Transmission of Government Spending Shocks in the Euro Area: Time Variation and Driving Forces*

Markus Kirchner[†], Jacopo Cimadomo[‡], Sebastian Hauptmeier[§]

February 2010

Abstract

This paper provides new evidence on the effects of government spending shocks and the fiscal transmission mechanism in the euro area for the period 1980-2008. Our contribution is two-fold. First, we investigate changes in the macroeconomic impact of government spending shocks using time-varying structural VAR techniques. The results show that the short-run effectiveness of government spending in stabilizing real GDP and private consumption has increased until the end-1980s but it has decreased thereafter. Moreover, government spending multipliers at longer horizons have declined substantially over the sample period. We also observe a weaker response of real wages and a stronger response of the nominal interest rate to spending shocks. Second, we provide econometric evidence on the driving forces behind the observed evolution of spending multipliers. We find that a higher ratio of credit to households over GDP, a smaller share of investment expenditures and a larger wage component have led to decreasing contemporaneous multipliers. At the same time, our results indicate that higher government debt-to-GDP ratios have negatively affected long-term multipliers.

Keywords: Government spending shocks; Fiscal transmission mechanism; Structural change; Bayesian analysis; Structural vector autoregressions; Time-varying parameter models

JEL classification: C32; E62; H30; H50

---

*We would like to thank Gianni Amisano, Peter Claeys, Wouter den Haan, Carlo Favero, Massimo Giuliodori, Fédéric Holm-Hadulla, George Hondroyiannis, Alexander Kriwoluzky, Marco Lombardi, Andreas Schabert, Petr Sedláček, Christian Stoltenberg, Mathias Trabandt, Ad van Riet, and Sweder van Wijnbergen for helpful comments and discussions. We are particularly grateful to Joan Paredes, Javier Pérez, and Diego Pedregal for providing us their quarterly fiscal database for the euro area. We also thank seminar participants at the University of Amsterdam, the European Central Bank, and the ECB Public Finance Workshop "Challenges for fiscal sustainability in the EU". The views expressed in this paper are those of the authors, and do not necessarily reflect those of the European Central Bank. Any remaining errors are the sole responsibility of the authors.

[†]University of Amsterdam, Department of Economics, Roetersstraat 11, 1018 WB Amsterdam, The Netherlands; e-mail: markus.kirchner@uva.nl

[‡]European Central Bank, Fiscal Policies Division, Kaiserstrasse 29, 60311 Frankfurt am Main, Germany; e-mail: jacopo.cimadomo@ecb.int

[§]European Central Bank, Fiscal Policies Division, Kaiserstrasse 29, 60311 Frankfurt am Main, Germany; e-mail: sebastian.hauptmeier@ecb.int

1 Introduction

Fiscal policy has been rediscovered as a tool for short-run economic stabilization. Governments around the world have enacted unprecedented fiscal stimulus packages to counter the severe economic downturn triggered by the financial crisis. For instance, the fiscal stimulus adopted within the European Economic Recovery Plan (EERP) is expected to reach about 1% of the EU's GDP in 2009 and 0.9% in 2010, and it is largely expenditure based (see European Commission, 2009). At the current conjuncture, however, there is a high degree of uncertainty concerning the effectiveness of government expenditure policies in stabilizing economic activity. The theoretical and empirical literature on the effects of government spending shocks reflects this uncertainty and it is rather inconclusive so far, especially as regards the euro area.[1]

Against this background, this paper offers two original contributions. First, we uncover changes in the effects of government spending shocks in the euro area over the period 1980–2008, using the tools of Bayesian time-varying parameters VAR (TVP-VAR) analysis.[2] Second, we provide econometric evidence on the driving forces of the observed time evolution of government spending multipliers and, in a broader sense, the fiscal transmission mechanism. In particular, using regression analysis we relate spending multipliers at different horizons to a set of macroeconomic indicators and to the composition of spending. The underlying idea is that time variation–caused by structural change–may reveal new facts about the macro impact of government expenditure policies. To the best of our knowledge, this is the first paper which investigates time variation in the effects of government spending shocks through an application of state-of-the-art Bayesian techniques, and which at the same time provides empirical evidence on the driving factors behind the changing patterns of spending multipliers by means of a systematic exploitation of state dependency.

We believe that the TVP-VAR methodology outperforms simpler methods including sub-sample or rolling-windows estimation for several reasons. Most importantly, structural changes might not be easily identified a priori, or they may take the form of processes that last several years. In addition, fiscal multipliers might change in a non-monotonic way. Finally, dating a

[1] The related literature is discussed in detail in Section 2.

[2] These tools have been applied previously to investigate changes in the effects of monetary policy in the U.S. and the relation to the "Great Moderation" (see e.g. Cogley and Sargent, 2001, 2005; Primiceri, 2005; Benati and Mumtaz, 2007; Canova and Gambetti, 2009; Galí and Gambetti, 2009), and the implications of structural change for macroeconomic forecasts (see D'Agostino, Gambetti, and Giannone, 2009).

break and determining the size of rolling windows would have to be arbitrary to some extent. Indeed, one can think of numerous structural changes which might impact on the effectiveness of fiscal policy. A choice of sub-samples for one of them (e.g. monetary policy regime changes) is unlikely to fit another (e.g. trade integration).

We focus on the euro area since sub-sample instability should be an imminent fact given significant structural changes experienced since the 1980s. Examples include the adoption of the Maastricht Treaty in 1992, the run-up to the Economic and Monetary Union (EMU), the introduction of the single currency, and the single monetary policy since 1999. Such events should enhance the scope for time variation and help the identification of the driving forces of the fiscal transmission mechanism.

Based on a newly available quarterly fiscal data set developed by Paredes, Pérez, and Pedregal (2009), fixed parameters VAR estimations over the full 1980-2008 sample suggest that, on average, government spending shocks have an expansionary short-run impact and moderately contractionary long-term effects on output and the components of private demand in the euro area. However, our time-varying approach allows to uncover important changes in the macroeconomic impact of government spending shocks. In particular, our results show that short-run government spending multipliers on real GDP and private consumption have increased until the end-1980s but they have decreased thereafter. Moreover, the expansionary effects of government spending have become more short-lived over time. Long-term multipliers have decreased substantially over the sample period. The effectiveness of spending based fiscal expansions in stimulating economic activity thus appears to be particularly low in the current decade. In addition, we show that smaller spending multipliers coincide with a weaker response of real wages and a stronger response of the short-term nominal interest rate.

With respect to the driving forces of the fiscal transmission mechanism, our evidence points towards households' access to credit as the most important determinant of the size of contemporaneous spending multipliers. In particular, we find that a one percentage point increase in the ratio of credit provided to households over GDP leads to a decline between 0.04 and 0.06 points in contemporaneous multipliers. This result provides support for recent arguments suggesting that access to credit and non-Ricardian behavior by households matter for the size of fiscal multipliers. Regarding the composition of government spending, we find that a smaller share of investment expenditures and a larger wage component have led to declining short-run multipliers. Our results therefore indicate that government investment may have an additional positive aggregate supply effect in addition to the aggregate demand effect of government goods purchases. The fact that wage

payments are associated with lower multipliers provide support for a recent arguments stating that government wage expenditures may have adverse effects in an imperfect labor market through their impact on reservation wages (see Alesina and Ardagna, 2009). Finally, we find that the level of government debt is the main determinant of the long-term effect of government spending. A one percentage point increase in the ratio of government debt over GDP leads to a decline by about 0.01 points in spending multipliers after five years. This result suggests that, given higher initial government financing needs, sustained deficits after a spending shock may lead to rising concerns on the sustainability of public finances and expectations of a larger future consolidation, which depresses private demand.

The remainder of the paper is organized as follows. Section 2 reviews the related literature. Section 3 describes our econometric model, the estimation method, the data and the structural identification approach. Section 4 presents the VAR results. It first discusses results from a time invariant VAR and then the evidence from the TVP-VAR. Section 5 investigates the driving forces of the fiscal transmission mechanism. It first provides an account of existing views on the transmission mechanism and then reveals the determinants underlying the observed time variation in spending multipliers using simple regression analysis. Section 6 concludes.

2 Related Literature

On the theoretical side no consensus has been achieved so far concerning the impact of government spending shocks on the main macroeconomic variables. General equilibrium models currently used to evaluate the effects of government spending tend not to be robust in their predictions (cf. Cogan, Cwik, Taylor, and Wieland, 2009). Neoclassical models with optimizing agents and fully flexible prices typically indicate a rise in output and employment but a fall in private consumption and real wages following an exogenous increase in government goods purchases (see e.g. Baxter and King, 1993). New Keynesian sticky-price models models can generate an increase in real wages, depending on the monetary regime (see Linnemann and Schabert, 2003). However, basic versions of these models also tend to predict a crowding out of private consumption, unless additional features are included which dampen the negative wealth effect of a fiscal expansion. Examples include non-Ricardian consumers (Galí, Lopéz-Salido, and Vallés, 2007), imperfect substitutability between public and private consumption (Linnemann and Schabert, 2004), and small wealth effects on labor supply (Monacelli and Perotti, 2008).

On the empirical side the effects of government spending shocks are typically investigated within the structural VAR framework.[3] Alternatives include the event-study approach by Ramey and Shapiro (1998), or more recently, Ramey (2009). Despite an increasing number of papers in this field, many open questions do remain. In particular, the effects of government spending shocks in the euro area are largely unexplored. Indeed, even though fiscal policy in the euro area is still mostly a country-specific matter, the aggregate impact of fiscal policy is of high practical relevance for policy makers.[4] Initiatives such as the EERP also indicate an interest in co-ordinated fiscal policy in Europe, although the impact of such actions remains uncertain.

The scarcity of empirical results for the euro area as a whole and also for euro area countries has been due to the lack of quarterly fiscal data, a limitation which has been overcome recently through a newly available quarterly fiscal database for the euro area compiled by Paredes, Pérez, and Pedregal (2009). This data set, which covers the period 1980Q1–2008Q4, is coherent with official annual and quarterly national accounts data, as far as quarterly fiscal data is available from national accounts (mostly for the period 1999Q1 onwards). Based on this data set, Burriel, de Castro, Garrote, Gordo, Paredes, and Pérez (2009) show that the qualitative responses of macroeconomic variables to fiscal shocks in a (weighted) representative euro area country compare well with standard results for the U.S. and previous results for some EU countries.

There is also disagreement on whether fiscal policy, and in particular government spending increases, have lost power in stimulating economic activity over time, and if so to what extent and why. In particular, the literature lacks empirical tests of potential explanations for the changing effects of government spending shocks. Blanchard and Perotti (2002) find that the size of spending multipliers on output in the U.S. varies considerably across subperiods. However, this paper does not provide a clear-cut explanation, based on econometric results, for the observed changes. Similar accounts of instability, based on sub-sample or rolling-windows estimation, can be found in Perotti (2005), Bénassy-Quéré and Cimadomo (2006), Bilbiie, Meier, and Mueller (2006), and Caldara and Kamps (2008). These studies generally

[3]See e.g. Fatás and Mihov (2001), Blanchard and Perotti (2002), Perotti (2005), Bénassy-Quéré and Cimadomo (2006), Caldara and Kamps (2008), and Mountford and Uhlig (2009).

[4]In fact, several DSGE models and other quantitative models of the euro area do already explicitly account for aggregate fiscal variables and aggregate fiscal data. See, for example, Smets and Wouters (2003), Fagan, Henry, and Mestre (2005), Christoffel, Coenen, and Warne (2008), Ratto, Roeger, and in 't Veld (2009), and Forni, Monteforte, and Sessa (2009).

conclude that the responses of the U.S. and of some European economies to fiscal policy shocks have become weaker in the post-1980 period. In a DSGE framework, Bilbiie, Meier, and Mueller (2006) show that the more active monetary policy in the Volcker-Greenspan period and increased asset market participation can account for the observed decline in spending multipliers in the U.S. after 1980. Finally, Perotti (2005) suggests that relaxation of credit constraints, increasing financial market sophistication, a stronger real interest rate response, and changes in monetary policy could explain the decline in the effects of government spending on GDP and its components. Again, however, the above papers provide relatively little econometric evidence in order to support potential explanations for changes in the effects of government spending shocks.

3 Econometric Methodology

Our empirical approach uses the techniques of Bayesian inference. We prefer a Bayesian approach over estimation by classical statistical methods for a number of reasons. Most importantly, this approach facilitates the estimation of time variation in multivariate linear structures and stochastic volatility models. As discussed by Primiceri (2005), Bayesian methods are the natural choice for the estimation of unobserved component models of this type where the distinction between parameters and shocks is less clear than in other models. The main advantage of Bayesian techniques in this context is however related to the high dimensionality of such an estimation problem. Although it would in principle be possible to write up the likelihood for the problem, it is a hard task to maximize it over a large number of parameters. By using prior information and by splitting up the original problem into a few smaller steps, Bayesian methods deal efficiently with the high dimension of the parameter space.[5]

3.1 Reduced-form VAR

We consider two alternative specifications of a reduced-form VAR of lag order p. The first version has fixed parameters:

$$y_t = b_0 + B_1 y_{t-1} + \cdots + B_p y_{t-p} + u_t, \quad t = 1, \ldots, T \tag{1}$$

where the vector y_t includes government spending, output, private consumption, the short-term interest rate, and possibly other macroeconomic indica-

[5]In addition, the Bayesian approach allows for a conceptually clean way of calculating statistics of interests such as error bands for impulse responses (see Sims and Zha, 1999).

tors. The vector b_0 collects the VAR intercepts, and the B_i, $i = 1, \ldots, p$, are matrices of coefficients. The vector of innovations u_t is Gaussian white noise with mean zero and covariance R.

In the second version of the VAR, we generalize specification (1) and allow for time-varying coefficients and heteroskedastic innovations.[6] The first aspect allows for changes in the propagation of shocks through the economy. The second aspect is introduced in order to allow for changes in the distribution of the underlying stochastic shocks. Both features are supposed to capture structural changes such as shifts in private sector behavior and/or changes in the conduct of policy. Hence:

$$y_t = b_{0,t} + B_{1,t} y_{t-1} + \cdots + B_{p,t} y_{t-p} + u_t, \quad t = 1, \ldots, T \tag{2}$$

where $u_t \sim \mathrm{NID}(0, R_t)$. Stack the VAR intercepts and coefficients by equations in a vector $\beta_t = \mathrm{vec}(Z_t')$, where $Z_t = [b_{0,t}, B_{1,t}, \ldots, B_{p,t}]$ and $\mathrm{vec}(\cdot)$ is the column stacking operator. This state vector of coefficients is assumed to follow a driftless random walk:[7]

$$\beta_t = \beta_{t-1} + \varepsilon_t \tag{3}$$

where $\varepsilon_t \sim \mathrm{NID}(0, Q)$. The innovation covariance is R_t, which can be decomposed using a triangular factorization of the form

$$R_t = A_t^{-1} H_t (A_t^{-1})' \tag{4}$$

where A_t^{-1} is lower triangular with ones on the main diagonal and H_t is diagonal. Stack the elements below the main diagonal of A_t row-wise in a vector α_t. Collect the diagonal elements of H_t in a vector h_t. Like the coefficient states, the covariance and volatility states are modeled as (geometric) random walks:[8]

$$\begin{aligned} \alpha_t &= \alpha_{t-1} + \nu_t \\ \log h_t &= \log h_{t-1} + \omega_t \end{aligned} \tag{5}$$

[6] Our specification of the TVP-VAR follows Cogley and Sargent (2001, 2005) and Primiceri (2005). We apply some additional restrictions on the hyperparameters which are discussed below.

[7] Compared to alternative specifications such as regime switching models, the random walk specification has the advantage that it allows for smooth shifts as opposed to discrete breaks in the states of the model. As discussed by Primiceri (2005), regime switching models may well capture some of the rapid shifts in policy but they seem less suitable for describing changes in private sector behavior where aggregation usually smoothes most of the changes, or learning dynamics of both private agents and policy makers.

