018) find that financial uncertainty is likely to be a relevant driver of the US business cycle. Notably, 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.

[21] The EU index is available here: http://www.policyuncertainty.com/Replication_Files.zip . A full documentation on the construction of this index is provided in Baker, Bloom, and Davis (2016), who also compare the characteristics of the EPU and EU indices. The EU series is quarterly. We then construct a monthly counterpart via quadratic-match average, which performs a proprietary local quadratic interpolation of the low frequency data to fill in the high observations.

moment shock coming from the United States. However, a number of US first moment shocks - among others, technology, fiscal, and monetary policy shocks - are likely to be at play and influence the Canadian economy. While the separate identification of each of these shocks is left to future research, it is clearly important to control for a composite of these first moment disturbances to minimize the probability of an upward bias as regards the contribution of US EPU shocks on the Canadian unemployment. To tackle this issue, we proceed as follows. First, we use principal component analysis to extract common factors from the 134 monthly US indicators included in the FRED-MD dataset, recently compiled by McCracken and Ng (2016).[22] Second, we consider the first factor in terms of contribution to the variance of the series belonging to the FRED-MD and add it as first variable to our baseline vector. This two-step procedure is meant to emulate the Factor-Augmented VAR (FAVAR) approach proposed by Bernanke, Boivin, and Eliasz (2005) for the identification of a monetary policy shock.

Commodity/oil prices. Canada is a resource-rich country which exports oil and other commodities. As documented by Charnavoki and Dolado (2014), energy products represented 23.5% of total merchandise exports in 2010, while other basic products and materials related to the agriculture sector, forestry and mining accounted for about 40% of those exports. Charnavoki and Dolado (2014) investigate the relevance of shocks to commodity prices for the Canadian economy via a structural dynamic factor model. They measure commodity prices by computing the common factor of a range of indices for energy, food, agricultural raw materials, base metals, and fertilizers, and find commodity price shocks to be an important driver of the Canadian business cycle. Among commodity prices, oil price represents a particularly relevant factor for business cycle fluctuations in a small open economy like Canada both for its direct impact on Canadian exports and for its potential indirect impact via its effects on the US economy. Hamilton (2003) shows that oil price fluctuations have preceded all US recessions included in our sample. Hence, oil prices may very well be an important driver of both uncertainty and real activity in the US and Canada. Our baseline framework does not feature commodity and/or oil prices. We then run three robustness checks in which we enrich our baseline vector with, alternatively, Charnavoki and Dolado's (2014) global commodity price factor, and with two measures of oil prices: the producer price index for crude petroleum and refiner acquisition cost for imported oil.[23]

[22] The database can be downloaded from the website http://research.stlouisfed.org/econ/mccracken/sel/.

[23] As regards Charnavoki and Dolado's measure, we compute the common factor of the commodity price indexes as documented in their paper. The database constructed by Charnavoki and Dolado that contains the series employed to construct such common factor is downloadable from the website

Figure 5 displays the outcome of our robustness checks. A few comments are in order. First, our baseline finding is robust across all robustness checks. In particular, it is clearly robust to the employment of our US EPU dummy, something that points to a spillover effect actually due to US EPU shocks. Second, shocks other than the US EPU one evidently affect the Canadian unemployment rate. The baseline response of unemployment is dampened in most scenarios, with a reduction of the peak response of about 40% in a variety of models. In particular, the models accounting for first moment US shocks, broad economic uncertainty, and fluctuations in oil and commodity prices are those that return the lowest peak reactions of unemployment to an EPU shock. Third, out of the above mentioned models, those incorporating information on EU, oil, and commodity prices predict the peak of unemployment to come after a few months and imply the lowest integral of the response of unemployment. Differently, the model featuring first moment shocks predict unemployment to peak after one year and a somewhat slower speed of convergence towards the steady state. Fourth, the across-model heterogeneity of unemployment responses observed in busts is larger than in booms. This suggests that model misspecification due to the omission of relevant macroeconomic indicators is likely to be more important when studying US EPU spillovers in the context of Canadian busts, compared to booms.

6 EPU shocks: Contribution and transmission mechanism

The results documented so far speak in favor of the fact that variations in the US EPU index can be associated to fluctuations in real activity, inflation, and the short-term interest 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 two counterfactual exercises aiming at isolating the role of the Canadian EPU *vis-à-vis* the role of bilateral trade for the transmission of US EPU shocks to the Canadian

of the *American Economic Journal: Macroeconomics*. The producer price index has been downloaded from the St. Louis Fed FRED website, while the refiner acquisition cost has been downloaded from the Energy Administration Information website. The reason for using both measures, alternatively, in our robustness checks is due to the possibly different effects of oil shocks obtained by using these different price measures, as highlighted by Hamilton (2003, 2011) and Kilian and Vigfusson (2011) respectively. For this reason, in Figure 5 we label the producer price inedx as "oil Hamilton" and the refiner acquisition cost as "oil Kilian". Both series have been deflated using the CPI. Results are robust to using nominal instead of real prices (evidence available upon request).

economy.

6.1 Generalized Forecast Error Variance Decomposition

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.[24] The first three rows of each panel report the contribution of US EPU, Canada EPU, and monetary policy shocks in our baseline specification, while the fourth row reports the contribution of US EPU shocks in the model augmented with economic uncertainty.[25] We begin by looking at the FEVD conditional to economic busts obtained with our baseline specification. A number of considerations are in order. First, as shown by the first row of the Table, in bad times US EPU shocks explain 26% of the volatility of the Canadian unemployment rate. Hence, EPU spillovers are quantitatively important to explain the dynamics of a key labor market variable such as the unemployment rate. Interestingly, movements in the Canadian EPU index explain about 23% of the Canadian unemployment rate. These numbers points to EPU shocks coming from the US as being as important as domestic Canadian EPU shocks, a result consistent with Klößner and Sekkel's (2014) evidence pointing to policy spillovers from the United States to Canada. Moreover, uncertainty is important in general, given that it is responsible for about 49% of the variation in unemployment at a 2-year horizon. Second, looking at the FEVD in booms shows that the role of uncertainty is relevant in bad times only. Indeed, these figures dramatically drop to 9% (US EPU shocks) and 5% (Canadian EPU shocks) when it comes to explaining unemployment during expansionary phases of the Canadian business cycle. A similar result holds true as regards industrial production, with uncertainty shocks explaining about 8% (US EPU) and 15% (Canadian EPU) in busts, and about 2% and 4% 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, 14%, 17%, and 14% in busts while it ranges from 4% to 6% in booms. Again, independently of the state of the economy, these figures are found to be fairly in line with the contribution

[24] A FEVD analysis focusing on a 12-month horizon delivers very similar results, which are available upon request.

[25] The reason for reporting results from the model augmented with economic uncertainty is that, among all models reported in the robustness checks section, this is the one that returns a less pronounced response of unemployment.

of the Canadian EPU shocks.

Another result of our FEVD analysis regards the drivers of the EPU indices employed in our analysis. As reported in Table 2, about 62% of the volatility of the US EPU index in busts is driven by its own innovation, while the contribution of the Canadian EPU is about 8%. Looking at booms, Canadian EPU explains an even lower share of the US EPU (about 3%), which is instead mostly explained (about 72% of total volatility) by its own shock.[26] Differently, the contribution of US EPU innovations to the volatility of the Canadian EPU index is 32% in busts and 31% in booms. This information is consistent with Granger causality tests conducted with a linear bivariate framework modeling the two EPU indices. Such tests support the causality running from the US EPU index to the Canadian one at any conventional level (the p-value of the null hypothesis of non-causality is 0.00), while they reject the causality running from the Canadian EPU index to the US one (p-value: 0.36).[27] This result supports a novel reading of the role of big countries like the US 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. The figures collected in Table 2 clearly point to a much smaller role played by monetary policy shocks as regards unemployment, with a contribution of about 7% during downturns (about one fourth compared to external uncertainty shocks) and about 3% in booms (vs. 9% by US EPU shocks'). The contribution of monetary policy shocks to the volatility of inflation reads 16% in busts and 14% in booms, 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 36%, much larger than the 5% due to monetary policy shocks. This gap is much smaller in booms, with the former shocks being responsible for about 9% of the variance of the real exchange rate against a contribution of about 3% by monetary policy shocks. Not surprisingly, the main driver of the short-term in-

[26] Notice that here we are referring to the volatility of the EPU indexes, not to that of the innovations to such indexes. Such innovations, which are those we use to compute the GIRFs documented in the previous Section and the FEVD reported in this Section, are - by construction - exogenous under the assumption of our VAR being rich enough from an informational standpoint.

[27] We model a linear VAR(6) as suggested by the Akaike lag-length criterion. Moreover, a simple regression of the Canadian uncertainty index on a constant and lagged US EPU returns an adjusted R^2 of 0.33, a signal of high predictive power of the US EPU index on the Canadian counterpart.

terest 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.

Finally, we check if our findings are robust to the controls we employed to produce the GIRFs documented in Figure 5. Following Baker, Bloom, and Davis (2016), we take the model embedding the broad definition of Economic Uncertainty as a reference. Table 2 documents the contribution of the US EPU shocks to the volatility of the variables in our baseline vector conditional on the EU control. Not surprisingly, the figures corresponding to this scenario point to a more limited role played by US EPU shocks for the volatility of the Canadian unemployment rate. However, such contribution is still as large as 13% in busts, while it is a sixth of it (2%) in booms. This exercise suggests two things. First, US EPU shocks are likely to be a composite of pure policy-related uncertainty shocks and more general economic uncertainty shocks in models that do not feature a broader measure of uncertainty, like the EU indicator. Second, our results still hold when this control is added to our baseline vector. Importantly, our Appendix shows that our main results are robust across all the different models discussed above.

6.2 Transmission mechanism: The uncertainty spillover channel

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, and via this channel affect domestic business cycle indicators, in particular unemployment.[28] An equally plausible transmission channel from US EPU to real activity in Canada would work via trade. Uncertainty in the US would depress domestic consumption and investment, hence US demand of imports from Canada, and via this channel increase unemployment in Canada. The conjecture that fluctuations in uncertainty occurring in the US both foster uncertainty and depress net exports in Canada is confirmed by the impulse responses of Canadian EPU and net exports to a shock to the US EPU index, obtained with our STVAR model augmented with a measure

[28] 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.

of bilateral trade balance, which are shown in Figure 6.[29] The left panel of Figure 6 plots the Canadian EPU impulse response during economic busts. The Canadian index significantly increases after a US EPU shock, before quickly going back to the pre-shock level. The right panel of Figure 6 plots the response of Canadian net exports to US EPU shocks. As one can see, net exports display a significant, short-lived decrease as a consequence of a jump in US economic policy uncertainty.[30] This gives us a broader picture of the effects of a shock to the level of US economic policy uncertainty on Canada: a US policy uncertainty shock triggers an increase in Canadian policy uncertainty, a decrease in Canadian net exports to the US, *and* a temporary downturn of real activity. One possible interpretation of these facts is that there might be two alternative, not mutually exclusive, transmission mechanisms at work simultaneously, i.e., one that transmits US EPU shocks to the Canadian real economy via uncertainty spillovers, the other that works via trade.