[8] Modeling volatilities and covariances separately instead of directly modeling the elements of the variance-covariance matrix ensures that R_t is always positive definite.

where $\nu_t \sim \text{NID}(0, S)$ and $\omega_t \sim \text{NID}(0, W)$. Following Primiceri (2005) both the diagonal elements and the off-diagonal elements of the reduced-form covariance matrix can drift over time, thus allowing for changes in the contemporaneous relations among the endogenous variables.

The joint distribution of shocks is postulated as $[u_t, \varepsilon_t, \nu_t, \omega_t]' \sim \text{NID}(0, V_t)$, where V_t is block diagonal with blocks R_t, Q, S and W. Notice that an unrestricted covariance matrix would drastically increase the number of parameters and complicate the estimation problem. The assumption of independence of R_t and the hyperparameters implies that innovations to the VAR parameters are uncorrelated with the VAR innovations. This assumption seems plausible. The VAR innovations capture business cycle events, policy shocks, or measurement errors. Such short-term movements should be unrelated to long-term institutional changes and other changes in the structure of the economy, which are captured by changes in the VAR parameters. For example, the introduction of the single currency in the euro area should not have been related to technology shocks, government spending shocks, etc.

We make the additional assumption that Q, S and W are diagonal in order to reduce the dimensionality of the problem further and to simplify inference. The assumption of (block) diagonality of S ensures that the rows of A_t evolve independently such that the covariance states can be estimated row by row (cf. Primiceri, 2005). Diagonality of W implies that the volatility states are independent such that the simple univariate algorithm of Jacquier, Polson, and Rossi (1994) can be applied to each element of u_t in order to estimate the volatility states. The reduction of estimated parameters resulting from the diagonality restrictions on Q and S helps to save degrees of freedom in our relatively short euro area data set. We furthermore show in Appendix D that restricting all hyperparameter matrices to be diagonal tends to improve the performance of the estimation algorithm.

3.2 Estimation method

Both versions of the reduced-form VAR are estimated by Bayesian methods. For the version with fixed parameters, our prior and posterior for the coefficient matrices b_0 and B_i, $i = 1, \ldots, p$, and the covariance matrix R belong to the Normal-Wishart family with a diffuse prior centered on OLS estimates over the full sample.[9] For the TVP-VAR, we apply a variant of the Gibbs sampler (see Geman and Geman, 1984; Smith and Roberts, 1993).[10] We

[9]For details on the estimation using the Normal-Wishart distribution, see Uhlig (2005).

[10]See Cogley and Sargent (2001, 2005), Primiceri (2005), Benati and Mumtaz (2007), Galí and Gambetti (2009), and D'Agostino, Gambetti, and Giannone (2009) for applications of Gibbs sampling algorithms to TVP-VARs.

briefly outline the main steps and refer to Appendix A for details on the estimation algorithm. The Gibbs sampler iterates on four steps, sampling in each step from lower dimensional conditional posteriors as opposed to the joint posterior of the whole parameter set.

(a) VAR coefficients. Conditional on the data and a history of covariance and volatility states, the observation equation (2) is linear with Gaussian innovations and a known covariance matrix. The VAR coefficients can thus be sampled using the Kalman filter and a backward recursion, as described in Carter and Kohn (1994) and Cogley and Sargent (2001).

(b) Elements of A_t. Conditional on the data and a history of coefficient and volatility states, equation (2) can be rewritten as $A_t u_t = v_t$, with $\text{cov}(v_t) = H_t$. This is a linear Gaussian state space system with independent equations, due to the (block) diagonal structure of S (see Primiceri, 2005). The algorithm of Carter and Kohn (1994) can thus be applied equation by equation to sample the elements of A_t on each row below the main diagonal.

(c) Elements of H_t. Conditional on the data and a history of coefficient and covariance states, the orthogonalized innovations v_t are observable. Given the diagonal structure of W, we sample the diagonal elements of H_t using the univariate algorithm of Jacquier, Polson, and Rossi (1994) element by element, following Cogley and Sargent (2005).

(d) Hyperparameters. Conditional on the data and the parameter states, the state innovations ε_t, ν_t and ω_t are observable. This allows to draw the hyperparameters (i.e. the elements of Q, S and W) from their respective distributions.

Under relatively weak regularity conditions (see Roberts and Smith, 1994) and after a sufficiently long burn-in period, iterations on these steps produce a realization from the joint posterior distribution. We generate 60,000 draws from the Gibbs sampler, of which we burn the first 50,000 to let the Markov chain converge to its ergodic distribution. Of the remaining 10,000 draws, we keep every 10th draw in order to break the autocorrelation of draws.[11] This leaves us with 1,000 draws from the joint posterior distribution of the model parameters. Appendix C investigates the convergence properties of the Markov chain, concluding that these properties are overall satisfactory.

We follow conventional choices in the TVP-VAR literature in the calibration of the priors. The choices made are similar as in Primiceri (2005) but we have a somewhat more conservative stance on the degrees of freedom of the prior distributions which we set to the minimum value allowed for the priors

[11] The Gibbs sampler is a dependence chain algorithm. However, independent draws should be used when calculating statistics of interest such as posterior means and impulse responses.

to be proper. See Appendix B for details on the calibration of the priors.

Unlike most previous TVP-VAR studies, we do not impose a prior restriction on the VAR coefficients saying that draws which do not satisfy stationarity conditions are discarded. Cogley and Sargent (2001) have proposed such a restriction for U.S. monetary policy, the argument being that the Fed conducted monetary policy in a purposeful way thus ruling out unstable paths of inflation (see Cogley and Sargent, 2005). Such a point is harder to defend for aggregate euro area fiscal data since there may have been fiscal instability in some countries in the past. The potential downside of not imposing the stationarity conditions is that this may exaggerate the amount of time variation in the data due to a potentially large amount of unstable draws. We therefore check the robustness of our results to the imposition of the stationarity conditions in Section 4.3.

3.3 Data description

Our baseline VAR includes data on real government spending, real GDP, real private consumption, and the short-term nominal interest rate for the euro area covering the period 1980Q1–2008Q4. Real GDP is our measure of economic activity, and private consumption is included since it is the largest component of aggregate demand and a key determinant of the size of government spending multipliers (see Galí et al., 2007). The short-term interest rate is added to this small-scale VAR in order to assess the effects of monetary-fiscal policy interactions and potential changes thereof.[12] We also investigate the impact of government spending shocks on a broader set of macroeconomic indicators, i.e. real private investment, the real wage, real net taxes and the Harmonized Index of Consumer Prices (HICP).[13] These variables are added all together in the fixed coefficients VAR. In the specification of the TVP-VAR we are constrained by the need to avoid overparameterization and exhausting available degrees of freedom. Therefore, the additional variables are added one at a time to the baseline specification which limits the number of variables in the VAR to a maximum of five indicators.

As Burriel et al. (2009), we use a newly available quarterly fiscal data set compiled by Paredes et al. (2009) in order to construct a measure of government spending. Paredes et al. (2009) employ intra-annual fiscal data,

[12] As argued by Perotti (2005), the long-term interest rate may have a closer relation to private consumption and investment decisions than the short rate and it may therefore be more appropriate to include the long rate in the VAR. Doing so did not lead to any significant changes in our results.

[13] We use the HICP instead of the GDP deflator to assess the response of prices to spending shocks, due to its closer link to monetary policy decisions in the euro area.

mostly on a cash basis, in a mixed-frequencies state space model to obtain quarterly fiscal data for the above-mentioned period. By construction these data are coherent with annual and quarterly national accounts data, as far as quarterly fiscal data is available from national accounts. The main advantage of this new data set is that it avoids the endogenous bias that arises if fiscal data interpolated on the basis of general macroeconomic indicators were used with macroeconomic variables to assess the impact of fiscal policies. Other macroeconomic data for the euro area are taken from the ECB's Area-Wide Model database (Fagan et al., 2005) and the Bank of International Settlements macroeconomic series.

In order to enhance comparability with the previous literature, our data definitions closely follow related fiscal VAR studies (see e.g. Blanchard and Perotti, 2002; Perotti, 2005; Caldara and Kamps, 2008; Mountford and Uhlig, 2009; Burriel et al., 2009). In particular, government spending is defined as the sum of general government final consumption expenditure and gross investment. Net taxes are defined as non-interest general government revenue net of transfers. Transfers are thus considered as negative taxes.[14] Private investment has been computed by deducting government investment from the Paredes et al. (2009) data set from total economy investment from the Area-Wide Model database. All of the above variables are expressed in per capita terms.[15] The real wage is measured in hourly terms. The data are seasonally adjusted and the GDP deflator is used to obtain real variables. Both the fixed coefficients VAR and the TVP-VAR are estimated in levels and prior to the estimation all variables except the interest rate were transformed into natural logarithms.[16] Figure 1 shows the data used in the baseline VAR specification, expressing government spending and private consumption as shares of GDP.

3.4 Structural interpretation

The reduced-form VAR attempts to capture a structural representation with uncorrelated shocks. The reduced-form innovations are therefore linear trans-

[14]Following Burriel et al. (2009) transfers include all expenditure items except government consumption, government investment and interest payments. The general government primary balance is therefore obtained as the difference between net taxes and spending as defined above.

[15]The labor force is used as a proxy for total population, since quarterly data on total population is not available for the entire sample period.

[16]The fixed coefficients VAR includes a quadratic time trend in order to account for the presence of trends in real variables and the nominal interest rate. Note that a deterministic trend becomes redundant in the TVP-VAR.

formations of some underlying structural shocks e_t with $\mathrm{E}[e_t e_t'] = I$, i.e.

$$u_t = Ce_t, \quad t = 1, \ldots, T$$

for the time invariant VAR and

$$u_t = C_t e_t, \quad t = 1, \ldots, T$$

for the TVP-VAR. In particular, the residuals in the equation for government spending can be considered as linear combinations of three types of shocks (see Blanchard and Perotti, 2002): (1) The automatic response of spending to movements in the business cycle, prices, and interest rates. (2) The systematic discretionary response of spending to macroeconomic developments. (3) Random discretionary innovations to spending, which are the truly structural government spending shocks of interest. Without restrictions on the matrices C or C_t and therefore the reduced-form covariance matrix, the above system is not identified since many combinations of structural shocks can generate the same reduced-form innovations.

We identify government spending shocks by assuming that government spending is predetermined in a system with output, consumption, the interest rate, and possibly other macroeconomic variables. We thereby follow Fatás and Mihov (2001) who estimate a recursive VAR where government spending is ordered first and where the innovation in the first equation of the VAR is interpreted as a structural government spending shock.[17] The desired linear combination is then achieved by a Cholesky decomposition, i.e. $R = CC'$ and $R_t = C_t C_t'$ where C and C_t are lower triangular matrices. All variables in the VAR are therefore allowed to respond contemporaneously to government spending shocks but government spending does not react within a quarter to shocks to other variables in the system. The fact that our definition of spending does not include interest payments justifies ordering spending before the interest rate. The fact that government spending is defined net of transfer payments justifies the assumption of acyclicality, i.e. there is no automatic reaction of spending to movements in the business cycle. Similarly, due to implementation lags in policy-making a discretionary fiscal response to a change in the economy is unlikely to occur. Both assumptions are also made in the identification scheme due to Blanchard and Perotti (2002) and Perotti (2005), who use institutional information about the tax and transfer system in the identification of structural shocks. When more variables are included

[17] See Caldara and Kamps (2008) for a comparative study of alternative identification methods including the recursive approach of Fatás and Mihov (2001), the identification scheme due to Blanchard and Perotti (2002), the sign restrictions approach of Mountford and Uhlig (2009), and the event-study approach of Ramey and Shapiro (1998).

in the VAR, the assumption that government spending does not react within a quarter to shocks to those variables can be justified on similar grounds, i.e. spending is not affected contemporaneously by shocks originating in the private sector.

Impulse responses of the endogenous variables included in the VAR to a one-time structural shock to government spending are then computed as follows. In the time invariant case, given a posterior (empirical) distribution of $R = CC'$ the matrix C gives the contemporaneous responses (at horizon $k = 0$) of the endogenous variables to unitary shocks e_0. Given a distribution of VAR coefficients, model (1) with $u_k = 0$ can then be used to calculate impulse responses at horizons $k = 1, 2, 3, \ldots$. In the time-varying parameters case, we apply a local approximation to the impulse responses at time t, following e.g. Galí and Gambetti (2009). That is, the matrices C_t are computed from the posterior distribution of reduced-form covariance matrices $R_t = C_t C_t'$, which give the contemporaneous impulse responses to unitary shocks e_t at time t. The posterior distribution of VAR coefficients at time t is then applied to calculate the implied responses at horizon $t + k$, for $k = 1, 2, 3, \ldots$, using model (2) with $u_{t+k} = 0$. This leaves us with a posterior distribution of impulse responses where the responses of all variables to spending shocks hitting the economy at time t are allowed to vary over time.

A recent criticism of the structural VAR approach for identifying fiscal policy shocks centers on the fact that this approach often yields qualitatively different results for the U.S. (in terms of the direction of the responses of private consumption and real wages to spending shocks) than the event-study approach of Ramey and Shapiro (1998) which builds on military episodes in order to identify exogenous spending shocks. Ramey (2009) points out that these differences can be traced back to differences in the timing with which news about spending increases arrives if such spending increases are anticipated in advance of their implementation. The challenge posed by fiscal anticipation effects to structural VAR methods is that they may not only mismeasure the timing of shocks but their moving average representation may have non-fundamental roots such that structural fiscal shocks cannot be recovered from past fiscal data (see Leeper, Walker, and Yang, 2009a).

However, the main alternative to structural VAR identification, the event-study approach of Ramey and Shapiro (1998), cannot be applied in a straightforward manner in the context of our study.[18] Comparably large and easily identified (military or other) spending increases as in the U.S. have been

[18]Other alternatives to structural VAR methods include an approach based on flipping non-fundamental roots using Blaschke matrices suggested by Mertens and Ravn (2009) and a DSGE model based approach suggested by Kriwoluzky (2009) who estimates a vector moving average model in order to circumvent the issue of non-invertibility.

absent in aggregate euro area data over the observed sample. A major exception are the fiscal stimulus packages announced and adopted within the EERP in 2009–2010, but this period is not part of our sample. Even if this period was included in the sample, the results from an application of the event-study approach would likely be driven by very few isolated episodes such as this one or the German reunification. Furthermore, whether fiscal shocks are truly unanticipated or not matters only if anticipated and unanticipated fiscal shocks have different effects. This is a controversial empirical issue, largely revolving around the importance of financial constraints and other frictions. Perotti (2005) cites empirical evidence showing that private consumption displays large contemporaneous responses to income tax refunds and changes in social security taxes, although both are predictable. Finally, anticipation effects are unlikely to undermine the main results of this paper on the time variation in spending multipliers. While anticipation effects might bias the estimated impulse responses, it is not clear whether and why such effects have changed over time.

4 The Effects of Spending Shocks

We organize the discussion of results in this section as follows. Section 4.1 presents the results for the fixed parameters structural Bayesian VAR (BVAR), in order to give an impression of the impact of government spending shocks over the full sample. Section 4.2 presents the evidence from the identified TVP-VAR on time variation in the effects of government spending shocks in the euro area. Section 4.3 investigates the robustness of the TVP-VAR results to imposing a stationarity condition on the VAR coefficients.

4.1 Time invariant impulse responses

Figure 2 reports the estimated impulse responses due to the identified government spending shocks to the four endogenous variables y_t of equation (1) in the baseline specification, together with their 16% and 84% probability bands. Following Blanchard and Perotti (2002), we report the responses of output, consumption and spending (and later on investment and net taxes) to the spending shock in terms of (non-accumulated) multipliers. That is, the original impulse responses of the responding variables are divided by the impact response of government spending and the result is divided by the ratio of government spending and the responding variable. The rescaled impulse responses can thus be interpreted to give the reaction of the responding variable, in percent of real GDP, to a spending shock leading to an initial

increase in the level of government spending of size 1% of real GDP. For the time invariant BVAR the ratio is evaluated at the sample mean. For the TVP-VAR below we take the ratio in the respective quarter.[19]

The government spending shock is estimated to induce a positive response of government spending for about 20 quarters after the shock. The initial reaction of output is positive, the estimated response being about 0.54% due to an increase in government spending of size 1% of GDP. The output response remains positive with 64% probability for 5 quarters after the shock, and the point estimate turns negative after 8 quarters in order to drop to -0.34% in the medium run (13 quarters after the shock) before returning to the baseline. The spending shock also leads to a short-run crowding-in of private consumption. The point estimate of the impact multiplier is 0.24, and the response of consumption is estimated to be positive with 64% probability during 5 quarters after the shock. Similarly as for output, however, consumption is being crowded out in the medium run and the response drops to -0.22% of GDP after 15 quarters before slowly returning to its initial level. The nominal interest rate hardly responds to the spending shock in the initial period, but it then starts to rise and peaks at 0.23 percentage points 5 quarters after the shock and then slowly declines again. The response is estimated to be positive with 64% probability during around 3 years.

In a next step we extend the baseline specification by a broader set of macroeconomic indicators which typically appear in fiscal VAR studies. The impulse responses from an estimated BVAR in government spending, output, consumption, investment, the real wage, net taxes, the HICP, and the nominal interest rate are reported in Figure 3. As a consequence of a government spending shock leading to a rise in the level of government spending of size 1% of GDP, net taxes increase by about 0.8% of GDP on impact indicating an overall fiscal expansion since the aggregate primary deficit increases. Net taxes also return more quickly to baseline than the level of spending such that the shock remains expansionary over the full horizon of the impulse response. Output again tends to rise in the short to medium run before declining below its initial level, and similarly for private consumption and investment. The responses of output and the components of private demand

[19]The following example should clarify the concept. Suppose the spending shock leads to a 2% increase in government spending. Since the share of spending over GDP is roughly 25%, this corresponds to a spending increase of about 0.5% of GDP. Say output increases by 1% and consumption increases by 0.5%, i.e. by 0.25% of GDP since the share of consumption over GDP is approximately 50%. The share of spending over consumption is thus roughly 50%. The corresponding multipliers (increases in % of GDP due to a 1% of GDP increase in spending) would be calculated as $(1/2)/0.25 = 2$ for output and $(0.5/2)/0.5 = 0.5$ for consumption.

are however estimated with relatively little precision, compared to the baseline VAR. The point estimates of the impact multipliers are 0.55 (output), 0.23 (consumption) and 0.06 (investment). The real wage increases by approximately 0.1% on impact and remains above its initial level during more than 12 quarters after the shock. Consumer prices show a muted response in the initial period but they start to increase shortly after the spending shock, indicating an inflationary impact of the fiscal expansion. Monetary policy reacts by increasing the nominal interest rate, whose response resembles that in the baseline specification.

Overall, these results indicate that government spending shocks have had on average expansionary effects on output and the components of private demand as well as real wages in the euro area over the period 1980–2008. It is important to note, however, that the expansionary effects of the spending shock are relatively short-lasting since in the medium to long run output declines as the components of private demand are being crowded out. The increase in the nominal interest rate is consistent with an offsetting reaction of monetary policy to the fiscal expansion in order to reduce inflationary pressure. In general these results compare well with the results of previous structural VAR studies on the euro area. In particular, they are broadly similar to those of Burriel et al. (2008), the main previous fiscal VAR study for the euro area as a whole using a similar data set. Burriel et al. (2008) also find a positive impact of government spending shocks on GDP and private consumption in the short to medium run and a decline in the medium to long run, an increase in the aggregate primary deficit, and a relatively persistent increase in interest rates.

4.2 Uncovering time variation

The time-varying nature of model (2) allows to examine impulse responses for each quarter available in the sample. We start by looking at responses in three selected quarters at the beginning, towards the middle and at the end of the sample, i.e. 1980Q4, 1995Q4 and 2008Q4. State-dependent impulse responses of output, consumption and the nominal interest rate to government spending shocks in these quarters are reported in Figure 4. As in the time invariant case, shocks are normalized to lead to an initial increase in the level of spending of size 1% of GDP at each point of time.

The contemporaneous responses of output and consumption to a government spending shock are larger at the beginning of the sample than at the end of the sample. The point estimates of the impact multipliers are 0.72 (output) and 0.37 (consumption) in 1980Q4 compared to 0.42 (output) and 0.28 (consumption) in 2008Q4. Moreover, the responses of ouput and con-

sumption have lost persistence over time. The effect of a spending shock on output was positive during 6 to 7 quarters in 1980Q4, but only during 4 to 5 quarters in 1995Q4 and 3 to 4 quarters in 2008Q4. The time-varying techniques also uncover increasingly damaging long-run consequences of the fiscal expansion. The response of GDP at a horizon of five years was -0.69% in 1980Q4, but it has declined to -1.62% in 2008Q4. A less expansionary effect on output goes along with a much larger (by a factor of 2 to 3) medium to long term crowding out of consumption. Furtermore, while the estimated impact multipliers tend to be positive with 64% probability at the beginning of the sample, the probability bands are wider and they include the zero line at the end of the sample. The decline in the long-term multipliers is significant as most of the probability mass has shifted downwards. We also note a change in the response of the nominal interest rate. The initial reaction of the interest rate to a spending shock was negative in 1980Q4, close to zero in 1995Q4, and positive in 2008Q4.

The conclusions from Figure 4 are confirmed in Figure 5, which shows state-dependent impulse responses over the whole sample. Only the fourth-quarter response in each year is reported such that the first impulse response reported refers to 1980Q4 while the last one refers to 2008Q4. One can again observe that the effect of government spending on output and consumption has become significantly weaker and less persistent over time, and that the nominal interest rate tends to respond more strongly to spending shocks. Yet the figure also reveals that the expansionary short-run effect on output and consumption peaks towards the end of the 1980s before declining until the most recent decade. Long-term multipliers have steadily declined over the observed sample. It is also obvious that the shape of the response of government spending to the spending shock has remained rather stable over time. The persistence of spending does not show any important time variation.

In Figure 6 we plot the impulse responses of all variables over time at selected horizons, i.e. the contemporaneous responses, the responses after one year and the responses after five years. Again, we observe that the response of government spending has remained relatively stable over time. The impact multiplier on output was slightly below one in the period 1980–1985, it increased above one in the period 1985–1990, and it then decreased to values below 0.5 until 2008. At a horizon of five years the multipliers on output and consumption tend to have declined substantially from values between -0.7 and -1 in the 1980s to values between -1.4 and -1.7 in the recent decade. In general the output multiplier follows the movements of the multiplier on private consumption. The initial reaction of the interest rate was negative until around 1999–2002, and it turned positive afterwards. The medium to long run response of the interest rate has also increased over

time. A stronger response of the nominal interest rate–consistent with a less accommodative stance of monetary policy towards the fiscal expansion–thus seems to have contributed to the observed decline in spending multipliers.

We also investigate the time-varying effects of government spending shocks on a broader set of macroeconomic indicators, adding one at a time private investment, net taxes, the real wage and the HICP to the estimated VAR. Figures 7 and 8 show the estimated state-dependent impulse responses. We observe a small positive short-term effect of spending shocks on private investment, and a medium to long-term crowding out. Similarly as the multipliers on output and consumption, the multiplier on private investment was larger in the first part of the sample. Yet the decline in the estimated multiplier has started to take place somewhat later after the year 2000. We note that the reaction of net taxes to government spending shocks has remained comparably stable over time, and the response is in general smaller than 1% of GDP indicating that the primary deficit has always increased due to the spending shock. A smaller overall fiscal expansion can thus not hold as an explanation for smaller spending multipliers. The response of the real wage shows more time variation. It was positive for several quarters after the shock throughout the sample, but we observe a larger initial reaction and a more persistent response in the first part of the sample. In general the real wage response is similar to the consumption response and it was strongest towards the end of the 1980s.

The response of prices has remained relatively stable over time. Since the nominal interest rate reacts more strongly to government spending shocks, this implies that the real interest rate has tended to increase more due to spending shocks. Agents save more and consume and invest less which means that private demand decreases. Firms respond by decreasing output. If prices are sticky, real wages tend to increase after an expansionary spending shock but given a weaker response of private demand they seem to have done so less in more recent times. Disposable income would therefore respond less strongly to the fiscal expansion and if liquidity constraints play a role, consumers would tend to consume less which reinforces the negative effect on private demand. This and other possible determinants of the observed time variation in spending multipliers will be addressed further in Section 5.

4.3 Robustness: imposing stationarity conditions

Cogley and Sargent (2001) have proposed to impose a prior restriction on the VAR coefficients saying that draws from the Gibbs sampler which do not satisfy stationarity conditions are discarded. We have argued above that such a restriction is difficult to defend for aggregate euro area fiscal data

since there may have been instability in the effects of fiscal policy in some countries in the past. The potential downside of not imposing the stationarity conditions is that this may exaggerate the amount of time variation in the data due to a potentially large amount of unstable draws. We therefore check the robustness of our main results to imposing the stationarity conditions.

Formally, the random walk process (3) for the VAR coefficients β_t, $t = 1, \ldots, T$, characterizes the conditional density $f(\beta_t|\beta_{t-1}, Q)$. Following Cogley and Sargent (2001), we introduce an indicator function $I(\beta_t)$ which rejects unstable draws not satisfying standard eigenvalue stability conditions and which thus enforces stationarity of the estimated TVP-VAR at each point of time. The VAR coefficients are thus postulated to evolve according to

$$p(\beta_t|\beta_{t-1}, Q) = I(\beta_t) f(\beta_t|\beta_{t-1}, Q)$$

Figure 9 shows state-dependent impulse response in the baseline VAR following a positive government spending shock of size 1% of GDP, with the stationarity conditions imposed. A comparison with Figure 5 indicates no significant differences to the previous results with respect to the effects on output and private consumption. The interest rate response shows somewhat less high-frequency variation but the broad patterns are similar to the previous results. Overall, imposing the stationarity conditions does therefore lead to very little changes in the main results documented above.

5 The Fiscal Transmission Mechanism

This section exploits the results obtained so far with the aim of disentangling the determinants of the effects of fiscal policy in the euro area. We first provide an account of existing views on the fiscal transmission mechanism in Section 5.1. We then investigate the driving forces of time variation in spending multipliers using simple regression analysis in Section 5.2. Section 5.3 investigates the robustness of the regression results.

5.1 Views on the transmission mechanism

Several potential determinants of the effectiveness of fiscal policy have received some attention recently: (i) the level of government debt, (ii) asset market participation and access to credit, (iii) the degree of trade openness. Using regression analysis we relate these factors to the observed time variation in spending multipliers. In addition, we study the effects of the composition of government spending according to (iv) the share of government investment and (v) the wage component of total spending. Before turning to

the results we provide an account of existing views on the fiscal transmission mechanism according to the above-mentioned determinants.

(i) Government debt. Experience from past fiscal consolidations suggests the possibility that in times of fiscal stress, characterized by high debt-to-GDP ratios, an economy's response to fiscal shocks changes qualitatively. That is, positive consumption growth was observed after prolonged and substantial deficit cuts. This is the hypothesis of "expansionary fiscal contractions" brought about by Giavazzi and Pagano (1990).[20] Investigating a quarterly panel of 19 OECD countries, Perotti (1999) finds that the effect of spending shocks on consumption switches signs if the initial financing needs of the government are small. He argues that the effect from initial conditions comes from the convexity of tax distortions: a (larger) expected increase in taxation tomorrow causes a (larger) decline in wealth and (larger) fall in consumption today. A significant and sustained reduction of government spending may then lead consumers to expect a permanent future tax cut and an increase in permanent income, leading to a rise in private consumption.

(ii) Credit. Another important channel through which spending shocks may affect the economy is the degree of asset market participation and the stringency of credit constraints. In the standard neoclassical model and in the basic New Keynesian model, expansionary government spending shocks tend to generate a crowding out of private consumption and therefore relatively small multipliers on GDP. The reason is the negative wealth effect on consumers induced by higher future tax payments, which makes them save more and consume less due to the consumption smoothing objective. However, credit constraints and limited asset market participation may dampen this effect and induce non-Ricardian behavior by consumers. If private agents consume a high share of their after-tax income, or if they are constrained in their access to credit, they do not or cannot save against a higher future tax burden. Galí et al. (2007) show that a government spending shock can generate an increase in aggregate consumption in a New Keynesian model conditional on having a relatively large fraction of liquidity constrained consumers (around 30%–50% of the population).

In addition, it has recently been argued that fiscal policy may be more effective in stabilizing real economic activity in periods of recessions. The reason is that in recessions credit constraints might bind across a wider range

[20]See also Giavazzi, Jappelli, and Pagano (2000). Giavazzi and Pagano (1990) study episodes of large fiscal consolidations in Denmark during 1983–1986 and in Ireland during 1987-1989. In these episodes the cyclically adjusted deficit as a share of GDP declined by 9.5% and 7.2% relative to the preconsolidation year and yet private consumption increased by 17.7% and 14.5% cumulatively. Alesina and Perotti (1996) identify similar episoes in several other European countries and Canada during the 1980s.

of agents, which will affect the transmission of fiscal policy shocks.[21] Roeger and in 't Veld (2009) allow for credit constrained households along the lines of the financial accelerator literature, thus allowing the stringency of credit constraints to vary over the business cycle. They argue that fiscal policy becomes a more effective tool for short-run economic stabilization, since the marginal propensity to consume out of current income increases during recessions. Tagkalakis (2008) provides empirical evidence for asymmetric effects of fiscal policy for a panel of nineteen OECD countries over the period 1970–2002. He shows that a tax cut in downturns increases private consumption more than a tax cut in upturns, whereas a spending shock has a larger effect on private consumption in downturns.

(iii) Openness. It is often claimed that fiscal multipliers depend on the degree of openness to trade.[22] In very open economies, domestic output will remain largely unaffected by a fiscal expansion since a large fraction of the intended stimulus falls on imports. Multipliers should then be smaller in Europe than, for instance, in the U.S. since Europe is more open towards international trade. Using a panel VAR approach, Beetsma, Giuliodori, and Klaassen (2008) show that a 1% of GDP increase in public spending in the European Union leads to a fall of the trade balance by 0.5% of GDP on impact and a peak fall of 0.8% of GDP, due to rising imports and falling exports. This compares to a 1.2% impact effect and a 1.6% peak rise in GDP due to the spending shock. With respect to time variation in fiscal multipliers, the effects of an increase in spending on GDP are then expected to be smaller the higher the degree of openness. Below we use the import share as a proxy for the degree of openness since the theory indicates that imports are the channel through which openness to trade should effect fiscal multipliers, whereas the effects of exports are less clear-cut.

(iv) Government investment. Although not all empirical studies find a growth-enhancing effect of public capital, there is now more consensus than in the past that public capital furthers economic growth.[23] It therefore seems important to investigate whether any change in the composition of spending according to consumption and investment expenditures has contributed to a

[21] In fact, empirical evidence suggests that asset market constraints on households and/or firms are more severe in recessions than in expansions. See Stiglitz and Weiss (1981), Bernanke and Gertler (1989), Fissel and Japelli (1990), Bernanke, Gertler, and Gilchrist (1996), and García, Lusardi, and Ng (1997).

[22] For instance, Perotti (2005) touches upon this claim but he argues that the increase in openness is probably too small to account for the decline in spending multipliers in OECD economies.