We shed light on the relative role played by the Canadian EPU index and bilateral trade in transmitting US policy-related shocks by conducting two counterfactual simulations. In the first one, the Canadian EPU index is not allowed to respond to systematic movements in US EPU, while net exports are left free to react. In the second one, it is Canadian net exports that do not react to US EPU fluctuations, while Canadian EPU is left free to respond. If the main driver of unemployment fluctuations in Canada is the response of Canadian economic policy uncertainty to US EPU, and not US economic policy uncertainty *per se* or via its impact on net exports, then the first counterfactual should return a more moderate responses of the Canadian unemployment rate to a US EPU shock compared to both the baseline response of unemployment and the counterfactual one obtained by switching off the "trade channel".

Figure 7 summarizes the results obtained by estimating the baseline version of our STVAR as in Section 4, augmented with both economic uncertainty, ordered second, and net exports, ordered last. The Figure reports, conditional to economic busts, the response of Canada's unemployment rate to a one standard deviation US EPU shock

[29] The series of US imports from Canada (IMPCA) and US exports to Canada (EXPCA) are downloaded from the FRED database. Real net exports are defined as the difference between imports and exports divided by US CPI. The variable is ordered last in our VAR. The estimated VAR includes also economic uncertainty, ordered second.

[30] The analysis is aimed at describing the transmittion channel of US EPU shocks in Canada during periods of slack. Therefore, for the sake of clarity of exposition, we do not plot the responses of Canadian EPU and net exports in booms. However, the full set of responses in booms and busts, available upon request, confirms our baseline findings: both Canadian EPU and net exports react more strongly to US EPU shocks in recessionary times.

in four scenarios: the full model; the model without net exports; the counterfactual response obtained by muting only the response of Canada EPU to US EPU; the counterfactual response obtained by muting only the response of net exports to US EPU. Two main findings arise. First, adding net exports to the baseline model leaves the peak response of unemployment virtually unaffected, while it makes its response milder at larger horizons. Second, the peak response of unemployment is halved in the counterfactual scenario in which the Canadian EPU index does not respond to movements in US EPU, while it remains virtually unchanged in the counterfactual scenario where net exports do not react to systematic movements in US EPU.[31] Overall, the results from both counterfactuals suggests that the negative spillover of US economic policy uncertainty shocks on real activity in Canada is mainly due to the reaction of Canadian economic policy uncertainty, while the trade channel plays a minor role. We interpret this evidence as consistent with the "economic policy uncertainty spillovers channel".

7 EPU spillovers: The case of the UK

The results obtained so far document a significant economic uncertainty spillover effect originating in the United States for the Canadian economy. It is of interest to investigate whether this finding is specific to Canada, or rather it applies also to other economies, not necessarily as much integrated with the US as Canada is. One interesting alternative case is provided by the UK for two main reasons. First, the UK is a much larger economy compared to Canada, so in principle less prone to spillovers of shocks originating in other countries. According to the IMF, the UK GDP in 2015 was equal to 2,849,345 millions of US$, almost twice the size of Canadian GDP, which was equal to 1,552,386 millions (again, measured by US$). Second, despite sharing similar characteristics with the United States, it is far from being as much trade integrated as Canada. In terms of bilateral trade with the US, in 2015 11% of UK total exports were imported by the Unites States (74% the share for Canada), and 6.7% of UK imports came from the United States (55% the share for Canada).

We then replicate the analysis conducted in Section 4 for Canada with UK data. We estimate our nonlinear framework (1)-(3) to model the following vector: $\boldsymbol{X}_t = [EPU_t^{US}, EPU_t, \overline{\Delta IP}_t, u_t, \pi_t, R_t, \Delta\epsilon_t]'$, where EPU_t^{US} is the US EPU uncertainty index,

[31] If anything, shutting down net exports makes the response of unemployment more persistent at longer horizons. This finding is consistent with Charnavoki and Dolado (2014), who document a Dutch-disease type of effect at business cycle frequencies for Canada.

while the remaining variables, which refer to the United Kingdom, are the country-specific economic policy uncertainty index EPU_t, a six-month moving average of the monthly growth rate of industrial production (percentualized and annualized) $\overline{\Delta IP}_t$, the unemployment rate u_t, the CPI inflation rate π_t (y-o-y percentualized growth rate of the monthly index), the policy rate R_t, and the bilateral real exchange rate $\Delta\epsilon_t^{UK} \equiv \pi_t^{US} + \Delta s_t^{UK,US} - \pi_t^{UK}$ constructed by considering the inflation rates in the two countries and combining it with $\Delta s_t^{UK,US}$, which is the y-o-y growth rate of the UK/US nominal exchange rate.[32] The sample size, dictated by the availability of our proxy for the policy rate, is 1959:M1-2014:M10

Figure 8 plots the response of industrial production, unemployment, inflation and the policy rate to a one-standard deviation shock in US EPU conditional on booms and busts in the United Kingdom.[33] An unexpected increase in US economic policy uncertainty triggers a negative response in UK real activity. As for the case of Canada, such response is larger if the shock originated in the United States when the UK economy was already in a period of slack. Inflation reacts negatively, with some lags, again more markedly in economic bad times. Monetary policy is found to induce a reduction in the interest rate, and such monetary policy response is stronger in busts. Figure 9 plots the difference between the impulse responses. The differences in the reaction of both unemployment and industrial production, as well as that of the policy rate, are found to be statistically significant at 68% confidence level. Overall, the results for the UK confirm that economic policy uncertainty shocks originating in the US can spillover onto other economies and trigger an asymmetric response of real activity depending on the stance of the business cycle. Hence, evidence in favor of the economic policy uncertainty spillovers channel documented in the previous Section is not confined to a small-open economy like Canada that is linked to the US by an intense trading activity. Indeed, we find that it could very well be a relevant transmission channel of US EPU shocks also for bigger and more trade-independent economies (in terms of relationship with the US) such as the United Kingdom.

[32] For the United Kingdom, the policy rate is the discount rate. The correlation between this rate and the 3-month rate on UK Treasury securities reads 0.99 in our sample. All the UK series were downloaded from the Federal Reserve Bank of St. Louis' website.

[33] Our results do not depend on the specific choice of a six-lags moving average of the monthly growth rate of industrial production, i.e., they are robust to alternative modeling choices (evidence available upon request).

8 Conclusions

We investigate the spillover effects of a jump in US economic policy uncertainty for the Canadian business cycle. 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 during periods of busts experienced by the countries that receive the external uncertainty shocks. 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 to US EPU signal the existence of an "economic policy uncertainty spillover channel", i.e., spikes in US economic policy uncertainty foster uncertainty in Canada and, via this channel, lead to a temporary slowdown of Canada's real activity. This result is shown to be robust to the possible presence of a simultaneous "trade channel", working via fluctuations in bilateral trade. Finally, similar analysis conducted for the UK confirms that US EPU shocks are likely to be among the drivers of unemployment when bad times are in place also in economies that are relatively larger and less integrated with the United States.

From a policy perspective, our evidence suggests that uncertainty about future policy actions in influential countries like the US may be costly not only for such countries but can importantly spillover on to small-open economies like Canada. As discussed by Davis (2015), the large increase in the number of norms and regulations that the US 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 manage 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 US of implementing the policies suggested by Davis (2015) and Baker, Bloom, and Davis (2016) may be larger than those typically estimated when considering the US case in isolation.

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Date	Event	Baseline	Dummy
Oct. 1987	Black Monday	X	X
Jan. 1991	Gulf War I	X	
Nov. 2000	Bush election	X	X
Sep. 2001	9/11	X	X
Mar. 2003	Iraq invasion	X	
Jan. 2008	Large interest rate cuts	X	X
Sep. 2008	Lehman Brothers' bankruptcy	X	X
Oct. 2008	Global Financial Crisis	X	
July 2010	Mid-term elections	X	X
Sep. 2010	Mid-term elections	X	X
July 2011	Debt Ceiling	X	X
Aug. 2011	Debt Ceiling	X	X
Nov. 2012	Fiscal cliff	X	X
Oct. 2013	Government shutdown	X	X

Table 1: **Major US and Global Economic Policy Uncertainty Shocks.** Baseline: Dates corresponding to positive realizations of the estimated shocks exceeding 2 standard deviations according to our baseline model. Dummy: Dates selected by focusing on domestic (US) events only.

					Busts			
Shock/Variable		EPU_t^{US}	EPU_t	ΔIP_t	u_t	π_t	R_t	$\Delta\epsilon_t$
$\widetilde{\varepsilon}^{EPU_t^{US}}$		0.62	**0.32**	0.08	**0.26**	0.14	0.17	0.14
$\widetilde{\varepsilon}^{EPU_t}$		**0.08**	0.36	0.15	**0.23**	0.10	0.14	0.22
$\widetilde{\varepsilon}^{R_t}$		*0.04*	*0.06*	*0.07*	*0.07*	*0.16*	*0.51*	*0.05*
$\widetilde{\varepsilon}^{EPU_t^{US}}$	*EU*	0.35	**0.15**	0.03	**0.13**	0.08	0.05	0.05

					Booms			
Shock/Variable		EPU_t^{US}	EPU_t	ΔIP_t	u_t	π_t	R_t	$\Delta\epsilon_t$
$\widetilde{\varepsilon}^{EPU_t^{US}}$		0.72	**0.31**	0.02	**0.09**	0.04	0.06	0.04
$\widetilde{\varepsilon}^{EPU_t}$		**0.03**	0.44	0.04	**0.05**	0.05	0.09	0.05
$\widetilde{\varepsilon}^{R_t}$		*0.02*	*0.05*	*0.10*	*0.03*	*0.14*	*0.62*	*0.03*
$\widetilde{\varepsilon}^{EPU_t^{US}}$	*EU*	0.33	**0.08**	0.00	**0.02**	0.02	0.01	0.01

					Linear			
Shock/Variable		EPU_t^{US}	EPU_t	ΔIP_t	u_t	π_t	R_t	$\Delta\epsilon_t$
$\widetilde{\varepsilon}^{EPU_t^{US}}$		0.88	**0.35**	0.02	**0.17**	0.05	0.09	0.06
$\widetilde{\varepsilon}^{EPU_t}$		**0.02**	0.55	0.04	**0.07**	0.07	0.12	0.06
$\widetilde{\varepsilon}^{R_t}$		*0.01*	*0.03*	*0.06*	*0.05*	*0.12*	*0.67*	*0.01*
$\widetilde{\varepsilon}^{EPU_t^{US}}$	*EU*	0.45	0.15	0.01	0.02	0.01	0.02	0.00

Table 2: **Forecast Error Variance Decomposition: US 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, with the exception of those in the fourth line of each case (Busts, Booms, Linear), which refer to the model with the Economic Uncertainty measure placed after the US EPU index in the vector.

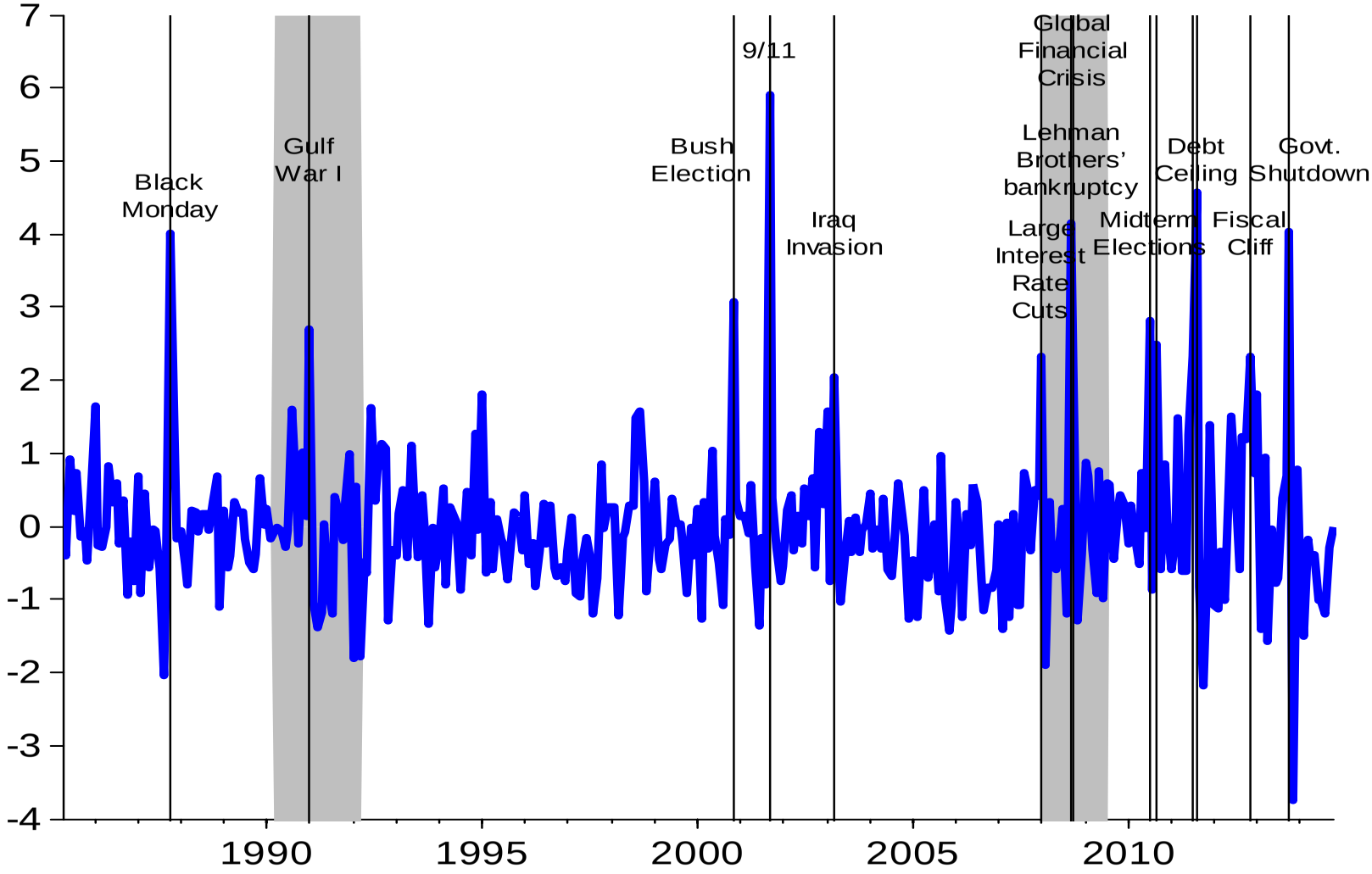

Figure 1: **US EPU shocks**. Blue line: US EPU shocks estimated with the US-Canada STVAR as explained in the text. Black vertical lines: Realizations of the shock larger than two standard deviations. Grey vertical bars: Canadian recessions as dated by the ECRI.

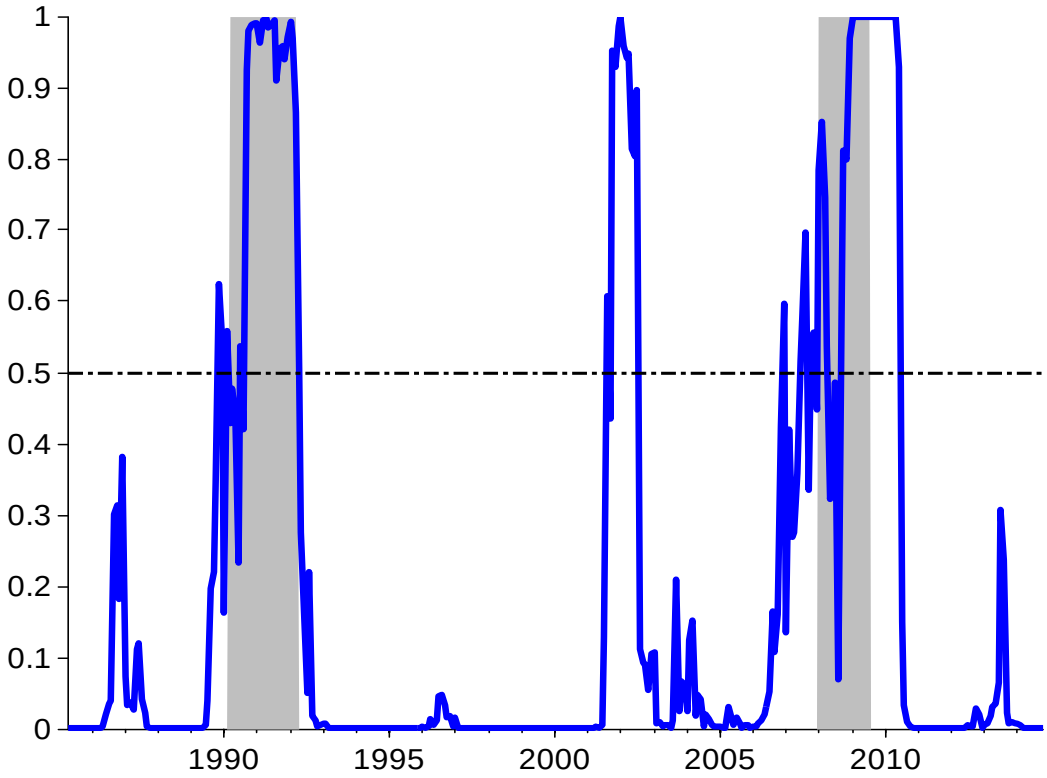

Figure 2: **Probabilities of Economic Busts for Canada as Estimated by the STVAR model.** Sample: 1985:M1-2014:M10**.** Function [1-F(z)] estimated jointly with the baseline STVAR model. Transition indicator z: 18-month moving average of the monthly growth rate of the Canadian industrial production index. Grey vertical bars indicate recessions as dated by the Economic Cycle Research Institute.

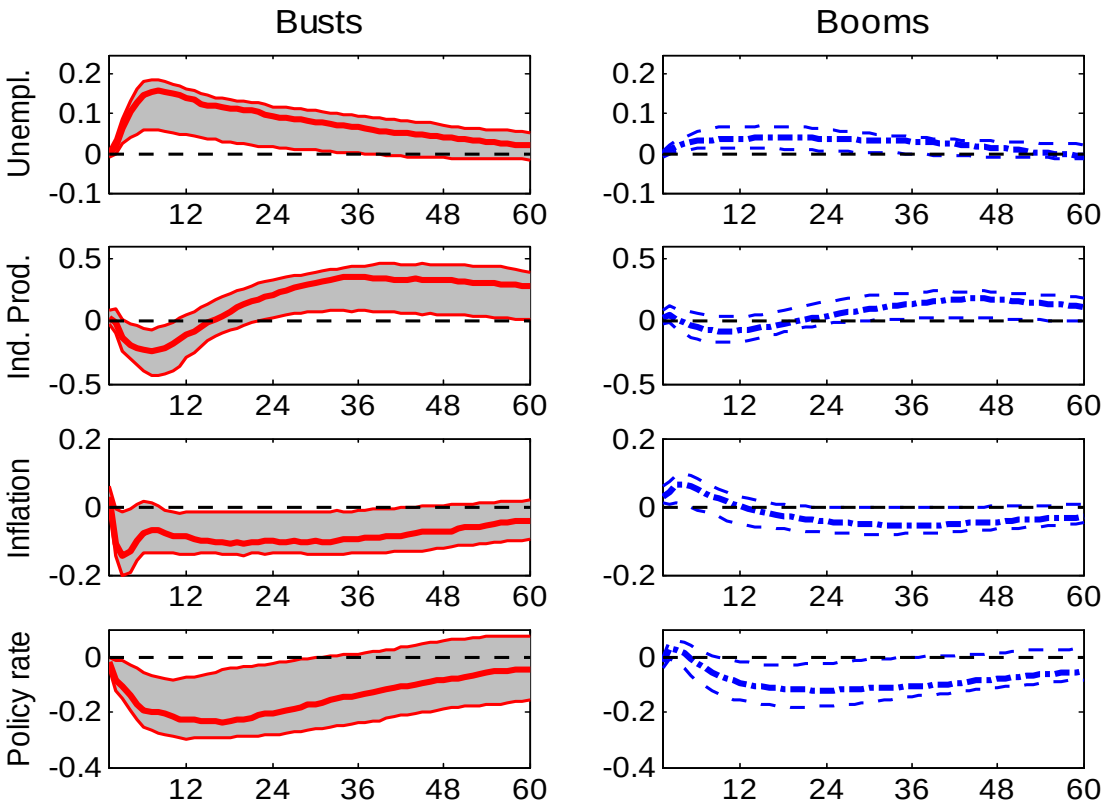

Figure 3: **Effects of a shock to the US EPU Index on the Canadian economy.** Sample: 1985:M1-2014:M10. Median generalized impulse responses to a one-standard deviation shock to the US EPU index hitting the Canadian economy in busts (red solid line) and booms (blue dashed-dotted line). 68% confidence 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.