[23] See Romp and de Haan (2007) for a survey of empirical studies on the link between public capital and economic growth.

changing spending multiplier. From a theoretical perspective, general equilibrium models that account for public capital typically predict that increases in government investment can generate larger fiscal multipliers than increases in government consumption, so long as public capital is even slightly productive (see e.g. Baxter and King, 1993; Pappa, 2005; Straub and Tchakarov, 2007). The reason is that government investment has the aggregate demand effect of government absorption but also an additional aggregate supply effect by enhancing production and the marginal productivity of private labor and capital. On the other hand, Leeper, Walker, and Yang (2009b) have recently provided evidence that government investment projects in the U.S. are subject to substantial implementation lags. Delays between the authorization of a government spending plan and the completion of an investment project lead to a smaller or negative short-term multiplier, even if public capital is productive. Private investment and employment are then postponed until the public capital is on line such that in the short run private investment is lower and labor impacts may be small or even negative. Output can therefore fall in the short run in response to an increase in government investment.

(v) Wage component. More than half of government spending in the euro area consists of wage payments to government employees. Several studies emphasize that the distinction between goods purchases and employee compensation is important when assessing the impact of spending shocks on the macroeconomy. Finn (1998) shows that shocks to government employment tend to have different effects than shocks to goods purchases in the neoclassical model. Employment shocks tend to raise the real wage and thus act as a transfer to households, which dampens the (negative) wealth effect on consumption and labor supply. Pappa (2005) shows that government employment shocks have similar effects in a New Keynesian model. Using structural VAR analysis, Perotti (2007) shows that the responses of output and private consumption in the U.S. tend to be larger in response to a government employment shock compared to a goods spending shock.

An alternative interpretation of the effects of government employment and wages has recently been provided by Alesina and Ardagna (2009) in the context of an imperfect labor market, which is typically absent in both the standard neoclassical model and the basic New Keynesian model. They argue that a decrease in government employment reduces the probability of finding a job if not employed in the private sector, and a decrease in government wages decreases the worker's income if employed in the public sector. In both cases, the reservation utility of union members goes down and the wage demanded by the union for private sector workers decreases, increasing profits, investment and competitiveness. According to this argument, an increase in the wage component of government spending could lead to a smaller effect on

output. There is thus still some disagreement about the effects of government employment and wages, and we investigate below whether any adjustments in the composition of government spending in terms of goods purchases and employee compensation might have contributed to the observed changes in the effects of government spending shocks in the euro area.

5.2 Driving forces of time variation

Several testable hypotheses can be derived from Section 5.1. First, the effects of spending shocks on output and consumption are smaller the higher the initial debt-to-GDP ratio. Second, the effect of spending shocks are higher if households are more restricted in their access to credit, or if actual output is below potential output. Third, a higher share of imports over GDP leads to smaller spending multipliers. Fourth, a higher government investment share can lead to higher spending multipliers but if implementation lags play a role short-term multipliers can be smaller. Fifth, a higher wage share can result in larger or smaller effects on economic activity according to the degree of labor market competitiveness. In this section we address those various hypotheses by means of simple regression analysis. We apply Bayesian linear regressions, using the government spending multipliers on output and consumption from Section 4.2 as dependent variables.[24] We distinguish both short-term effects on contemporaneous multipliers and long-term effects on multipliers after five years. The fact that the dependent variables are estimated parameters, which may lead to biased standard errors, is addressed in Section 5.3.

Figure 10 shows the explanatory factors used in the regression analysis. The lagged aggregate euro area debt-to-GDP ratio is used to measure the initial financing needs of euro area governments. Access to credit is measured by the lagged ratio of credit to households over GDP.[25] The state of the business cycle is approximated by the lagged HP-filtered output gap. Lagged values are used to address reverse causation from spending multipliers on output and the business cycle. As discussed in Section 5.1, the

[24] We specify diffuse normal priors with mean zero and standard deviation 10^6 for the regression coefficients. In all regressions we control for a constant and a linear trend, in order to address potential concerns of spurious causation. Controlling for quadratic trend instead of a linear trend did not lead to any significant changes in the results. We furthermore account for the possible presence of heteroskedastic disturbances using diffuse priors on the variance terms. The regressions are estimated using a Gibbs sampling algorithm with 1,100 draws dropping the first 100 draws, see Geweke (1993) for details.

[25] We use the amount of outstanding loans to households in each quarter. The quarterly credit ratio is computed by dividing this measure by the sum of nominal GDP over the last four consecutive quarters. The data on loans were obtained from the Bank of International Settlements macroeconomic series.

ratio of imports over GDP (in lagged terms) is used to assess the impact of changes in the degree of openness on spending multipliers. Finally, we use the contemporaneous shares of government investment and employee compensation over total spending in order to assess the impact of changes in the composition of spending.

A first set of regression results is documented in Table 1. Using the medians of the contemporaneous multipliers on output and consumption as dependent variables, the explanatory factors are added one by one to the regression equations. The point estimates of the regression coefficients are the means of their posterior distribution. The statistical "significance" of the regression coefficients is measured in terms of the posterior probability that they are non-positive (non-negative) if their point estimates are positive (negative). The results indicate that an increase in the share of government debt over GDP has had a negative impact on contemporaneous spending multipliers. A one percentage point increase in the debt-to-GDP ratio has caused on average a decline in the multiplier on output by 0.01 points in all regression specifications considered except (1), whose explanatory power is however relatively low. Given an increase in the debt ratio by about 30 percentage points over the period 1980–2008, this is a fairly large contribution. The effect on the multiplier on consumption is closer to zero. We also estimate a negative effect on the size of spending multipliers of an increase in the ratio of credit to households over GDP. A one percentage point increase in the credit ratio leads on average to a decline (measured in terms of percent of GDP) in the spending multiplier on output (consumption) between 0.04 and 0.06 points (between 0.01 and 0.02 points). The credit ratio has increased from 30% in 1980 to almost 60% in 2008, such that increasing credit availability is estimated to have contributed substantially to the observed decline in spending multipliers. The output gap does however enter with an unexpected positive sign, but the coefficient on the multiplier on output is only positive with 90% probability in the largest regression model (6).

A rise in the share of imports over GDP is estimated to have a negative effect on the size of spending multipliers, a one percentage point increase in the import share leading on average to a decline in the multipliers on output and consumption by 0.01 points (except in the largest regression model for consumption). Finally, the impact of an increase in the share of government investment in total spending is estimated to be positive whereas an increase in the share of wage payments in total spending leads to a decline in spending multipliers. In the largest regression model for the output (consumption) multiplier, a one percentage point increase of the investment share is estimated to cause an average increase in the multiplier by 0.07 points (0.01 points). A one percentage point increase in the wage share, however, leads

to an average decrease in the multiplier by 0.05 points (0.01 points). We note that the adjusted R^2 is highest in regressions (6) and (12), providing support for the inclusion of all explanatory variables considered.

The results using median long-term spending multipliers after five years as dependent variables are reported in Table 2. In the discussion of results, we focus on the largest regression models (6) and (12). The results show that the output gap now enters with a negative sign, whereas the impact of the credit ratio and the debt ratio remains negative. On average, a one percentage point increase in the output gap leads to a decline in the long-term multipliers on output and consumption by 0.02 points and 0.01 points, respectively. A one percentage point increase in the credit ratio or the debt ratio lead to a decline in the spending multiplier on output by 0.01 points. The marginal effects on the consumption multiplier are again closer to zero. The impact of the import share, the share of government investment in total spending, and the wage share on the long-term multiplier on output is negligible. On the other hand, the impact of imports on the consumption multiplier is positive in the long-term. Contrary to the short-term multiplier, a higher wage share is estimated to have a positive impact on the long-term multiplier on consumption albeit with a small coefficient.

5.3 Robustness: standard error adjustment

A note of caution on the regression results reported in Tables 1 and 2 is in place. We have used the point estimates of spending multipliers as dependent variables in those regressions. Thereby we neglect the fact that the multipliers are themselves estimated parameters. This may give a biased view of the importance of the restrictions implied by the explanatory variables and artificially produce significant effects even when the "true" ones are negligible (see Canova and Pappa, 2006). One should therefore account for the uncertainty in the dependent variables, i.e. one needs to adjust the standard errors of the regression coefficients. We address this issue in the following way. We use each of 1,000 multipliers in the posterior distribution from the identified TVP-VAR in turn as dependent variable. Similar as above, we then generate 1,100 draws from the Gibbs sampler and omit the first 100 draws for each regression. This leaves us with 1,000,000 draws from the posterior distribution of regression coefficients from which we compute means and posterior probabilities. The results for contemporaneous multipliers and long-term spending multipliers, respectively, are reported in Tables 3 and 4.

The results in Table 3 show that the point estimates are similar but the standard errors of the explanatory variables are larger than previously, leading to the conclusion that some of them do not have an impact (with

at least 90% probability) on contemporaneous spending multipliers. The coefficient on the ratio of credit to households over GDP is however still negative with at least 99% probability. In regression (6) a one percentage point increase in the credit ratio leads on average to a decrease by 0.06 points (0.04 points) in the spending multiplier on output (consumption). Similarly, the share of government investment in total spending keeps to have a positive effect on spending multipliers, and the effect of government wages is again negative. The output gap is again estimated to have a positive effect of contemporaneous multipliers. However, government debt and imports drop as potential explanations for the effects of government spending on output and consumption according to regressions (6) and (12).

Table 4 indicates that accounting for the uncertainty in the dependent variables has an even larger impact on our results for the long-term spending multipliers. Only the ratio of government debt over GDP remains with a non-negligible effect on long-term multipliers. For both the multiplier on output and the multiplier on consumption, a one percentage point increase in the debt ratio leads on average to a decline by 0.01 points in the multipliers, the effect being negative with at least 95% probability in all regression models. The remaining variables stay with their previous signs, but their standard errors are too large for them to be significant driving forces of the fiscal transmission mechanism.

In summary, the second-stage regressions indicate that (i) the level of government debt has an adverse impact on the size of spending multipliers especially in the long run whereas its short-term impact turns unimportant once we account for the uncertainty in estimated multipliers. (ii) The ratio of credit over GDP is the main driving force of the observed time variation in contemporaneous spending multipliers. However, this effect does not immediately feed through to a higher effect of government spending on output or consumption during recessions. The output gap only has the expected negative effect on spending multipliers in the long run, but this effect cancels once we adjust standard errors. (iii) The negative impact on (short-term) multipliers of the degree of openness–measured by the share of imports over GDP–disappears once we control for the uncertainty in the dependent variables. With respect to compositional effects, (iv) a higher share of government investment in total spending has a positive effect on the size of spending multipliers in the short run, even when standard errors are adjusted. Finally, (v) a larger wage component of government spending leads to smaller short-term spending multipliers.

6 Conclusions

This paper has specified and estimated time-varying parameters vector autoregressions, with the aim of investigating changes in the effects of government spending shocks in the euro area over the period 1980–2008 and revealing the driving forces of the evolution of spending multipliers.

Our results indicate that–despite a relatively stable total fiscal impulse–the effectiveness of spending shocks in stimulating economic activity has decreased over time. Short-run spending multipliers increased until the late 1980s when they reached values above unity, but they started to decline afterwards to values closer to 0.5 in the current decade. Long-term multipliers show a more than two-fold decline since the 1980s. These results suggest that other components of aggregate demand are increasingly being crowded out by spending based fiscal expansions. In particular, the response of private consumption to government spending shocks has become substantially weaker over time. We also document a weaker response of real wages, whereas the nominal interest rate shows a stronger reaction to spending shocks.

With respect to the driving forces of time variation, our evidence points towards access to credit as one of the main determinants of short-term spending multipliers. This finding lends empirical support to the view that access to credit matters for the effectiveness of discretionary fiscal stimulus. The argument is that the presence of credit constraints and limited asset market participation reduces the importance of Ricardian equivalence, since a larger share of agents cannot borrow or save immediately against a higher future tax burden. The result that real wages show a weaker response to spending shocks seems also consistent with this view. It implies that current income reacts less strongly to spending shocks, which leads to a smaller increase in the consumption of credit constrained consumers.

We also conclude that a lower share of government investment and a larger wage component in total spending may have contributed to the observed decline in short-term multipliers. These findings indicate that government investment may have an additional positive aggregate supply effect in addition to the aggregate demand effect of government goods purchases. However, implementation lags do not seem to affect the size of spending multipliers since in that case we would expect a smaller short-term impact and larger long-term effects. The negative effect of wage payments on spending multipliers is consistent with arguments on the potential adverse consequences of increases in government employment and wages in an imperfect labor market.

Finally, our results suggest that rising government debt is the main reason for declining spending multipliers at longer horizons, and thus more adverse long-run consequences of fiscal expansions. We interpret this finding as an

indication that further accumulating debt after a spending shock leads to rising concerns on the sustainability of public finances, such that agents may expect a larger fiscal consolidation in the future which depresses private demand and output. We also find that a stronger response of the short-term nominal interest rate goes along with declining spending multipliers. This result is consistent with an increasingly offsetting reaction of monetary policy to the expansionary fiscal shock.

An important issue for future research would be to investigate the cross-country dimension of time variation in fiscal multipliers. Next to the fact that it would be useful to assess the robustness of our results for the aggregate euro area at the country level, such an investigation could contribute to the present study by adding variation in fiscal multipliers as well as explanatory variables. This would facilitate the identification of the factors which determine the effectiveness of fiscal policy, thus helping to further enhance our understanding of the fiscal transmission mechanism.

References

Alesina, A., and R. Perotti (1996): "Reducing Budget Deficits," *Swedish Economic Policy Review*, 3, 13–34.

Alesina, A. F., and S. Ardagna (2009): "Large Changes in Fiscal Policy: Taxes Versus Spending," NBER Working Paper No. 15438.

Baxter, M., and R. G. King (1993): "Fiscal Policy in General Equilibrium," *American Economic Review*, 83, 315–334.

Beetsma, R., M. Giuliodori, and F. Klaassen (2008): "The Effects of Public Spending Shocks on Trade Balances and Budget Deficits in the European Union," *Journal of the European Economic Association*, 6, 414–423.

Bénassy-Quéré, A., and J. Cimadomo (2006): "Changing Patterns of Domestic and Cross-Border Fiscal Policy Multipliers in Europe and the US," CEPII Working Paper No. 24.

Benati, L., and H. Mumtaz (2007): "U.S. Evolving Macroeconomic Dynamics: A Structural Investigation," ECB Working Paper No. 746.

Bernanke, B., and M. Gertler (1989): "Agency Costs, Net Worth and Business Cycle Fluctuations," *American Economic Review*, 79, 14–31.

Bernanke, B., M. Gertler, and S. Gilchrist (1996): "The financial accelerator and the flight to quality," *Review of Economics and Statistics*, 78, 1–15.

Bilbiie, F. O., A. Meier, and G. J. Mueller (2006): "What Accounts for the Changes in U.S. Fiscal Policy Transmission?," ECB Working Paper No. 582.

Blanchard, O. J., and R. Perotti (2002): "An Empirical Characterization of the Dynamic Effects of Changes in Government Spending and Taxes on Output," *Quarterly Journal of Economics*, 117, 1329–1368.

Burriel, P., F. de Castro, D. Garrote, E. Gordo, J. Paredes, and J. J. Pérez (2009): "Fiscal policy shocks in the euro area: an empirical assessment," ECB Working Paper No. 1133.

Caldara, D., and C. Kamps (2008): "What are the effects of fiscal policy shocks? A VAR-based comparative analysis," ECB Working Paper No. 877.

Canova, F. (2007): *Methods for Applied Macroeconomic Research*. Princeton University Press, Princeton.

Canova, F., and M. Ciccarelli (2006): "Estimating Multi-Country VAR Models," ECB Working Paper No. 603.