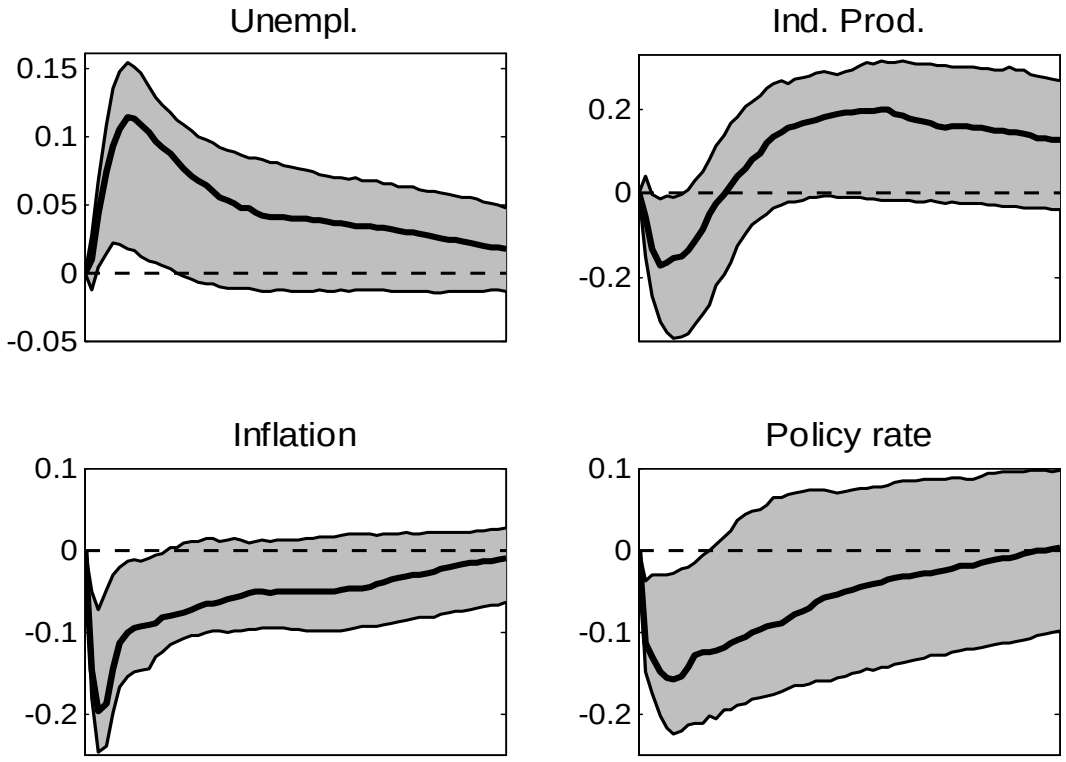

Figure 4: **Effects of a shock to the US EPU Index on the Canadian economy: Difference between states.** Sample: 1985:M1-2014:M10. Differences between median generalized impulse responses in busts and booms to a one-standard deviation shock to the US EPU Index. Median realizations identified via black 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.

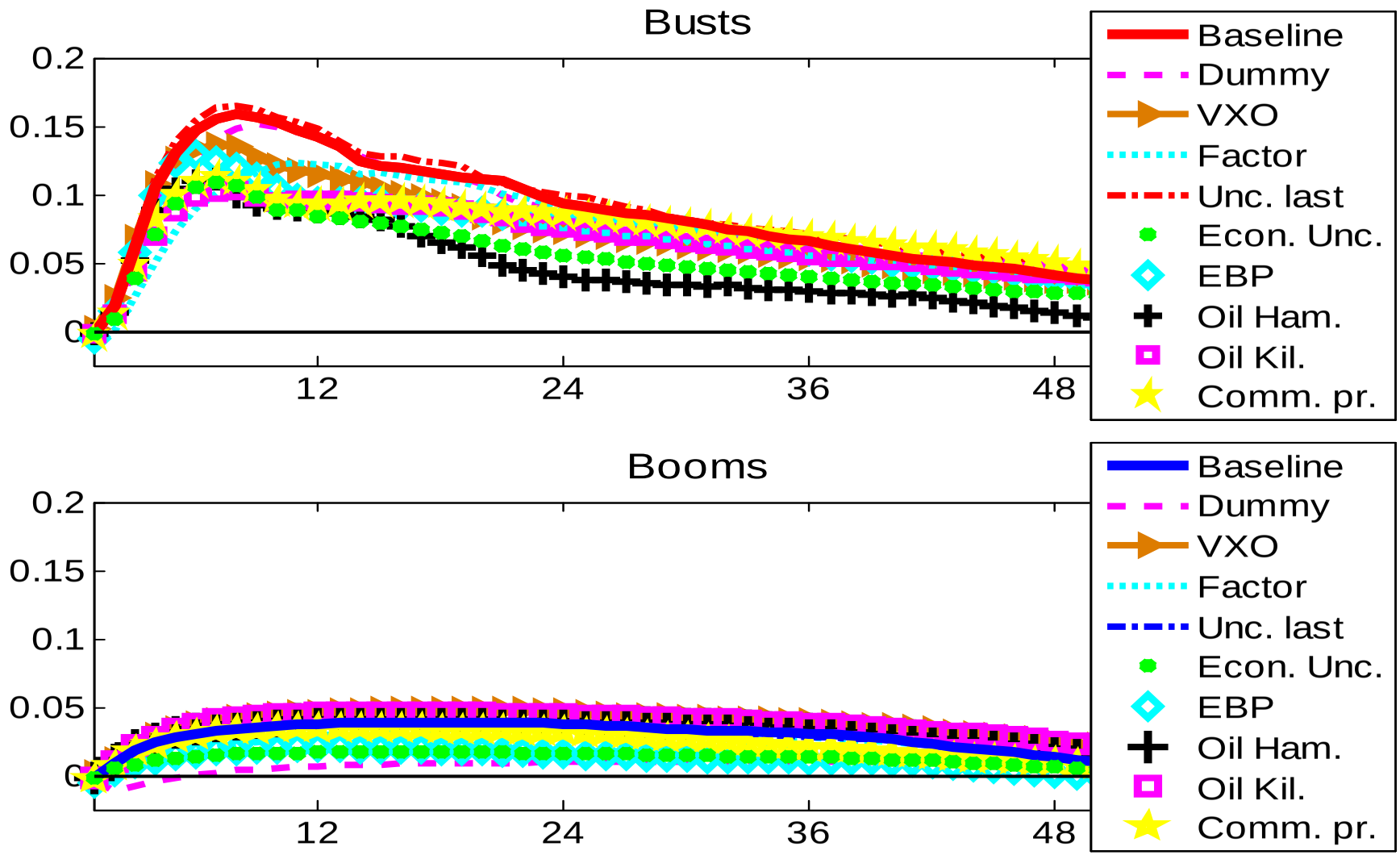

Figure 5: **Response of the Canadian unemployment rate to an EPU shock originating in the US: Robustness checks**. Responses of Canadian unemployment to a one standard deviation US EPU shock for the models discussed in Section 5.

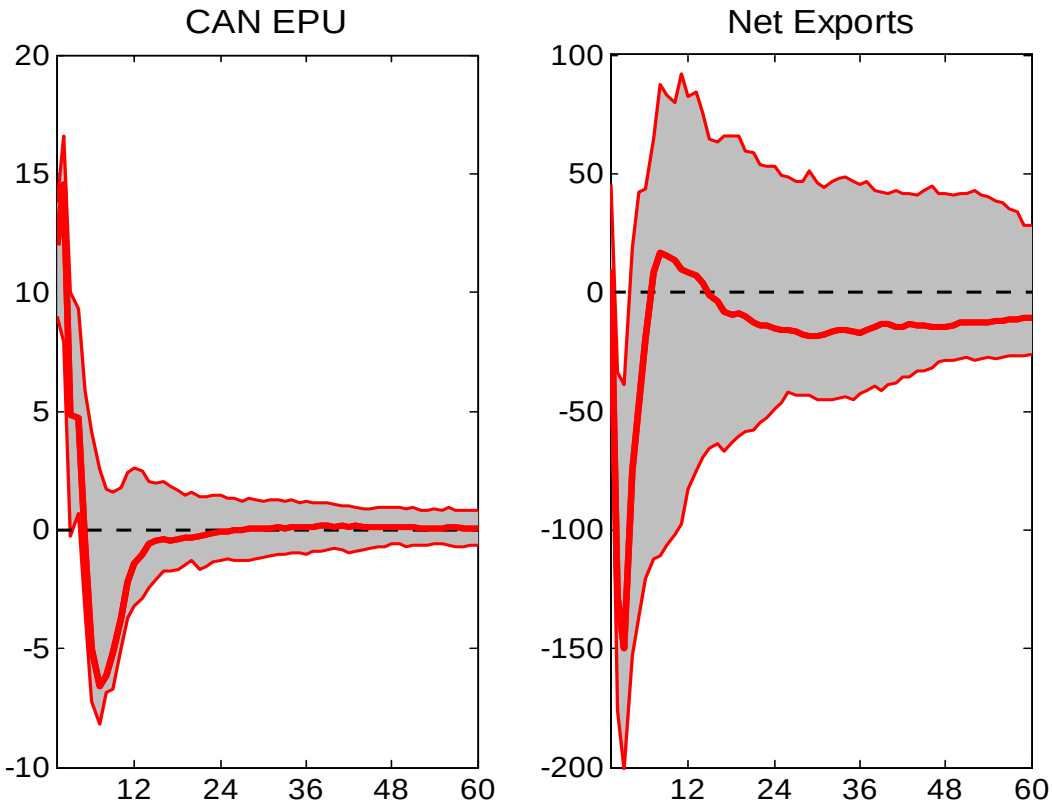

Figure 6: **Responses of Canadian EPU and net exports to US EPU shocks.** Generalized impulse responses to a one-standard deviation shock to the US EPU index hitting the Canadian economy in busts. 68% confidence bands denoted by shaded areas. Left panel: Canadian EPU. Right panel: net exports. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production index. Model featuring economic uncertainty ordered second and net exports ordered last in the vector.