Canova, F., and L. Gambetti (2009): "Structural changes in the US economy: Is there a role for monetary policy?," *Journal of Economic Dynamics and Control*, 33, 477–490.

Canova, F., and E. Pappa (2006): "The elusive costs and the immaterial gains of fiscal constraints," *Journal of Public Economics*, 90, 1391–1414.

Carter, C. K., and R. Kohn (1994): "On Gibbs Sampling for State Space Models," *Biometrika*, 81, 541–553.

Christoffel, K., G. Coenen, and A. Warne (2008): "The New Area-Wide Model of the Euro Area: A Micro-Founded Open-Economy Model for Policy Analysis and Forecasting," ECB Working Paper No. 944.

Cogan, J. F., T. Cwik, J. B. Taylor, and V. Wieland (2009): "New Keynesian Versus Old Keynesian Government Spending Multipliers," NBER Working Paper No. 14782.

Cogley, T., and T. J. Sargent (2001): "Evolving Post-World War II U.S. Inflation Dynamics," in *NBER Macroeconomics Annual*, vol. 16. MIT Press, Cambridge, MA.

——— (2005): "Drifts and Volatilities: Monetary Policies and Outcomes in the Post WWII U.S.," *Review of Economic Studies*, 8, 262–302.

D'Agostino, A., L. Gambetti, and D. Giannone (2009): "Macroeconomic Forecasting and Structural Change," ECARES Working Paper No. 2009-020.

European Commission (2009): "European Economic Forecast – autumn 2009," European Economy No. 10.

Fagan, G., J. Henry, and R. Mestre (2005): "An area-wide model for the euro area," *Economic Modelling*, 22, 39–59.

Fatás, A., and I. Mihov (2001): "The Effects of Fiscal Policy on Consumption and Employment: Theory and Evidence," CEPR Discussion Paper No. 2760.

Finn, M. G. (1998): "Cyclical Effects of Government's Employment and Goods Purchases," *International Economic Review*, 39, 635–657.

Fissel, G., and T. Japelli (1990): "Do liquidity constraints vary over time? Evidence from survey and panel data," *Journal of Money, Credit and Banking*, 22, 253–262.

Forni, L., L. Monteforte, and L. Sessa (2009): "The general equilibrium effects of fiscal policy: estimates for the euro area," *Journal of Public Economics*, 93, 559–585.

Galí, J., and L. Gambetti (2009): "On the Sources of the Great Moderation," *American Economic Journal: Macroeconomics*, 1, 26–57.

Galí, J., J. D. Lopéz-Salido, and J. Vallés (2007): "Understanding the Effects of Government Spending on Consumption," *Journal of the European Economic Association*, 5, 227–270.

García, R., A. Lusardi, and S. Ng (1997): "Excess sensitivity and asymmetries in consumption: an empirical investigation," *Journal of Money, Credit and Banking*, 29, 154–176.

Geman, S., and D. Geman (1984): "Stochastic Relaxation, Gibbs Distributions, and the Bayesian Restoration of Images," *IEEE Transactions on Pattern Analysis and Machine Intelligence*, 6, 721–741.

Geweke, J. (1992): "Evaluating the Accuracy of Sampling-Based Approaches to the Calculation of Posterior Moments," in *Bayesian Statistics*, ed. by J. M. Bernardo, J. Berger, A. P. Dawid, and A. F. M. Smith. Oxford University Press, Oxford.

——— (1993): "Bayesian Treatment of the Independent Student-t Linear Model," *Journal of Applied Econometrics*, 8, S19–S40.

Giavazzi, F., T. Jappelli, and M. Pagano (2000): "Searching for Non-linear Effects of Fiscal Policy: Evidence from Industrial and Developing Countries," *European Economic Review*, 2000, 1259–1289.

Giavazzi, F., and M. Pagano (1990): "Can Severe Fiscal Adjustments Be Expansionary?," in *NBER Macroeconomics Annual*, ed. by O. Blanchard, and S. Fischer. MIT Press, Cambridge, MA.

Jacquier, E., N. G. Polson, and P. E. Rossi (1994): "Bayesian Analysis of Stochastic Volatility Models," *Journal of Business and Economic Statistics*, 12, 371–389.

Kim, C., and C. R. Nelson (1999): *State-Space Models with Regime Switching*. MIT Press, Cambridge, MA.

Koop, G. (2003): *Bayesian Econometrics*. Wiley & Sons.

Kriwoluzky, A. (2009): "Pre-Announcement and Timing: The Effects of a Government Expenditure Shock," EUI Working Paper No. 38.

Leeper, E. M., T. B. Walker, and S.-C. S. Yang (2009a): "Fiscal Foresight and Information Flows," NBER Working Paper No. 14630.

——— (2009b): "Government Investment and Fiscal Stimulus in the Short and Long Runs," NBER Working Paper No. 15153.

Linnemann, L., and A. Schabert (2003): "Fiscal Policy in the New Neoclassical Synthesis," *Journal of Money, Credit and Banking*, 35, 911–929.

——— (2004): "Can fiscal spending stimulate private consumption?," *Economics Letters*, 82, 173–179.

Mertens, K., and M. O. Ravn (2009): "Empirical Evidence on the Aggregate Effects of Anticipated and Unanticipated U.S. Tax Policy Shocks," CEPR Discussion Paper No. 7370.

Monacelli, T., and R. Perotti (2008): "Fiscal Policy, Wealth Effects and Markups," NBER Working Paper No. 14584.

Mountford, A., and H. Uhlig (2009): "What are the Effects of Fiscal Policy Shocks?," *Journal of Applied Econometrics*, 24, 960–992.

Pappa, E. (2005): "New-Keynesian or RBC Transmission? The Effects of Fiscal Shocks in Labour Markets," CEPR Discussion Paper No. 5313.

Paredes, J., J. Pérez, and D. J. Pedregal (2009): "A quarterly fiscal database for the euro area based on intra-annual fiscal indicators," ECB Working Paper No. 1132.

Perotti, R. (1999): "Fiscal Policy in Good Times and Bad," *Quarterly Journal of Economics*, 138, 1399–1435.

——— (2005): "Estimating the Effects of Fiscal Policy in OECD countries," CEPR Discussion Paper No. 4842.

——— (2007): "In Search of the Transmission Mechanism of Fiscal Policy," NBER Working Paper No. 13143.

Primiceri, G. (2005): "Time Varying Structural Vector Autoregressions and Monetary Policy," *Review of Economic Studies*, 72, 821–852.

Raftery, A. E., and S. Lewis (1992): "How Many Iterations in the Gibbs Sampler?," in *Bayesian Statistics*, ed. by J. M. Bernardo, J. Berger, A. P. Dawid, and A. F. M. Smith. Oxford University Press, Oxford.

Ramey, V. A. (2009): "Identifying Government Spending Shocks: It's All in the Timing," *mimeo*, University of California San Diego.

Ramey, V. A., and M. Shapiro (1998): "Costly Capital Reallocation and the Effects of Government Spending," *Carnegie-Rochester Conference Series on Public Policy*, 48, 145–194.

Ratto, M., W. Roeger, and J. in 't Veld (2009): "QUEST III: An estimated open-economy DSGE model of the euro area with fiscal and monetary policy," *Economic Modelling*, 26, 222–233.

Roberts, G. O., and A. F. M. Smith (1994): "Simple Conditions for the Convergence of the Gibbs Sampler and Metropolis-Hastings Algorithms," *Stochastic Processes and their Applications*, 49, 207–216.

Roeger, W., and J. in 't Veld (2009): "Fiscal Policy with Credit Constrained Households," European Commission Economic Papers No. 357.

Romp, W., and J. de Haan (2007): "Public Capital and Economic Growth: A Critical Survey," *Perspektiven der Wirtschaftspolitik*, 8, 6–52.

Sims, C. A., and T. Zha (1999): "Error Bands for Impulse Responses," *Econometrica*, 67, 1113–1155.

Smets, F., and R. Wouters (2003): "An Estimated Dynamic Stochastic General Equilibrium Model of the Euro Area," *Journal of the European Economic Association*, 1, 1123–1175.

Smith, A. F. M., and G. O. Roberts (1993): "Bayesian Computation Via the Gibbs Sampler and Related Markov Chain Monte Carlo Methods," *Journal of the Royal Statistical Society Series B*, 55, 3–23.

Stiglitz, J., and A. Weiss (1981): "Credit Rationing with Imperfect Information," *American Economic Review*, 71, 393–410.

Straub, R., and I. Tchakarov (2007): "Assessing the Impact of a Change in the Composition of Public Spending," ECB Working Paper No. 795.

Tagkalakis, A. (2008): "The effects of fiscal policy on consumption in recessions and expansions," *Journal of Public Economics*, 92, 1486–1508.

Uhlig, H. (2005): "What are the effects of monetary policy on output? Results from an agnostic identification procedure," *Journal of Monetary Economics*, 52, 381–419.

A Details of the Gibbs sampler

This appendix outlines the details of the Gibbs sampling algorithm used for estimation of the TVP-VAR. The algorithm generates a Markov chain which is a sample the joint posterior distribution of the VAR parameters (i.e. coefficient states, covariance states, volatility states, and hyperparameters). It combines elements of Cogley and Sargent (2005), Primiceri (2005), and Benati and Mumtaz (2007), with a few additional restrictions on the structure of the hyperparameters. In what follows, x^t denotes the history of x up to time t, i.e. $x^t = [x_1', x_2', \ldots, x_t']'$, and T denotes the sample length. Furthermore, rewrite the observation equation (2) conveniently as

$$y_t = X_t'\beta_t + u_t \tag{6}$$

where $X_t' = I \otimes [1, y_{t-1}', \ldots, y_{t-p}']$. The estimation proceeds in four steps.

A.1 Drawing coefficient states β^T

Conditional on A^T and H^T one obtains a history R^T. Then, conditional on y^T, R^T and Q, the observation equation (2) is linear with Gaussian innovations and a known covariance matrix. The posterior density of the coefficients can be factored as[26]

$$f(\beta^T|y^T, R^T, Q) = f(\beta_T|y^T, R^T, Q) \prod_{t=1}^{T-1} f(\beta_t|\beta_{t+1}, y^t, R_t, Q) \tag{7}$$

where

$$\begin{aligned} \beta_t|\beta_{t+1}, y^t, R^T, Q &\sim N(\beta_{t|t+1}, P_{t|t+1}) \\ \beta_{t|t+1} &\sim \mathrm{E}[\beta_t|\beta_{t+1}, y^t, R^T, Q] \\ P_{t|t+1} &\sim \mathrm{E}[P_t|P_{t+1}, y^t, R^T, Q] \end{aligned}$$

The conditional means and variances can be computed using the Kalman filter and a backward recursion (see Carter and Kohn, 1994). The Kalman filter delivers

$$\begin{aligned} P_{t|t-1} &= P_{t-1|t-1} + Q \\ K_t &= P_{t|t-1}X_t(X_t'P_{t|t-1}X_t + R_t)^{-1} \\ \beta_{t|t} &= \beta_{t-1|t-1} + K_t(y_t - X_t'\beta_{t-1|t-1}) \\ P_{t|t} &= P_{t|t-1} - K_tX_t'P_{t|t-1} \end{aligned}$$

[26] We omit conditioning factors which are redundant in the respective step.

The initial values $\beta_{0|0}$ for this recursion are the OLS point estimates from the initial sample, and the initial value $P_{0|0}$ is their covariance matrix. The initial R_t is the OLS covariance matrix of the reduced-form VAR. The covariance matrix Q is a scaled version of the variance-covariance matrix of the coefficients.

The Kalman filter delivers as its last points $\beta_{T|T}$ and $P_{T|T}$. Draws from (7) are then obtained by a backward recursion. The first point in the backward recursion is a draw from $N(\beta_{T|T}, P_{T|T})$. The remaining draws are from $N(\beta_{t|t+1}, P_{t|t+1})$ where the means and variances are derived as follows:

$$\begin{aligned} \beta_{t|t+1} &= \beta_{t|t} + P_{t|t}P_{t+1|t}^{-1}(\beta_{t+1} - \beta_{t|t}) \\ P_{t|t+1} &= P_{t|t} - P_{t|t}P_{t+1|t}^{-1}P_{t|t} \end{aligned}$$

A.2 Drawing covariance states A^T

Conditional on y^T, β^T and H^T, write the system of equations (6) as

$$A_t(y_t - X_t'\beta_t) = A_t\widehat{y}_t = H_t^{1/2}v_t \tag{8}$$

Moreover, A_t is lower diagonal (with ones on the main diagonal) such that (8) can be rewritten as

$$\widehat{y}_t = Z_t\alpha_t + H_t^{1/2}v_t \tag{9}$$

where α_t is defined as in the main text and Z_t has the structure

$$Z_t = \begin{bmatrix} 0 & \cdots & \cdots & 0 \\ -\widehat{y}_{1,t} & 0 & \cdots & \vdots \\ 0 & (-\widehat{y}_{1,t}, -\widehat{y}_{2,t}) & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & (-\widehat{y}_{1,t}, \cdots, -\widehat{y}_{n-1,t}) \end{bmatrix}$$

where n denotes the number of variables in the VAR. The system of equations (9) has a Gaussian but non-linear state-space form. However, under the assumption of (block) diagonality of S the problem becomes linear (see Primiceri, 2005). The forward (Kalman filter) and backward recursions of the previous step can then be applied equation by equation. Hence, the procedure allows to recover α^T by

$$\begin{aligned} \alpha_{i,t|t+1} &= \mathrm{E}[\alpha_{i,t}|\alpha_{i,t+1}, y^t, \beta^T, H^T, S_i] \\ \Lambda_{i,t|t+1} &= \mathrm{var}[\alpha_{i,t}|\alpha_{i,t+1}, y^t, \beta^T, H^T, S_i] \end{aligned}$$

where $\alpha_{i,t}$ is the block of α_t corresponding to the i-th equation and S_i is the associated i-th block of S. The initial values for the Kalman filter are obtained from a decomposition of the OLS covariance matrix, using the prior mean and the prior variance of α_0 as described in Appendix B.