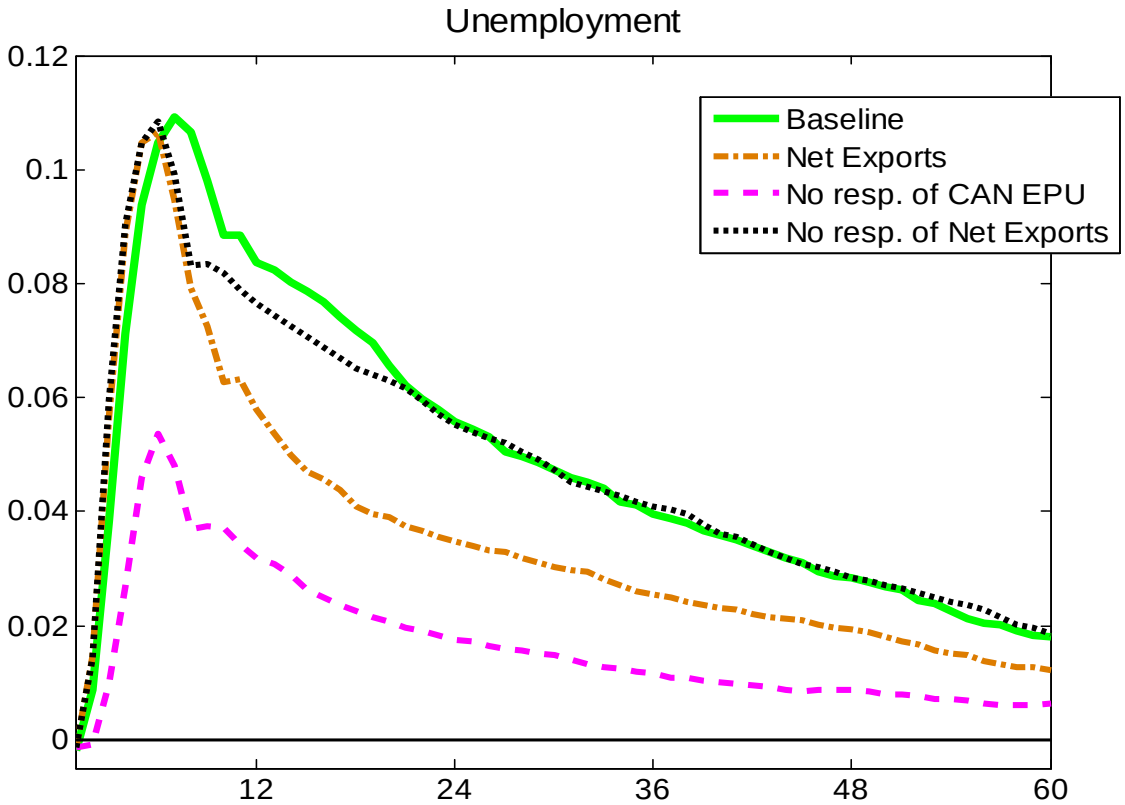

Figure 7: **Effects of a shock to the US EPU index on the Canadian economy: Role of the uncertainty and trade channels.** Sample: 1985:M1-2014:M10. Generalized impulse responses of the Canadian unemployment rate in busts. Solid green line: response of unemployment in the baseline model without net exports (labeled "Baseline"). Dashed-dotted brown line: response of unemployment in the baseline model including net exports (labeled "Net Exports"). Dashed purple line: couterfactual impulse response of unemployment obtained by zeroing the coefficients of Canada EPU to US EPU in the equation modeling Canada EPU (labeled "No resp. of CAN EPU"). Dotted black line: couterfactual impulse response of unemployment obtained by zeroing the coefficients of net exports to US EPU in the equation modeling net exports (labeled "No resp. of CAN EPU"). All models featuring Economic Uncertainty ordered second in the vector as a control.

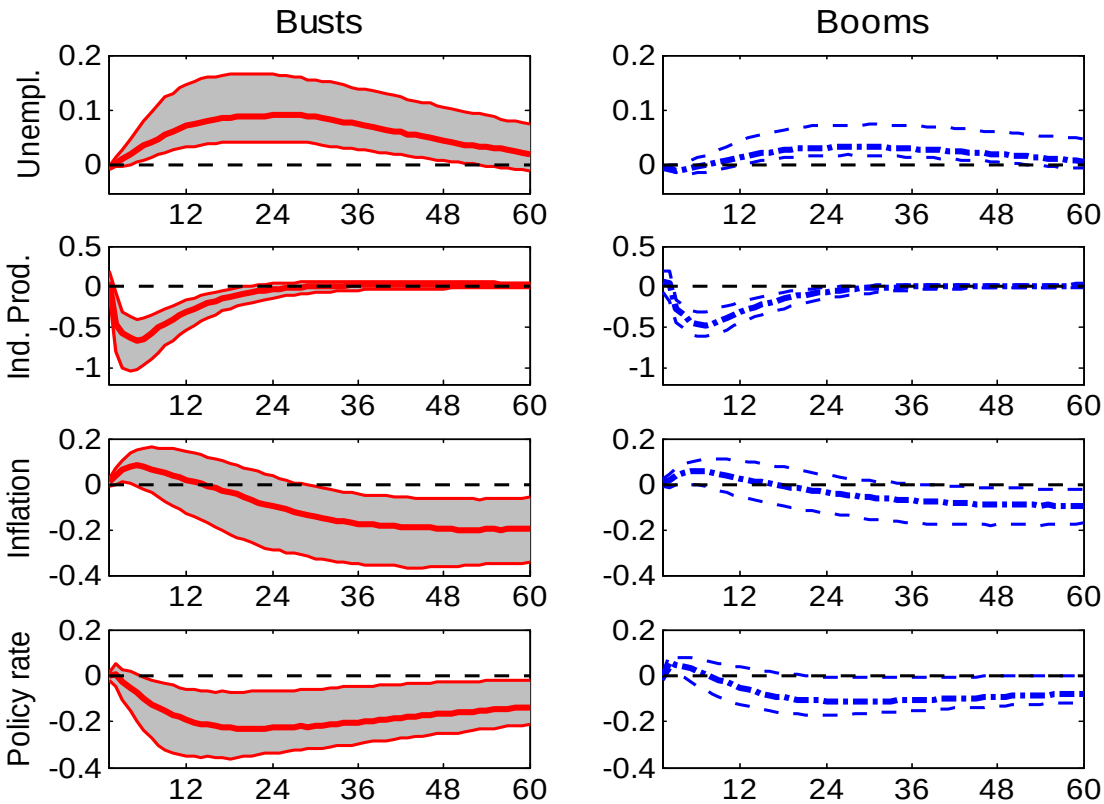

Figure 8: **Effects of a shock to the US EPU index on the UK economy.** Sample: 1959:M1-2014:M10. Median generalized impulse responses to a one-standard deviation shock to the US EPU index hitting the UK economy in busts (red solid line) and booms (blue dashed-dotted line). 68% confidence intervals denoted by shaded areas (busts) and dashed blue lines (booms). Transition indicator for the UK: 6-term moving average of the monthly growth rate of the UK industrial production index.

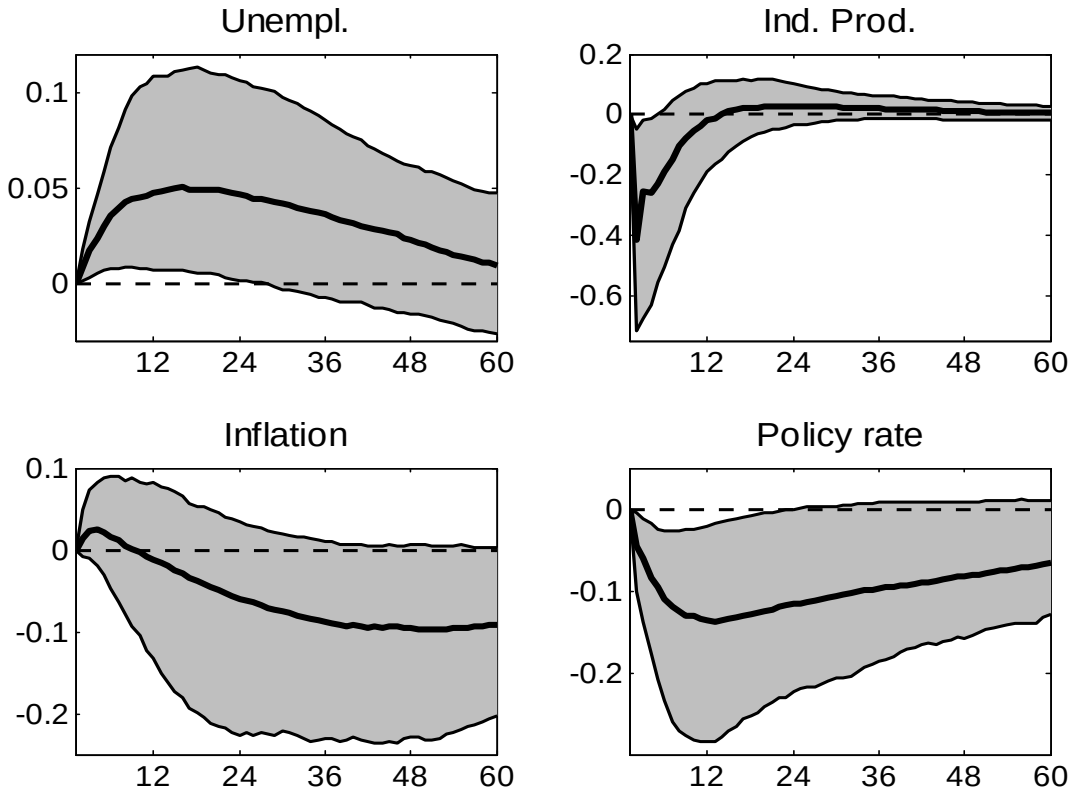

Figure 9: **Effects of a shock to the US EPU index on the UK economy: Difference between states.** Sample: 1959:M1-2014:M10. Differences between median generalized impulse responses in booms and busts to a one-standard deviation shock to the US EPU index. Median realizations denoted with black lines, 68% confidence intervals denoted with shaded areas. Transition indicator for the UK: 6-term moving average of the monthly growth rate of the UK industrial production index.

Appendix of the paper "Economic Policy Uncertainty Spillovers in Booms and Busts" by Giovanni Caggiano, Efrem Castelnuovo, and Juan Manuel Figueres

This Appendix reports further details about: the linearity test used to discriminate between a linear vs. a Smooth Transition VAR model; the computation algorithm of the GIRFs; the computation of the Generalized FEVD; a detailed discussion of our dating of the Canadian business cycle; extra robustness checks involving two different transition indicators alternative to the one used in the exercises documented in the main text; a robustness check involving a longer sample for Canada; and the Generalized FEVD for all robustness checks.

Linearity test. We test the null hypothesis of a linear VAR vs. the alternative of a Smooth Transition VAR model with a single transition variable using the LM testing procedure proposed by Teräsvirta and Yang (2014). Consider the following $p-$dimensional 2-regime $n-$th order Taylor approximation of the baseline logistic STVAR model:

$$\mathbf{X}_t = \mathbf{\Theta}_0' \mathbf{Y}_t + \sum_{i=1}^{n} \mathbf{\Theta}_i' \mathbf{Y}_t z_t^i + \boldsymbol{\varepsilon}_t \tag{A1}$$

where $\mathbf{X}_t$ is the $(p \times 1)$ vector of endogenous variables included in the baseline specification, $\mathbf{Y}_t = [\mathbf{X}_{t-1}| \ldots |\mathbf{X}_{t-k}|\boldsymbol{\alpha}]$ is the $((k \times p + q) \times 1)$ vector of exogenous variables (including endogenous variables lagged k times and a column vector of constants $\boldsymbol{\alpha}$), z_t is the transition variable, and $\mathbf{\Theta}_0$ and $\mathbf{\Theta}_i$ are matrices of parameters. In our case, the number of endogenous variables is $p = 7$, the number of exogenous variables is $q = 1$, and the number of lags is fixed to $k = 1$ to overcome the "curse of dimensionality" , as indicated in Teräsvirta and Yang (2014). Under the null hypothesis of linearity, $\mathbf{\Theta}_i = \mathbf{0}$ $\forall i$.

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

1. Estimate the restricted model $(\mathbf{\Theta}_i = \mathbf{0}, \forall i)$ by regressing $\mathbf{X}_t$ on $\mathbf{Y}_t$. Collect the residuals $\tilde{\mathbf{E}}$ and the matrix residual sum of squares $\mathbf{RSS}_0 = \tilde{\mathbf{E}}'\tilde{\mathbf{E}}$.

2. Run an auxiliary regression of $\tilde{\mathbf{E}}$ on $(\mathbf{Y}_t, \mathbf{Z}_n)$ where $\mathbf{Z}_n \equiv [\mathbf{Z}_1|\mathbf{Z}_2| \ldots |\mathbf{Z}_n] = [\mathbf{Y}_t' z_t|\mathbf{Y}_t' z_t^2| \ldots |\mathbf{Y}_t' z_t^n]$. Collect the residuals $\tilde{\mathbf{\Xi}}$ and compute the matrix residual sum of squares $\mathbf{RSS}_1 = \tilde{\mathbf{\Xi}}'\tilde{\mathbf{\Xi}}$.

3. Compute the test-statistic

$$\begin{aligned} LM &= Ttr\left\{\mathbf{RSS}_0^{-1}\left(\mathbf{RSS}_0 - \mathbf{RSS}_1\right)\right\} \\ &= T\left(p - tr\left\{\mathbf{RSS}_0^{-1}\mathbf{RSS}_1\right\}\right) \end{aligned}$$

Under the null hypothesis, the test statistic is distributed as a χ^2 with $np\left(kp+q\right)$ degrees of freedom.[1] For our model, we set $n = 3$, as suggested by Luukkonen, Saikkonen, and Teräsvirta (1988), and get a value of $LM = 241.12$ with a corresponding p-value equal to 0.00. Hence, we reject the null hypothesis of a linear specification versus the alternative of a STVAR. Results are robust to fixing the order of the Taylor approximation to $n < 3$.

Generalized Impulse Response Functions and confidence bands. We compute the Generalized Impulse Response Functions from our STVAR model by following the approach proposed by Koop, Pesaran, and Potter (1996). The algorithm features the following steps:

1. Given all available observations, with sample size $t = 1985 : M1, \ldots, 2014 : M10$, construct the set of all possible histories $\{\boldsymbol{\lambda}_{t-1,i} \in \boldsymbol{\Lambda}\}$ of length l, where l is the number of lags of the STVAR and $\boldsymbol{\lambda}_{t-1,i} = \{\mathbf{X}_{t-1}, \ldots, \mathbf{X}_{t-l}; z_{t-1}\}$. Then, $\boldsymbol{\Lambda}$ will contain $T - l$ histories $\boldsymbol{\lambda}_{t-1,i}$, with $T = 358$ and $l = 2$.[2]

2. Separate the set of all recessionary (busts) histories from that of all expansionary (booms) histories. Given the estimated threshold, $\widehat{c}$, for each $\boldsymbol{\lambda}_{t-1,i} \in \boldsymbol{\Lambda}$, if $z_{\boldsymbol{\lambda}_{t-1,i}} \leq \widehat{c}$, then $\boldsymbol{\lambda}_{t-1,i} \in \boldsymbol{\Lambda}^R$, where $\boldsymbol{\Lambda}^R$ is the set of all recessionary histories; if $z_{\boldsymbol{\lambda}_{t-1,i}} > \widehat{c}$, then $\boldsymbol{\lambda}_{t-1,i} \in \boldsymbol{\Lambda}^E$, where $\boldsymbol{\Lambda}^E$ is the set of all expansionary histories.[3]

3. Select at random one history $\boldsymbol{\lambda}_{t-1,i}$ from the set $\boldsymbol{\Lambda}^R$. Then, draw randomly with replacement from the empirical distribution of the residuals $\widehat{\boldsymbol{\varepsilon}}_t$, and get $\widehat{\boldsymbol{\varepsilon}}^{(j)*} = \left\{\widehat{\boldsymbol{\varepsilon}}_t^*, \widehat{\boldsymbol{\varepsilon}}_{t+1}^*, \ldots, \widehat{\boldsymbol{\varepsilon}}_{t+h}^*,\right\}$, where h is the maximum horizon of interest for the GIRFs and $\widehat{\boldsymbol{\varepsilon}}_{t+i}^*$ is a column vector of residuals of size p, where $p = 7$ is the dimension of the vector of endogenous variables of the baseline STVAR model.

[1] Notice that, since the transitional indicator is endogenous in our case, we do not include in $\mathbf{Z}_n$ the vector of constant terms $\boldsymbol{\alpha}$ to avoid perfect collinearity problems. As a consequence, the number of degrees of freedom is equal to np times the column dimension of $\mathbf{Z}_n$, and is equal to 147.

[2] The number of lags of the STVAR has been selected according to the AIC.

[3] The estimated threshold is $\widehat{c} = -0.78$. The estimated value of the slope parameter is $\widehat{\gamma} = 6.36$.

4. Orthogonalize the bootstrapped residuals to recover the structural shocks:

$$\mathbf{e}^{(j)^*} = \widehat{\mathbf{C}}^{-1}\widehat{\boldsymbol{\varepsilon}}^{(j)^*}. \tag{A2}$$

where $\widehat{\mathbf{C}}$ is the Cholesky factor of the residuals' variance-covariance matrix $\widehat{\boldsymbol{\Omega}}$.

5. Form another set of bootstrapped shocks, $\mathbf{e}^{(j)^\delta}$, which will be equal to $\mathbf{e}^{(j)^*}$ except for the first element of the first column, corresponding to the US EPU uncertainty shock at time t, which will be equal to the corresponding element in $\mathbf{e}^{(j)^*}$ plus δ, where δ is set to one-standard deviation of the orthogonalized residuals: $e_{(1,1)}^{(j)^\delta} = e_{(1,1)}^{(j)^*} + \delta$.

6. Transform back $\mathbf{e}^{(j)^*}$ and $\mathbf{e}^{(j)^\delta}$, and get the bootstrapped residuals:

$$\widehat{\boldsymbol{\varepsilon}}^{(j)^*} = \widehat{\mathbf{C}}\mathbf{e}^{(j)^*} \tag{A3}$$

and

$$\widehat{\boldsymbol{\varepsilon}}^{(j)^\delta} = \widehat{\mathbf{C}}\mathbf{e}^{(j)^\delta}. \tag{A4}$$

7. Conditional on the initial history $\boldsymbol{\lambda}_{t-1,i}$, use $(A3)$ and $(A4)$ to simulate the evolution of all the variables incorporated in the vectors $\mathbf{X}_{\boldsymbol{\lambda}_{t-1,i}}^{(j)*}$ and $\mathbf{X}_{\boldsymbol{\lambda}_{t-1,i}}^{(j)\delta}$ - endogenous transition indicator included - and compute the GIRF as:

$$GIRF\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}\right)^{(j)} = \mathbf{X}_{\boldsymbol{\lambda}_{t-1,i}}^{(j)^\delta} - \mathbf{X}_{\boldsymbol{\lambda}_{t-1,i}}^{(j)^*}.$$

8. Conditional on history $\boldsymbol{\lambda}_{t-1,i}$, repeat for $j = 1,\ldots,B$ vectors of bootstrapped residuals and get $GIRF^{(j)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}\right)$. Set $B = 500$.

9. Calculate the GIRF for $\boldsymbol{\lambda}_{t-1,i}$ as

$$\widehat{GIRF}^{(i)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}\right) = B^{-1}\sum_{j=1}^{B} GIRF^{(i,j)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}\right). \tag{A5}$$

10. Repeat steps 3 to 9 for $i = 1,\ldots,N = 500$ histories belonging to the set of recessionary histories, $\boldsymbol{\lambda}_{t-1,i} \in \boldsymbol{\Lambda}^R$, and get $\widehat{GIRF}^{(i,R)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}\right)$, where R denotes explicitly that we are conditioning upon recessionary (busts) histories.

11. Compute the recessionary GIRF as:

$$\widehat{GIRF}^{(R)}\left(h,\delta,\boldsymbol{\Lambda}^R\right) = N^{-1}\sum_{i=1}^{N}\widehat{GIRF}^{(i,R)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}\right).$$

12. Repeat all previous steps - 3 to 11 - for 500 histories belonging to the set of all expansions (booms) and get $\widehat{GIRF}^{(E)}\left(h,\delta,\boldsymbol{\Lambda}^{E}\right)$.

13. Estimate the confidence bands as follows. Generate a set of artificial data $\mathbf{Y}_t^*$ via the bootstrap procedure proposed by Kilian (1998). Use $\mathbf{Y}_t^*$ to estimate a STVAR model. Repeat steps $3-8$ for recessions (busts) and expansions (booms) and store the average realization $\widehat{GIRF}^{\left(Y_i^*,R\right)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}^*\right)$ and $\widehat{GIRF}^{\left(Y_i^*,E\right)}\left(h,\delta,\boldsymbol{\lambda}_{t-1,i}^*\right)$, respectively. Repeat this step for 500 sets of artificial data $\mathbf{Y}_t^*$. Compute the confidence bands by taking the 14th and the 86th percentiles of the empirical densities for each regime.

Generalized Forecast Error Variance Decomposition.

We calculate the FEVD for our STVAR model following the procedure proposed by Lanne and Nyberg (2016).

1. Draw a sequence of reduced form residuals $\boldsymbol{\varepsilon}^{(j)}$. Select an initial history $\boldsymbol{\lambda}_{t-1,i}$ from the set of recessionary (busts) histories $\boldsymbol{\Lambda}^{R}$.

2. Conditional on $\boldsymbol{\varepsilon}^{(j)}$ and $\boldsymbol{\lambda}_{t-1,i}$, compute $GIRF_{k,i}\left(h,\delta_{kt},\boldsymbol{\lambda}_{t-1,i}\right)^{(j)}$, where $h=1,\ldots,H$ is the horizon of interest, δ_{kt} denotes the shock to variable $k=1,\ldots,K$, and K is the number of endogenous variables. Set $\delta_{kt}=1$.

3. The contribution of shock k_1 to the forecast error variance of variable k_2 at horizon H, i.e. the GFEVD, conditional on $\boldsymbol{\varepsilon}^{(j)}$ and $\boldsymbol{\lambda}_{t-1,i}$ is given by:

$$\omega_{k_1k_2}\left(H,\boldsymbol{\lambda}_{t-1,i}\right)^{(j)}=\frac{\sum_{h=1}^{H}\left[GIRF_{k_2}\left(h,\delta_{k_1t},\boldsymbol{\lambda}_{t-1,i}\right)^{(j)}\right]^2}{\sum_{k=1}^{K}\sum_{h=1}^{H}\left[GIRF_{k_2}\left(h,\delta_{kt},\boldsymbol{\lambda}_{t-1,i}\right)^{(j)}\right]^2}$$

4. Repeat steps $2-3$ for $j=1,\ldots,J$ vectors of bootstrapped residuals $\boldsymbol{\varepsilon}^{(j)}$, thus generating J different $\omega_{k_1k_2}\left(H,\boldsymbol{\lambda}_{t-1,i}\right)^{(j)}$. Set $J=500$.