A.3 Drawing volatility states H^T

To sample the stochastic volatilities the univariate algorithm of Jacquier, Polson, and Rossi (1994) is applied each element of H_t. The orthogonalized residuals $v_t = A_t u_t$ are

observable conditional on y^T, β^T and A^T. We can use the univariate setting because the stochastic volatilities are assumed to be independent, following Cogley and Sargent (2005). Jacquier, Polson, and Rossi (1994) show that the conditional kernel is

$$f(h_{i,t}|h_{-i,t}, v_i^T, w_i) \propto f(h_{i,t}|h_{i,t-1}, h_{i,t+1}, v_i^T, w_i)$$

where w_i is the i-th diagonal element of W and $h_{-i,t}$ represents the vector of h's at all other dates. Using Bayes' theorem the above conditional kernel can be expressed as

$$\begin{aligned} f(h_{i,t}|h_{i,t-1}, h_{i,t+1}, v_i^T, w_i) &\propto f(u_{i,t}|h_{i,t}) f(h_{i,t}|h_{i,t-1}) f(h_{i,t+1}|h_{i,t}) \\ &\propto h_{i,t}^{-1.5} \exp\left(-\frac{v_{i,t}^2}{2h_{i,t}}\right) \exp\left(-\frac{(\ln h_{i,t} - \mu_{i,t})^2}{2\sigma_{ic}^2}\right) \end{aligned} \tag{10}$$

where $\mu_{i,t}$ and σ_{ic}^2 are the conditional mean and variance of $h_{i,t}$ implied by (5) and knowledge of $h_{i,t-1}$ and $h_{i,t+1}$. For a geometric random walk these parameters are

$$\mu_{i,t} = 0.5(\log h_{i,t-1} + \log h_{i,t+1}) \quad \text{and} \quad \sigma_{ic}^2 = 0.5 w_i$$

In practice $h_{i,t+1}$ is taken from the previous Gibbs iteration.[27] Jacquier, Polson, and Rossi (1994) propose a Metropolis step instead of a Gibbs step, because the normalizing constant is expensive to calculate in (10). Hence, one draws from a stand-in density and then uses the conditional likelihood $f(u_{i,t}|h_{i,t})$ to calculate the acceptance probability for that draw. Cogley and Sargent (2005) suggest to use the log-normal implied by (5) as the stand-in density:

$$g(h_{i,t}) \propto h_{i,t}^{-1} \exp\left(-\frac{(\log h_{i,t} - \mu_{i,t})^2}{2\sigma_{ic}^2}\right)$$

The acceptance probability for the m-th draw is

$$\begin{aligned} q_m &= \frac{f(v_{i,t}|h_{i,t}^m) g(h_{i,t}^m)}{g(h_{i,t}^m)} \frac{g(h_{i,t}^{m-1})}{f(v_{i,t}|h_{i,t}^{m-1}) g(h_{i,t}^{m-1})} \\ &= \frac{(h_{i,t}^m)^{-1/2} \exp\left(-v_{i,t}^2/2h_{i,t}^m\right)}{(h_{i,t}^{m-1})^{-1/2} \exp\left(-v_{i,t}^2/2h_{i,t}^{m-1}\right)} \end{aligned}$$

where $h_{i,t}^m = h_{i,t}^{m-1}$ if the draw is rejected. This algorithm is applied on a date by date basis to each element of u_t. The formulas are slightly different for the first and last element. For the first element we have

$$\mu_{i1} = \sigma_{ic}^2 \left(\frac{\mu_{i0}}{\sigma_{hi0}^2} + \frac{\log h_{i,t+1}}{w_i}\right) \quad \text{and} \quad \sigma_{ic}^2 = \frac{\sigma_{hi0}^2 w_i}{\sigma_{hi0}^2 + w_i}$$

and the acceptance probability is 1 since there is no previous draw. For the last element we have

$$\mu_{iT} = \log h_{i,t-1} \quad \text{and} \quad \sigma_{ic}^2 = w_i$$

where the prior on the distribution of $\log h_0$, providing values for the mean μ_{i0} and the variance σ_{hi0}^2, is described in Appendix B.

[27] In the first iteration, we use squared orthogonalized residuals $v_{i,t}^2$ in order to initialize the volatilities, which are obtained from the application of the OLS estimates from the initial sample on the actual sample.

A.4 Drawing hyperparameters

The hyperparameters of the model are the covariance matrices of the innovations, i.e. Q (coefficient states), S (covariance states), and W (volatility states). Conditional on y^T, β^T, A^T and H^T, these state innovations are observable. Since the hyperparameters are assumed to be independent, each covariance matrix can be drawn from its respective distribution.

Since we have restricted the hyperparameter matrix Q to be diagonal, its diagonal elements q_i have univariate inverse Gamma distributions with scale parameter $\gamma^q_{i,1}$ and degrees of freedom δ^q_1:

$$f(q_i|y^T, \beta^T) = \text{IG}\left(\frac{\gamma^q_{i,1}}{2}, \frac{\delta^q_1}{2}\right)$$

where $\delta^q_1 = \delta^q_0 + T$ and $\gamma^q_{i,1} = \gamma^q_{i,0} + \sum_{t=1}^T \varepsilon^2_{i,t}$ (see e.g. Kim and Nelson, 1999).

Similarly, restricting S to be diagonal, each of its diagonal elements s_i has an inverse Gamma distribution with scale parameter $\gamma^s_{i,1}$ and degrees of freedom δ^s_1:

$$f(s_i|y^T, A^T) = \text{IG}\left(\frac{\gamma^s_{i,1}}{2}, \frac{\delta^s_1}{2}\right)$$

where $\delta^s_1 = \delta^s_0 + T$ and $\gamma^s_{i,1} = \gamma^s_{i,0} + \sum_{t=1}^T \nu^2_{i,t}$.

Finally, the diagonal elements w_i of W have univariate inverse Gamma distributions with scale parameter $\gamma^w_{i,1}$ and degrees of freedom δ^w_1:

$$f(w_i|y^T, H^T) = \text{IG}\left(\frac{\gamma^w_{i,1}}{2}, \frac{\delta^w_1}{2}\right)$$

where $\delta^w_1 = \delta^w_0 + T$ and $\gamma^w_{i,1} = \gamma^w_{i,0} + \sum_{t=1}^T \omega^2_{i,t}$.

A.5 Summary

The Gibbs sampling algorithm is summarized as follows:

1. Initialize R^T, Q, S, and W.
2. Draw coefficients β^T from $f(\beta^T|y^T, R^T, Q)$.
3. Draw covariances A^T from $f(A^T|y^T, H^T, S)$.
4. Draw volatilities H^T from $f(H^T|y^T, \beta^T, A^T, W)$.
5. Draw hyperparameters from $f(q_i|y^T, \beta^T)$, $f(s_i|y^T, A^T)$, and $f(w_i|y^T, H^T)$.
6. Go to step 2.

B Calibration of the priors

This appendix discusses the choice of our priors. We closely follow common choices in the TVP-VAR literature and impose relatively conservative priors, particularly on the amount of time variation in the data (see e.g. Cogley and Sargent, 2001; Cogley and Sargent, 2005; Primiceri, 2005; Benati and Mumtaz, 2007). However, unlike most previous studies those priors are not calibrated based on OLS estimates from an initial "training

sample" which is then discarded. This would mean sacrificing part of our already relatively short sample. Instead, we calibrate our priors based on OLS estimates from the full sample. Such a strategy is suggested by Canova (2007) and Canova and Ciccarelli (2006) for cases where a training sample is not available. A fixed-coefficient VAR model is thus estimated by OLS (equation by equation) on the full sample from 1980Q1–2008Q4.

B.1 VAR coefficients

Let $\hat{\beta}$ denote the OLS estimate of the VAR coefficients, and $\widehat{\Xi}$ their covariance matrix. We set

$$\beta_0 \sim \mathrm{N}(\hat{\beta}, 4 \times \widehat{\Xi})$$

where the variance scaling factor increases the uncertainty about the size of the VAR coefficients in the initial sample versus the actual sample.

B.2 Elements of H_t

Denote the OLS estimate of the VAR covariance matrix as $\widehat{\Sigma}$. We apply a triangular decomposition of this matrix similar to (4), $\widehat{\Sigma} = \widehat{\Psi}^{-1}\widehat{\Phi}(\widehat{\Psi}^{-1})'$, and denote the vector of diagonal elements of $\widehat{\Phi}$ as ϕ_0. Our prior for the diagonal elements of the matrix H_t is

$$h_0 \sim \mathrm{N}(\phi_0, 10 \times I)$$

The variance scaling factor 10 is arbitrary but large relative to the mean ϕ_0.

B.3 Elements of A_t

Denote the vector of non-zero off-diagonal elements of $\widehat{\Psi}$ as ψ_0, ordered by rows. The prior for the elements of A_t is

$$\alpha_0 \sim \mathrm{N}\left(\psi_0, 10 \times \mathrm{diag}(\psi_0)\right)$$

where the variance of α_0 is scaled up taking into the magnitude of the respective elements of the mean ψ_0, as in Benati and Mumtaz (2007).

B.4 Hyperparameters

The prior on the diagonal elements of the coefficient state error variance Q is also inverse Gamma:

$$q_i \sim \mathrm{IG}\left(\frac{\gamma^q_{i,0}}{2}, \frac{\delta^q_0}{2}\right)$$

where $\gamma^q_{i,0} = k_Q \times \hat{\xi}_i$, where $\hat{\xi}_i$ denotes the i-th diagonal element of the OLS covariance matrix $\widehat{\Xi}$ and $k_Q = 10^{-4}$. Hence, our prior attributes only 0.01% of the uncertainty surrounding the OLS estimates to time variation following Cogley and Sargent (2001). The degrees of freedom δ^q_0 are set to 1, which is the minimum for the prior to be proper. We thus put as little weight on the prior as possible.

The prior on the diagonal elements of the hyperparameter matrix S for the covariance states is also inverse Gamma:

$$s_i \sim \mathrm{IG}\left(\frac{\gamma^s_{i,0}}{2}, \frac{\delta^s_0}{2}\right)$$

where $\gamma_{i,0}^s = k_S \times \hat{\psi}_i$, where $\hat{\psi}_i$ denotes the i-th diagonal element of the OLS covariance matrix $\widehat{\Psi}$ and $k_S = 10^{-2}$. Here we follow Primiceri (2005), who makes similar choices for a block diagonal structure of S. The degrees of freedom δ_0^s are again set to the minimum value of 1.

The prior on the diagonal elements of the variance W for the volatility states is inverse Gamma:

$$w_i \sim \text{IG}\left(\frac{\gamma_{i,0}^w}{2}, \frac{\delta_0^w}{2}\right)$$

where $\gamma_{i,0}^w = k_W$. We set $k_W = 10^{-4}$ and $\delta_0^w = 1$. The parameters of the distribution are the same as in Cogley and Sargent (2005) and Benati and Mumtaz (2007).

C Convergence of the Markov chain

This appendix assesses the convergence of the Markov chain produced by the Gibbs sampler. We apply three types of convergence checks to the VAR coefficients, the covariances, and the volatilities.[28] We omit the hyperparameters, since these are not the direct objects of the analysis in this paper.

The first convergence check is the diagnostics due to Raftery and Lewis (1992), which is used to assess the total number of iterations required to achieve a certain precision, and the minimum burn-in period and thinning factor. The parameters for the diagnostic are specified as follows: quantile = 0.025; desired accuracy = 0.025; required probability of attaining the required accuracy = 0.95. We generate a Markov chain with 5,000 draws as suggested by Raftery and Lewis (1992) which is used as an input for the diagnostics. Table 5 reports the diagnostics. For all three state vectors, the required number of runs is far below the total number of iterations actually applied. The same holds for the number of burn-in replications and the thinning factor. The choices made to generate the Markov chain therefore seem appropriate.

Our second convergence diagnostic are the inefficiency factors (IFs) for the posterior estimates of the parameters. The IF is the inverse of Geweke's (1992) relative numerical efficiency measure, i.e. $IF = 1 + 2\sum_{k=1}^{\infty} \rho_k$, where ρ_k is the k-th order autocorrelation of the chain. This diagnostic therefore serves to judge how well the chain mixes. Low autocorrelations suggest that the draws are close to independent, which increases the efficiency of the algorithm (Primiceri, 2005). We use a 4% tapered window for the estimation of the spectral density at frequency zero. Values of the IFs below or around 20 are regarded as satisfactory, according to Primiceri (2005). The left panels of Figure 11 report the IFs for the state vectors. The IFs are far below 20 for the coefficients and the covariances, but around 30-35 for the volatilities. Compared to the results reported e.g. in Primiceri (2005) and considering the higher dimensionality of our problem, however, these results still seem satisfactory.

The final convergence test applied is the convergence diagnostic (CD) due to Geweke (1992). This diagnostic is based on the idea that, if a sufficiently large number of draws have been taken, the posterior estimates based on the first half of draws should be essentially the same as the estimates based on the second half of draws. If they are very different, either too few draws have been taken and estimates are inaccurate or the effects of the initial values of the chain have not worn off (Koop, 2003). We therefore divide the 1,000 draws from the posterior distribution into a first set of $N_1 = 100$ draws, a middle set

[28] See Koop (2003), chapter 4, for a review of convergence diagnostics.

of 500 draws, and a last set of $N_2 = 400$ draws as suggested by Koop (2003). We drop the middle set of draws and therefore make it likely that the first and last set are independent of each other. The convergence diagnostic is given by

$$CD = \frac{\widehat{\theta}_1 - \widehat{\theta}_2}{\widehat{\sigma}_1/\sqrt{N_1} + \widehat{\sigma}_2/\sqrt{N_2}} \to \mathrm{N}(0,1)$$

by a central limit theorem, where $\widehat{\theta}_i$ and $\widehat{\sigma}_i/\sqrt{N_i}$ denote the posterior means of the parameters and their numerical standard errors based on the i-th set of draws, for $i = 1, 2$. We plot the p-values for the null hypothesis that the set of draws are the same in the right panels of Figure 11. The p-values are typically larger than conventional significance levels for the VAR coefficients and the covariances, indicating that a sufficiently large number of draws has been taken for these parameters. However, the null hypothesis is often rejected for the volatilities.

To summarize, the coefficients and covariances have in general better convergence properties than the volatilities. Since the focus of our analysis is on impulse responses which are determined by the contemporaneous relations among variables and the propagation mechanism rather than the size of stochastic shocks we conclude that the convergence properties of the Markov chain are satisfactory.

D Performance of the estimation algorithm

This appendix investigates the performance of the estimation algorithm using results obtained from a Monte Carlo exercise. It also motivates our prior choices and some of the additional restrictions imposed on the TVP-VAR, such as diagonality of the hyperparameter matrices Q and S. The Monte Carlo exercise consists of creating a bivariate data set y^T based on model (2) with $p = 2$ lags. This model has been simulated for 1050 periods with smoothly evolving "true" underlying states β^T, α^T and h^T. The idea is that the complete model (2)-(5) should be able to retrieve the underlying states based on the simulated data y^T. Since the focus of this paper has been on impulse responses, we evaluate the performance of the estimation algorithm in terms of its ability to reproduce the true impulse responses based on the underlying states.[29]

We show three figures obtained from this exercise. Figure 12 is based on a TVP-VAR with unrestricted Q and block diagonal S, as e.g. in Primiceri (2005). Figure 13 results from restricting Q to be diagonal and leaving S block diagonal. Finally, Figure 14 is based on a TVP-VAR which restricts both Q and S to be diagonal. In all figures, the upper left charts show the underlying "truth" whereas the upper right charts are based on our baseline choice of priors. The middle left and middle right charts are based on priors which allow for more and less time variation, respectively, varying the scaling factors k_Q, k_S and k_W which calibrate prior beliefs on the variance of shocks hitting the state equations (3)-(5). The lower left charts result from choosing values for k_Q, k_S and k_W close to zero, and the lower right charts are the impulse responses implied by a time invariant VAR estimated by OLS.[30]

[29] The specification of the TVP-VAR follows the main text, using 15,000 iterations of the Gibbs sampler and dropping the first 5,000 iterations. Details including the states used to simulate the model and the state estimates are available from the authors upon request.

[30] In each graph we plot the reduced-form impulse responses of the second variable due to an innovation in the first variable corresponding to every 5-th observation.

Figure 12 shows that the baseline specification of the TVP-VAR with unrestricted Q and block diagonal S has some trouble in reproducing the true impulse responses, whereas the specification with a larger prior time variation seems to come somewhat closer. As we reduce the prior scaling factors k_Q, k_S and k_W the model implied impulse responses quickly approach the OLS implied responses. Hence, although there is time variation in the true impulse responses and despite a relatively large sample, the estimation algorithm is not able to pick up this time variation. Figure 13 indicates that restricting the coefficient hyperparameter matrix Q to be diagonal helps the baseline specification to come closer to the truth, and even with smaller scaling factors k_Q, k_S and k_W the estimation algorithm still picks up some time variation. Finally, Figure 14 shows that restricting also the covariance hyperparameter matrix S to be diagonal makes the estimation algorithm more robust to the specific choice of priors. The baseline specification picks up the underlying truth fairly well, but also the specifications with somewhat more and somewhat less time variation do a good job in matching the truth. However, if prior scaling factors are set to zero the model implied impulse responses again resemble the OLS implied impulse responses.

Overall, this Monte Carlo exercise indicates a satisfactory performance of the estimation algorithm. Importantly, the exercise has shown that the performance of the algorithm tends to improve if the amount of estimated parameters is reduced by imposing restrictions on the hyperparameter matrices Q and S.