5. Obtain the GFEVD for shock k_1 and variable k_2 conditional on $\boldsymbol{\lambda}_{t-1,i}$ as:

$$\widehat{GFEVD}_{k_1,k_2}\left(H,\boldsymbol{\lambda}_{t-1,i}\right)=J^{-1}\sum_{j=1}^{J}\omega_{k_1k_2}\left(H,\boldsymbol{\lambda}_{t-1,i}\right)^{(j)}.$$

6. Repeat steps $1-5$ for $i=1,\ldots,I$ initial histories $\boldsymbol{\lambda}_{t-1,i}\in\boldsymbol{\Lambda}^{R}$, and get I values for $\widehat{GFEVD}_{k_1,k_2}\left(H,\boldsymbol{\lambda}_{t-1,i}\right)$.

7. The GFEVD for shock k_1 and variable k_2 in recessions (busts) is then given by:

$$\widehat{GFEVD}_{k_1,k_2}\left(H,\mathbf{\Lambda}^R\right) = I^{-1}\sum_{i=1}^{I}\widehat{GFEVD}_{k_1,k_2}\left(H,\boldsymbol{\lambda}_{t-1,i}\right).$$

8. Obtain the GFEVD for shock k_1 and variable k_2 in expansions (booms), $\widehat{GFEVD}_{k_1,k_2}\left(H,\mathbf{\Lambda}^E\right)$, by repeating steps $1-7$ conditioning on histories belonging to the set of expansions (booms).

Discussion on the dating of the Canadian Business cycle. As documented in Section 4 of the paper, our estimated logistic function point to a high probability of being in a recession for Canada in the early 2000s. However, according to the ECRI, such period is not a 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 A1, 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 (whose sign is switched in Figure A1 to ease the comparison with the evolution of industrial production and real GDP) 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%). The evolution of the employment rate confirms that the early 2000s slowdown affected the Canadian labor market less than in the two occasions classified as recessions. 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".

Alternative transition indicators. 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.[4] We then estimate two models which use - alternatively - two different transition indicators. The first model employs a moving average of the real GDP growth rate to replace industrial production in our VAR. The second model employs the common factor computed via a principal component analysis which considers four different business cycle indicators, i.e., the growth rates of industrial production and real GDP and the rates of unemployment and employment. Figure A2 shows that our results are robust to the employment of these alternative transition indicators.

Initial conditions to identify booms and busts. A somewhat different robustness check regards the role that initial conditions may play in computing our impulse responses. 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. These initial conditions are technically associated to the transition indicator z_{t-1}, which per each given t is compared with the estimated threshold $\widehat{c}$. In particular, values of $z_{t-1} > \widehat{c}$ ($z_{t-1} \leq \widehat{c}$) indicate that the economy is in a boom (bust). As in all nonlinear analysis of this kind, the risk to incorrectly classify 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 $|z_{t-1} - \widehat{c}| > 1/\delta$, where δ, with $\delta \epsilon \{1, 2\}$. These robustness checks are basically based on the selection of "extreme" realizations of the business cycle (say, deep downturns or solid booms). When $\delta = 2$, about 9% (65%) 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. Given that the relevant effects of uncertainty shocks are found in busts, we focus on realizations of z_{t-1} which are below

[4] Such measure is available at the following website: https://www.cdhowe.org/sites/default/files/attachments/other-research/pdf/Main-Economic-Indicators-used-to-Establish-the-Historical-Chronology-of-Canadian-Business-Cycles1%20%281%29.xls . We use the 12-month growth rate of the "Real GDP (2002 constant prices)" series.

the threshold. Figure A3 shows the outcome of this exercise. Evidently, our responses our robust to different selections of initial conditions.

Longer sample for Canada. As written in the text, the measure of EPU we use for Canada is based on information contained in newspapers only. An equivalent measure - the historical EPU series - is available for the US up to 2014, which justifies the end date of our sample. A highly correlated, updated series for the US is available starting from 1985. Figures A4 and A5 report the outcome of an exercise related to a robustness check conducted with this updated series, available until 2017. Our results are virtually unchanged with respect to those documented in this paper.

FEVD: Robustness checks. Table A1 reports the contribution of US EPU shocks for the Canadian indicators considered in our analysis across a number of different models. The main message is that our qualitative results are robust to a number of controls added to our baseline vector.

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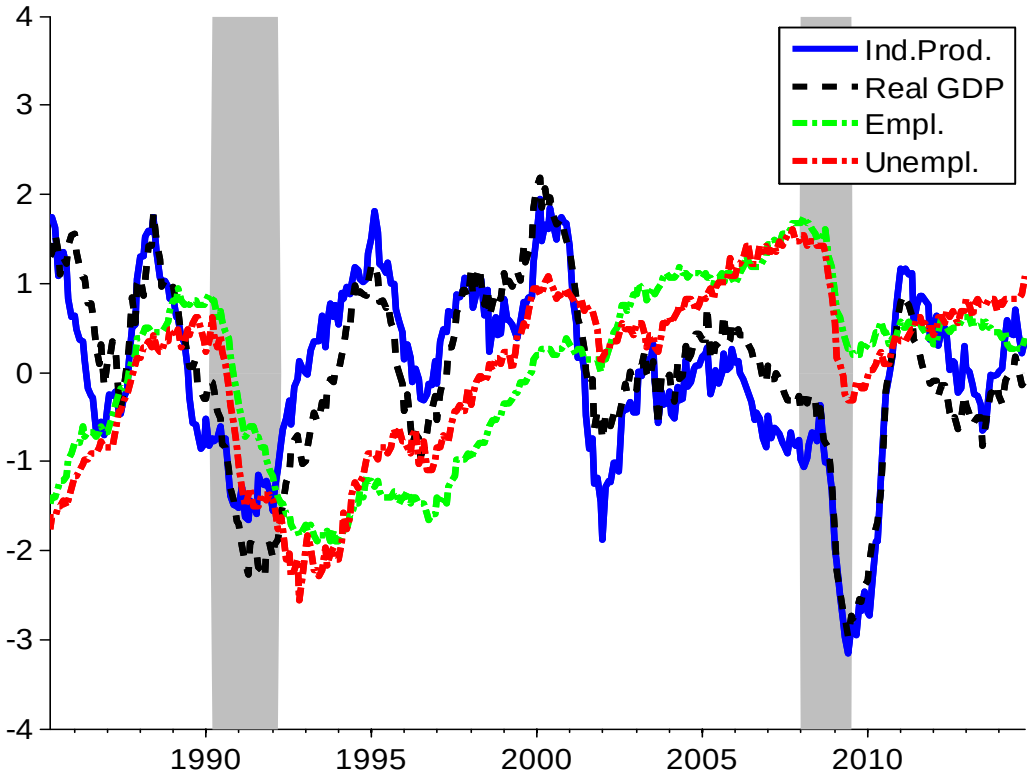

Figure A1: **Canada: Real Activity Indicators.** Sample: 1985:M1-2014:M10. Moving averages of the monthly growth rates of industrial production and real GDP consider eighteen and twelve terms, respectively. The sign of the unemployment rate was switched to highlight the correlation with the other real activity indicators. Grey vertical bars indicate recessions as dated by the Economic Cycle Research Institute. All indicators have been standardized so to have mean zero and unit variance.

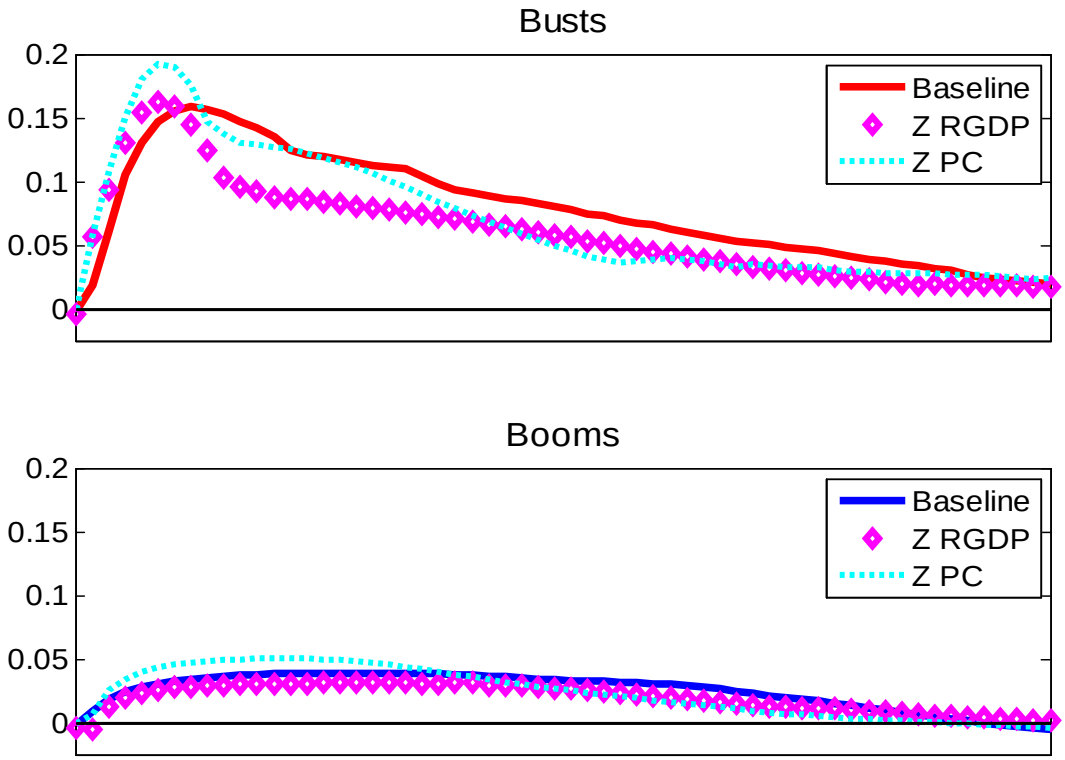

Figure A2: **State-contingent unemployment response: Robustness to alternative transition indicators.** Sample: 1985:M1-2014:M10. Moving averages of the monthly growth rates of real GDP and the principal component constructed by considering eighteen and twelve terms, respectively. Principal component (PC) constructed by considering four different business cycle indicators, i.e., the growth rates of industrial production and real GDP and the rates of unemployment and employment.