Table 1: Bayesian linear regressions, dependent variables are contemporaneous multipliers.[a,b,c,d]

	Multiplier on Output						Multiplier on Consumption					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Gov. Debt over GDP (-1)	0.01***	-0.01***	-0.01***	-0.01***	-0.01**	-0.01*	0.00***	-0.00***	-0.00***	-0.00***	-0.00*	-0.00*
	(0.00)		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Credit over GDP (-1)		-0.06***	-0.06***	-0.04***	-0.04***	-0.06***		-0.02***	-0.02***	-0.01***	-0.01***	-0.02***
		(0.00)	(0.00)	(0.00)	(0.01)	(0.01)		(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
Output Gap (-1)			0.02**	0.04***	0.03***	0.02*			0.01***	0.02***	0.01***	0.01**
			(0.01)	(0.01)	(0.01)	(0.01)			(0.00)	(0.01)	(0.01)	(0.01)
Imports over GDP (-1)				-0.01**	-0.01*	-0.01*				-0.01**	-0.01**	-0.00
				(0.01)	(0.01)	(0.01)				(0.00)	(0.01)	(0.01)
Investment Share					0.03*	0.07***					0.00	0.01**
					(0.02)	(0.02)					(0.01)	(0.01)
Wage Share						-0.05***						-0.01***
						(0.01)						(0.01)
Constant	0.82***	3.04***	2.94***	2.90***	2.10***	4.91***	0.46***	1.08***	1.04***	1.02***	0.93***	1.83***
	(0.12)	(0.17)	(0.18)	(0.17)	(0.52)	(0.88)	(0.03)	(0.06)	(0.06)	(0.06)	(0.20)	(0.37)
Trend	-0.01***	0.01***	0.01***	0.01***	0.01***	0.01***	-0.00***	0.00***	0.00***	0.00**	0.00*	0.00***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	112	112	112	112	112	112	112	112	112	112	112	112
Adjusted R^2	0.69	0.89	0.90	0.90	0.90	0.91	0.69	0.88	0.88	0.88	0.88	0.90

[a] Bayesian regressions allow for heteroskedastic errors following Geweke (1993). Dependent variables are posterior means of the posterior distribution of impulse responses from the identified TVP-VAR. All regressions are estimated using a Gibbs sampling algorithm with 1,100 draws and 100 omitted draws. This leaves us with 1,000 posterior draws of regression coefficients.

[b] Multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t).

[c] Point estimates are posterior means of the posterior distribution. Standard deviations are reported in parentheses. Asterisks indicate posterior probabilities that the regression coefficients are non-positive if the point estimates are positive or non-negative if the point estimates are negative (*less than 10%, **less than 5%, ***less than 1%.)

[d] Explanatory variables are measured in percent.

Table 2: Bayesian linear regressions, dependent variables are multipliers after five years.[a,b,c,d]

	Multiplier on Output						Multiplier on Consumption					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Gov. Debt over GDP (-1)	-0.01***	-0.01***	-0.01***	-0.01***	-0.01***	-0.01***	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Credit over GDP (-1)		-0.01***	-0.01***	-0.01***	-0.01***	-0.01***		-0.00**	-0.00**	-0.01***	-0.01***	-0.00**
		(0.00)	(0.00)	(0.01)	(0.01)	(0.01)		(0.00)	(0.01)	(0.01)	(0.00)	(0.00)
Output Gap (-1)			-0.02***	-0.02***	-0.02***	-0.02***			-0.00	-0.01*	-0.01**	-0.01***
			(0.01)	(0.01)	(0.01)	(0.01)			(0.01)	(0.01)	(0.00)	(0.00)
Imports over GDP (-1)				0.01*	0.01	0.01*				0.00**	0.00**	0.00***
				(0.00)	(0.00)	(0.00)				(0.00)	(0.00)	(0.00)
Investment Share					0.00	0.00					0.01**	0.00
					(0.01)	(0.01)					(0.01)	(0.01)
Wage Share						0.00						0.01**
						(0.00)						(0.00)
Constant	-0.40***	-0.07	-0.02	-0.02	-0.07	-0.14	-0.25***	-0.13***	-0.13***	-0.12***	-0.30***	-0.60***
	(0.03)	(0.12)	(0.12)	(0.12)	(0.29)	(0.44)	(0.01)	(0.05)	(0.05)	(0.05)	(0.13)	(0.19)
Trend	-0.01***	-0.00***	-0.00**	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***	-0.00***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	112	112	112	112	112	112	112	112	112	112	112	112
Adjusted R^2	0.97	0.98	0.98	0.98	0.98	0.98	0.98	0.98	0.98	0.98	0.98	0.98

[a] Bayesian regressions allow for heteroskedastic errors following Geweke (1993). Dependent variables are posterior means of the posterior distribution of impulse responses from the identified TVP-VAR. All regressions are estimated using a Gibbs sampling algorithm with 1,100 draws and 100 omitted draws. This leaves us with 1,000 posterior draws of regression coefficients.

[b] Multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t).

[c] Point estimates are posterior means of the posterior distribution. Standard deviations are reported in parentheses. Asterisks indicate posterior probabilities that the regression coefficients are non-positive if the point estimates are positive or non-negative if the point estimates are negative (*less than 10%, **less than 5%, ***less than 1%.)

[d] Explanatory variables are measured in percent.

Table 3: Bayesian linear regressions with standard errors adjusted for uncertainty in dependent variables, dependent variables are contemporaneous multipliers.[a,b,c,d]

	Multiplier on Output						Multiplier on Consumption					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Gov. Debt over GDP (-1)	0.01	-0.01	-0.01	-0.01	-0.00	-0.00	0.00	-0.01***	-0.01***	-0.01***	-0.00*	-0.00
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)
Credit over GDP (-1)		-0.06***	-0.06***	-0.04***	-0.04**	-0.06***		-0.04***	-0.03***	-0.02**	-0.02**	-0.04***
		(0.02)	(0.02)	(0.02)	(0.02)	(0.02)		(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Output Gap (-1)			0.03**	0.05***	0.04**	0.03***			0.02**	0.02***	0.03**	0.02**
			(0.02)	(0.02)	(0.02)	(0.02)			(0.01)	(0.01)	(0.02)	(0.01)
Imports over GDP (-1)				-0.02*	-0.01*	-0.01				-0.01	-0.01	-0.01
				(0.01)	(0.01)	(0.01)				(0.01)	(0.01)	(0.01)
Investment Share					0.03*	0.07**					0.01	0.04**
					(0.02)	(0.04)					(0.02)	(0.02)
Wage Share						-0.04**						-0.03**
						(0.03)						(0.02)
Constant	0.71	3.11***	2.99***	2.96***	2.11**	4.52**	0.78	2.09***	2.00***	1.98***	1.71***	3.48***
	(1.37)	(1.10)	(1.09)	(1.07)	(1.23)	(2.29)	(0.94)	(0.78)	(0.78)	(0.76)	(1.02)	(2.01)
Trend	-0.01***	0.01***	0.01***	0.01***	0.01**	0.01***	-0.01**	0.01***	0.01***	0.00**	0.00**	0.01***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	112	112	112	112	112	112	112	112	112	112	112	112

[a] Bayesian regressions allow for heteroskedastic errors following Geweke (1993). Standard error adjustment proceeds by using each of 1,000 multipliers in the posterior distribution from the identified TVP-VAR as dependent variable. All regressions are then estimated using a Gibbs sampling algorithm with 1,100 draws and 100 omitted draws. This leaves us with 1,000,000 posterior draws of regression coefficients.

[b] Multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t).

[c] Point estimates are posterior means of the posterior distribution. Standard deviations are reported in parentheses. Asterisks indicate posterior probabilities that the regression coefficients are non-positive if the point estimates are positive or non-negative if the point estimates are negative (*less than 10%, **less than 5%, ***less than 1%.)

[e] Explanatory variables are measured in percent.

Table 4: Bayesian linear regressions with standard errors adjusted for uncertainty in dependent variables, dependent variables are multipliers after five years.[a,b,c,d]

	Multiplier on Output						Multiplier on Consumption					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Gov. Debt over GDP (-1)	-0.00	-0.01**	-0.01**	-0.01**	-0.01**	-0.01**	-0.00	-0.01**	-0.01**	-0.01**	-0.01**	-0.01**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)	(0.03)	(0.03)
Credit over GDP (-1)		-0.02	-0.02	-0.03	-0.03	-0.02		-0.01	-0.01	-0.02	-0.02	-0.02
		(0.09)	(0.09)	(0.10)	(0.09)	(0.08)		(0.06)	(0.06)	(0.08)	(0.08)	(0.07)
Output Gap (-1)			-0.01	-0.02	-0.02	-0.01			-0.01	-0.02	-0.02	-0.02
			(0.10)	(0.11)	(0.11)	(0.12)			(0.11)	(0.11)	(0.08)	(0.09)
Imports over GDP (-1)				0.01	0.01	0.01				0.01	0.01	0.01
				(0.03)	(0.04)	(0.04)				(0.06)	(0.05)	(0.05)
Investment Share					-0.00	-0.01					0.00	-0.01
					(0.13)	(0.12)					(0.19)	(0.14)
Wage Share						0.01						0.01
						(0.08)						(0.09)
Constant	-0.67	0.02	0.03	0.05	0.08	-0.36	-0.56	-0.10	-0.09	-0.07	-0.10	-0.82
	(1.96)	(3.24)	(2.86)	(2.87)	(2.92)	(4.50)	(1.69)	(1.76)	(1.60)	(1.54)	(3.70)	(7.34)
Trend	-0.01***	-0.01***	-0.01**	-0.01**	-0.01**	-0.01**	-0.01***	-0.01***	-0.01**	-0.01**	-0.01**	-0.01**
	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)
Observations	112	112	112	112	112	112	112	112	112	112	112	112

[a] Bayesian regressions allow for heteroskedastic errors following Geweke (1993). Standard error adjustment proceeds by using each of 1,000 multipliers in the posterior distribution from the identified TVP-VAR as dependent variable. All regressions are then estimated using a Gibbs sampling algorithm with 1,100 draws and 100 omitted draws. This leaves us with 1,000,000 posterior draws of regression coefficients.

[b] Multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t).

[c] Point estimates are posterior means of the posterior distribution. Standard deviations are reported in parentheses. Asterisks indicate posterior probabilities that the regression coefficients are non-positive if the point estimates are positive or non-negative if the point estimates are negative (*less than 10%, **less than 5%, ***less than 1%.)

[d] Explanatory variables are measured in percent.

Table 5: Raftery and Lewis (1992) diagnostics.[a,b]

	Estim. Parameters	Thinning Factor	Burn-in Replic.	Total Runs
Coefficients	4068	1	2	150
Covariances	452	1	10	429
Volatilities	678	1	4	208

[a] Parameters for Raftery and Lewis (1992) diagnostics are quantile = 0.025; desired accuracy = 0.025; required probability of attaining the required accuracy = 0.95.

[b] Results are based on 5,000 iterations of the Gibbs sampler with zero burn-in replications and thinning factor 1.

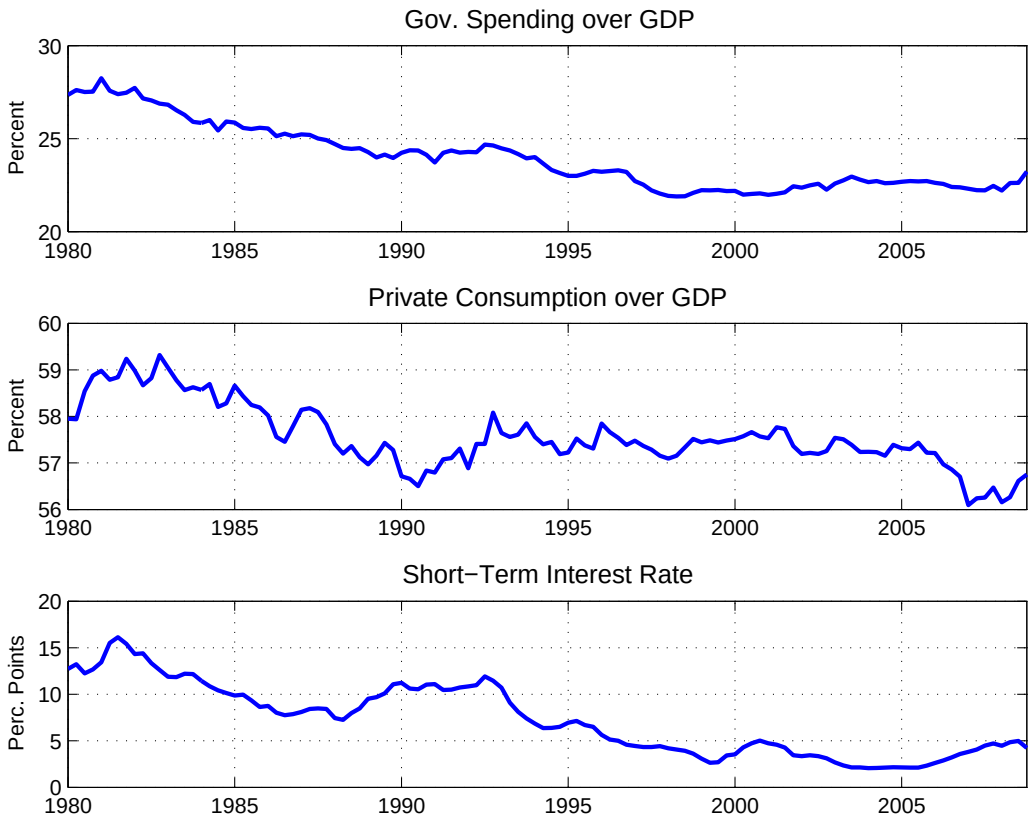

Figure 1: Data used in the baseline VAR. *Notes.* Government spending is defined as final general government consumption spending plus government investment; government spending and private consumption are expressed as nominal shares of GDP; the short-term nominal interest rate is measured in annual terms; source of fiscal data: Paredes, Pérez, and Pedregal (2009); source of remaining data: ECB's Area-Wide Model database.

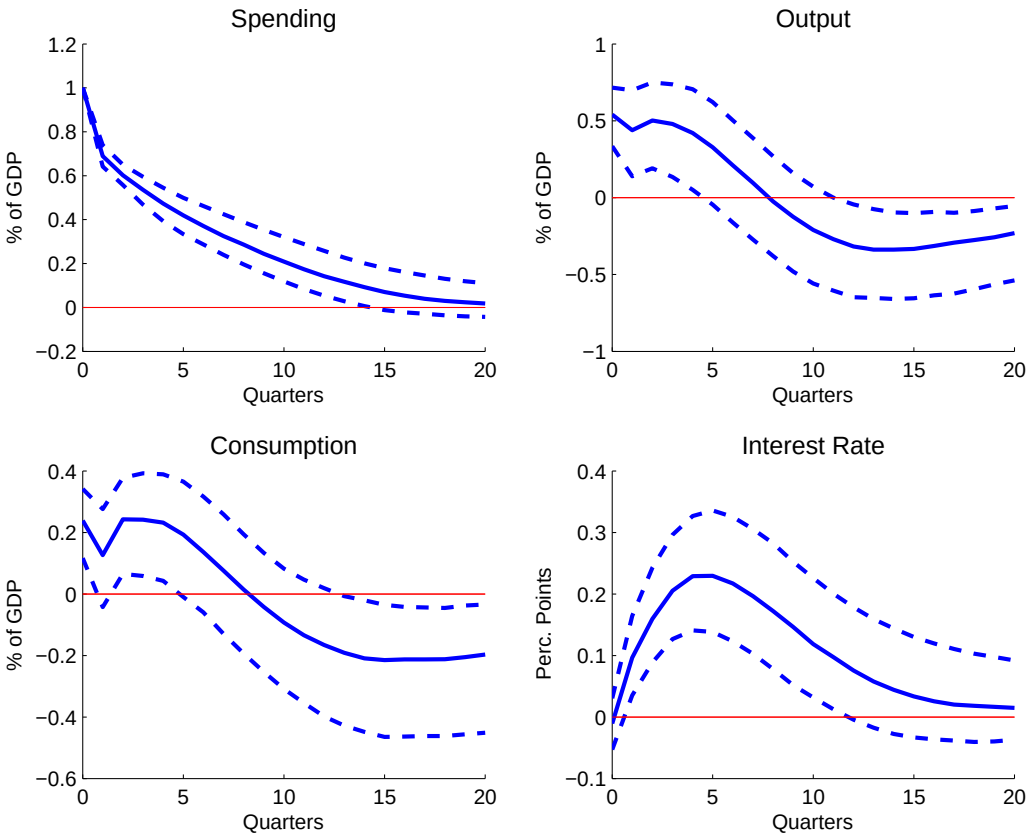

Figure 2: Impulse responses to a spending shock, baseline time invariant BVAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the responses of output, consumption and spending are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at horizon k = responding variable's response at horizon k/(spending response at horizon 0 $\times$ average ratio of spending to responding variable over sample); the response of the interest rate is reported in percentage points.