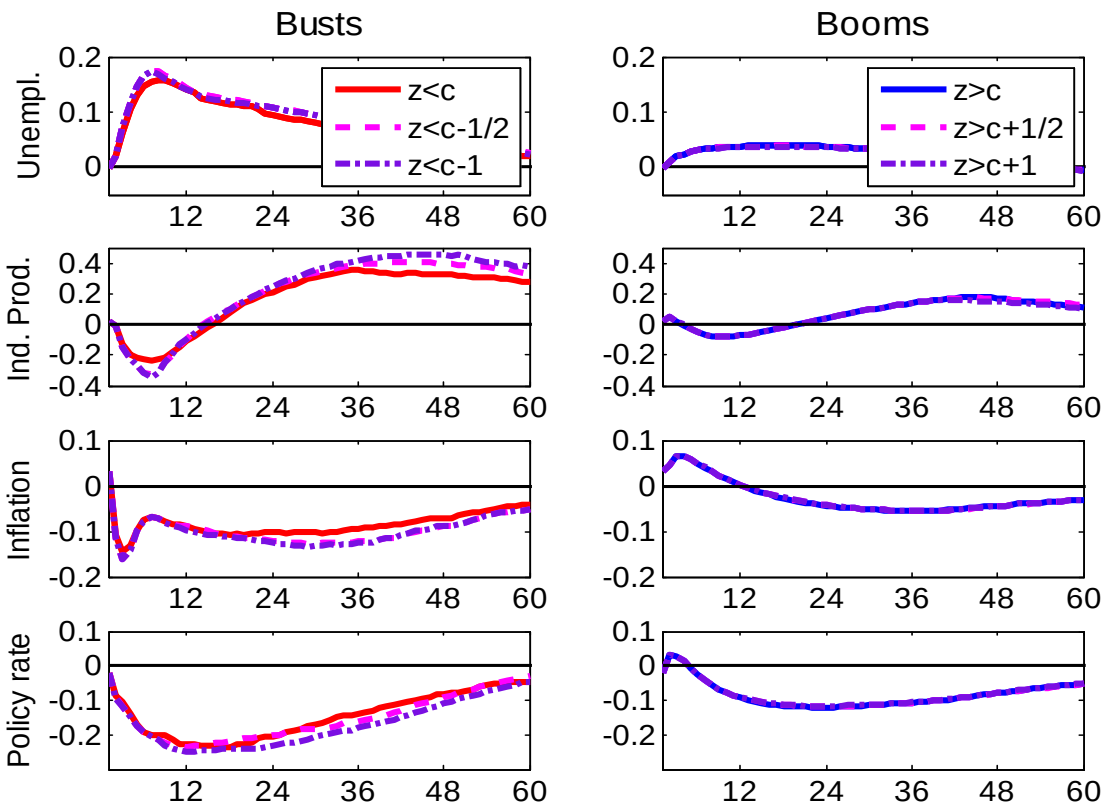

Figure A3: **Effects of a Shock to the US EPU Index on the Canadian economy: Role of Initial Conditions.** Sample: 1985:M1-2014:M10. Generalized median impulse responses to a one-standard deviation shock to the US EPU shock computed by considering three different sets of initial conditions identified with upward/downward deviations of sizes 0, 0.5, 1 with respect to the estimated threshold. Transition indicator for Canada: 18-term moving average of the monthly growth rate of the Canadian industrial production.

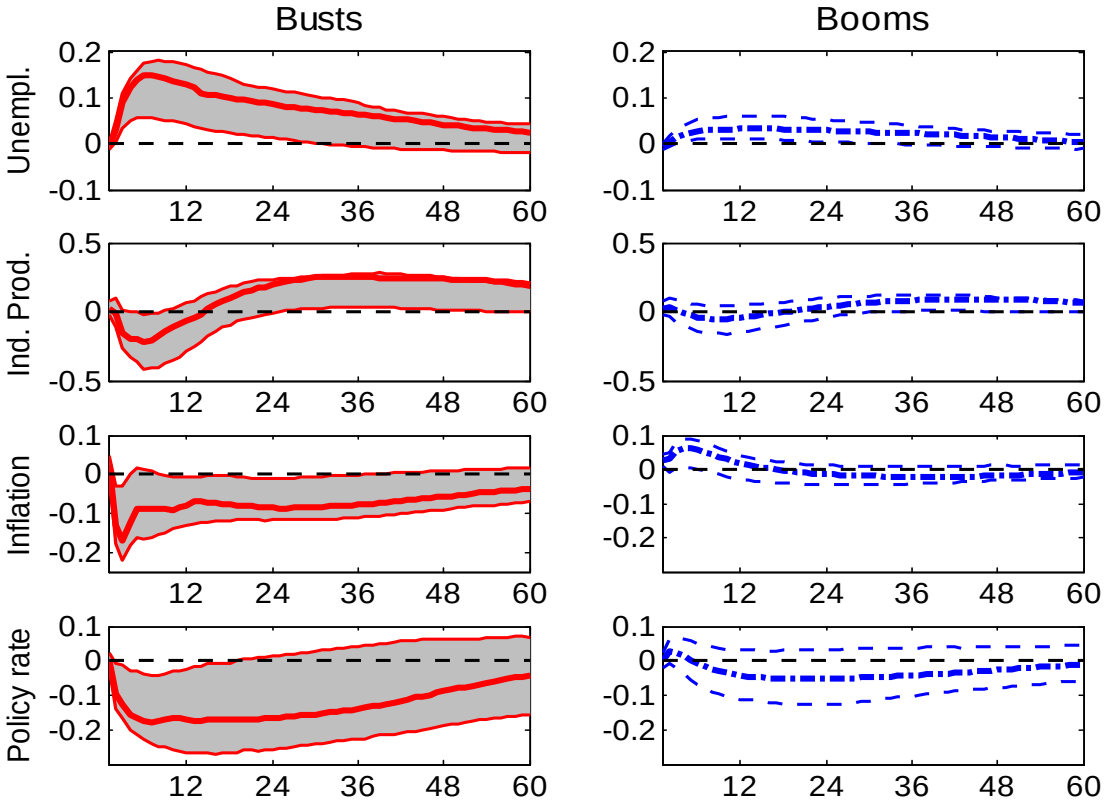

Figure A4: **Effects of a shock to the US EPU Index on the Canadian economy: Longer sample.** Sample: 1985:M1-2017:M1. Median generalized impulse responses to a one-standard deviation shock to the US EPU index hitting the Canadian economy in busts (red solid line) and booms (blue dashed-dotted line). 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.

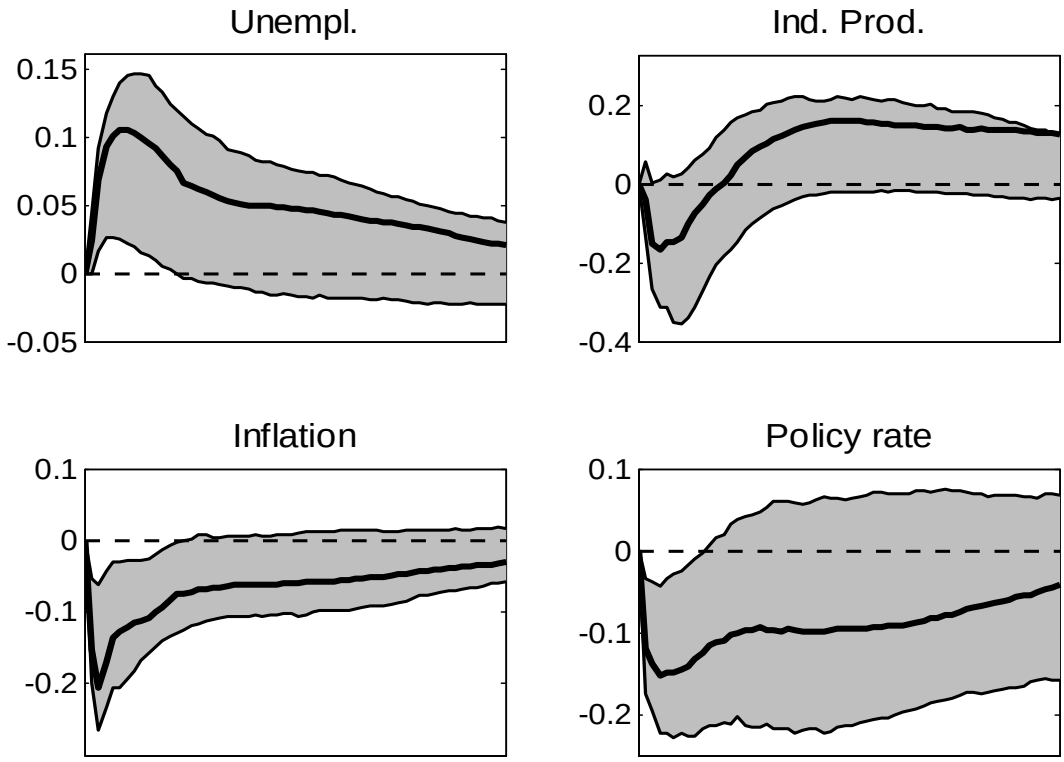

Figure A5: **Effects of a shock to the US EPU Index on the Canadian economy: Difference between states, longer sample.** Sample: 1985:M1-2017:M1. Differences between median generalized impulse responses in busts and booms to a one-standard deviation shock to the US EPU Index. Median realizations identified 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.

				Busts			
Model/Variable	EPU_t^{US}	EPU_t	ΔIP_t	u_t	π_t	R_t	$\Delta\epsilon_t$
Baseline	0.62	**0.32**	0.08	**0.26**	0.14	0.17	0.14
Dummy	0.88	**0.26**	0.08	**0.18**	0.19	0.22	0.12
VXO	0.38	**0.25**	0.07	**0.20**	0.13	0.13	0.14
Factor	0.40	**0.19**	0.03	**0.12**	0.09	0.09	0.07
Unc. last	0.60	**0.32**	0.07	**0.27**	0.14	0.13	0.11
Econ. Unc.	0.35	**0.15**	0.03	**0.13**	0.08	0.05	0.05
EBP	0.41	**0.23**	0.06	**0.14**	0.09	0.14	0.07
Oil Hamilton	0.46	**0.26**	0.06	**0.10**	0.08	0.07	0.05
Oil Kilian	0.41	**0.22**	0.03	**0.11**	0.09	0.08	0.06
Comm. Prices	0.50	**0.27**	0.04	**0.14**	0.10	0.10	0.08

				Booms			
Model/Variable	EPU_t^{US}	EPU_t	ΔIP_t	u_t	π_t	R_t	$\Delta\epsilon_t$
Baseline	0.72	**0.31**	0.02	**0.09**	0.04	0.06	0.04
Dummy	0.93	**0.19**	0.01	**0.02**	0.03	0.07	0.03
VXO	0.64	**0.33**	0.02	**0.10**	0.04	0.09	0.05
Factor	0.60	**0.23**	0.01	**0.04**	0.02	0.03	0.02
Unc. last	0.65	**0.26**	0.02	**0.08**	0.02	0.03	0.02
Econ. Unc.	0.33	**0.08**	0.00	**0.02**	0.02	0.01	0.01
EBP	0.52	**0.21**	0.01	**0.03**	0.02	0.03	0.01
Oil Hamilton	0.69	**0.30**	0.01	**0.06**	0.03	0.03	0.01
Oil Kilian	0.68	**0.28**	0.01	**0.07**	0.02	0.02	0.01
Comm. Prices	0.71	**0.29**	0.01	**0.05**	0.02	0.03	0.01

Table A1: **Forecast Error Variance Decomposition: US EPU Shocks, Different Scenarios.** 2 year-ahead forecast error variance decomposition. The figures reported in the table refer to the point estimates of the contribution of US EPU shocks to the forecast error variance decomposition of the variables included in the baseline STVAR.