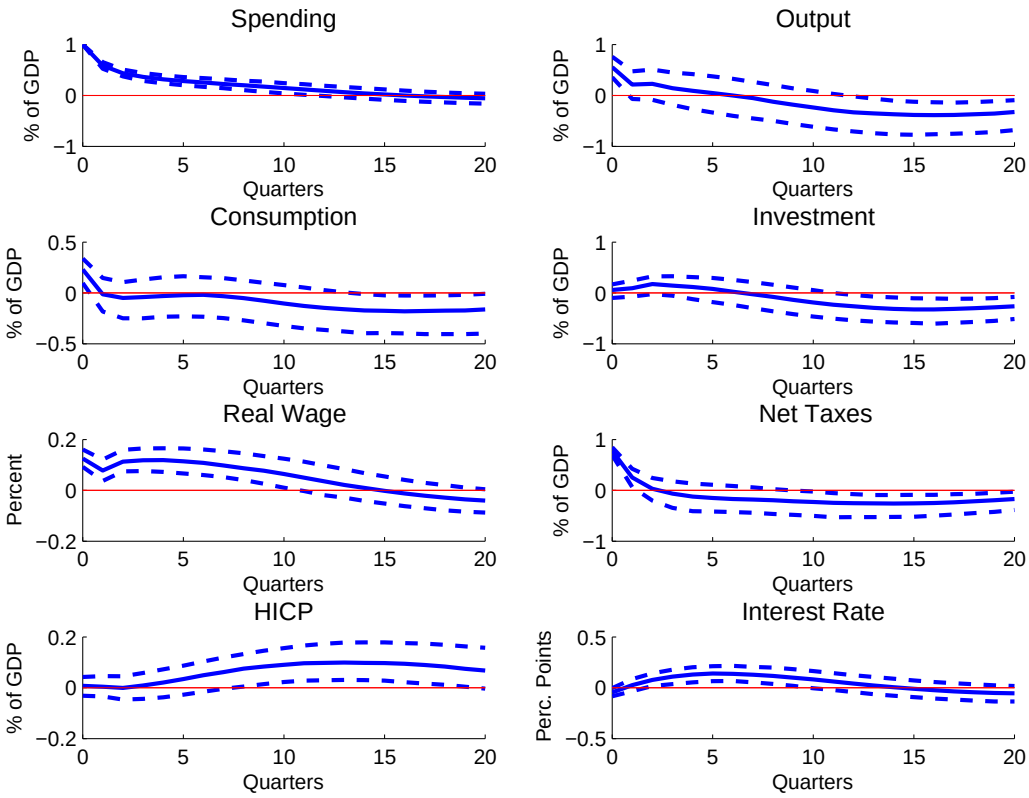

Figure 3: Impulse responses to a spending shock in, extended time invariant BVAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the responses of output, consumption, investment, net taxes and spending are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at horizon k = responding variable's response at horizon k/(spending response at horizon 0 $\times$ average ratio of spending to responding variable over sample); the response of the real wage and the HICP is measured in %; the response of the interest rate is reported in percentage points.

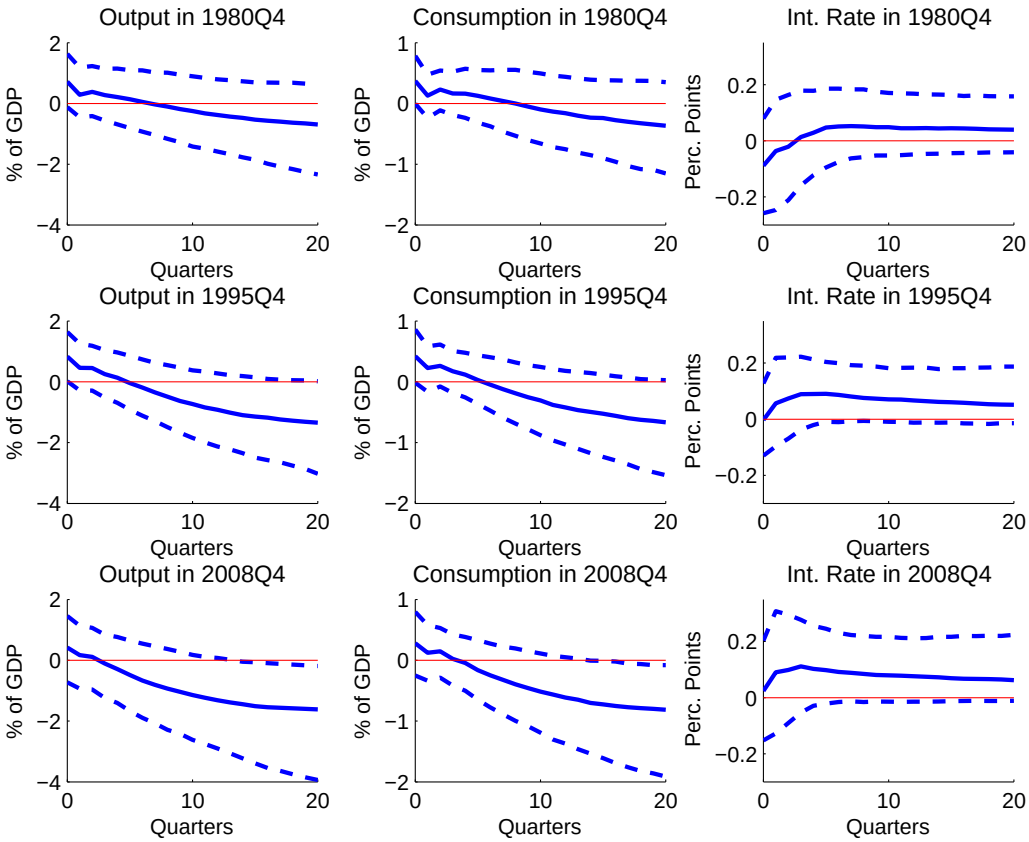

Figure 4: Impulse responses to a spending shock in selected quarters, baseline TVP-VAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the responses of output and consumption are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t); the response of the interest rate is reported in percentage points.

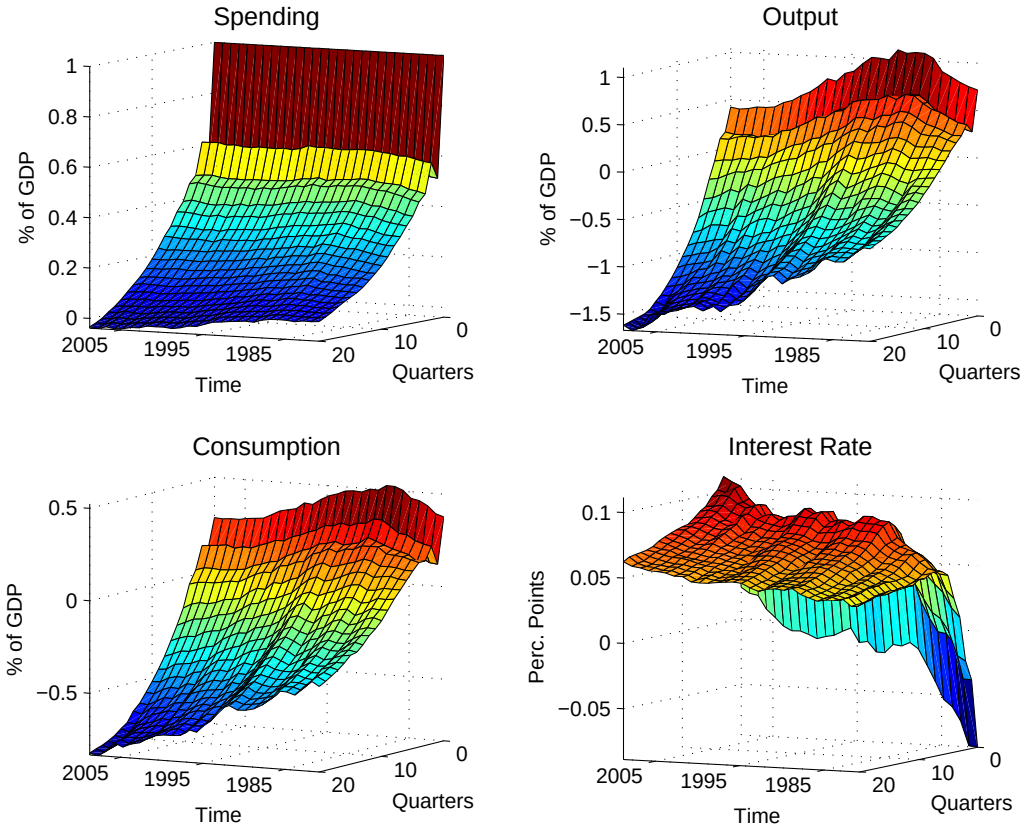

Figure 5: Impulse responses to a spending shock in each year of the sample, baseline TVP-VAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the responses of output, consumption and spending are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t); the response of the interest rate is reported in percentage points.

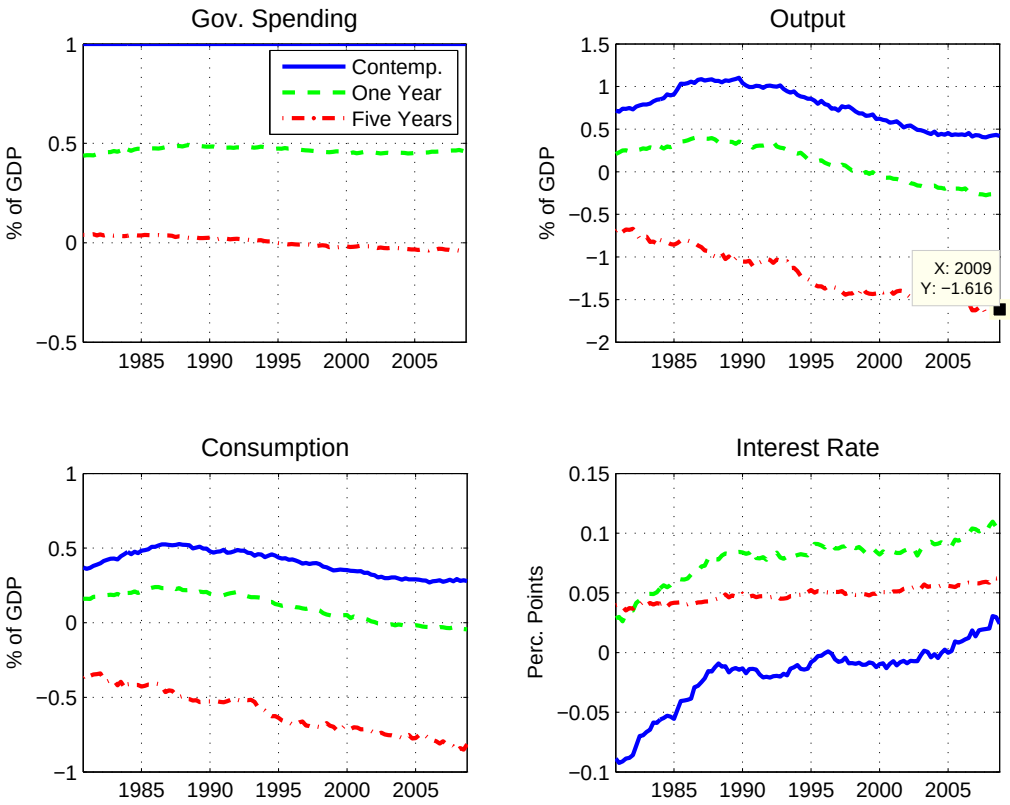

Figure 6: Impulse responses to a spending shock at selected horizons, baseline TVP-VAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the responses of output, consumption and spending are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t); the response of the interest rate is reported in percentage points.

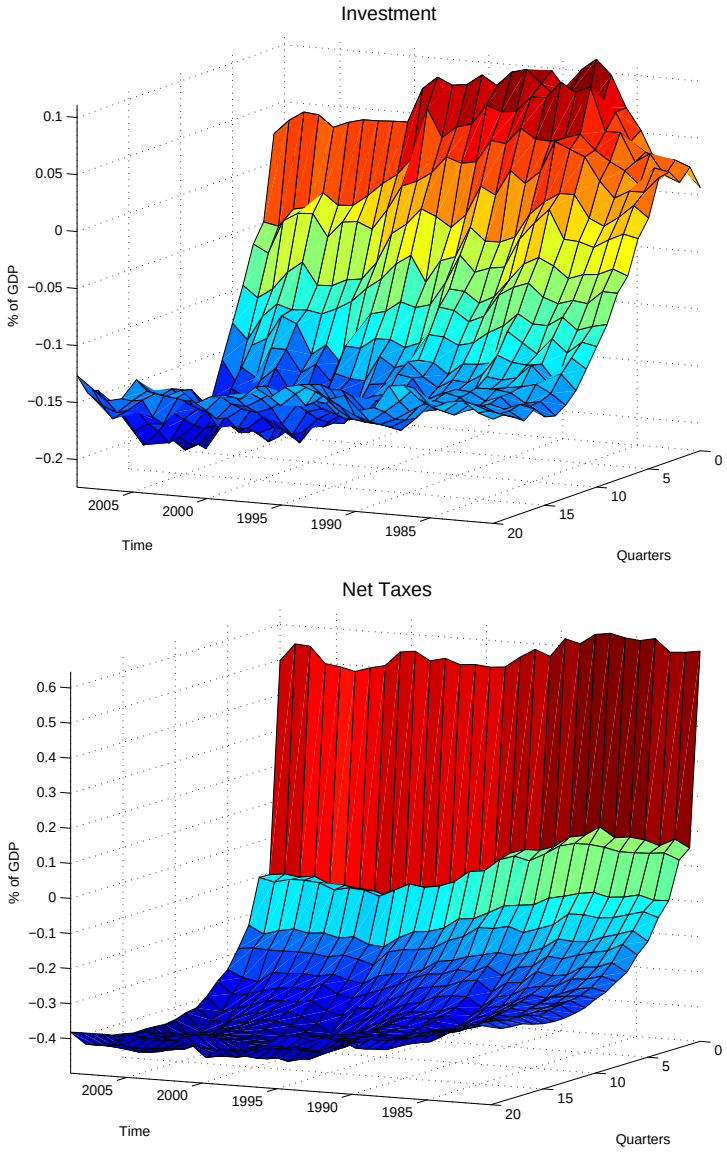

Figure 7: Impulse responses of private investment and net taxes to a spending shock in each year of the sample, extended TVP-VAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the impulse responses are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon 0 $\times$ ratio of spending to responding variable at time t).

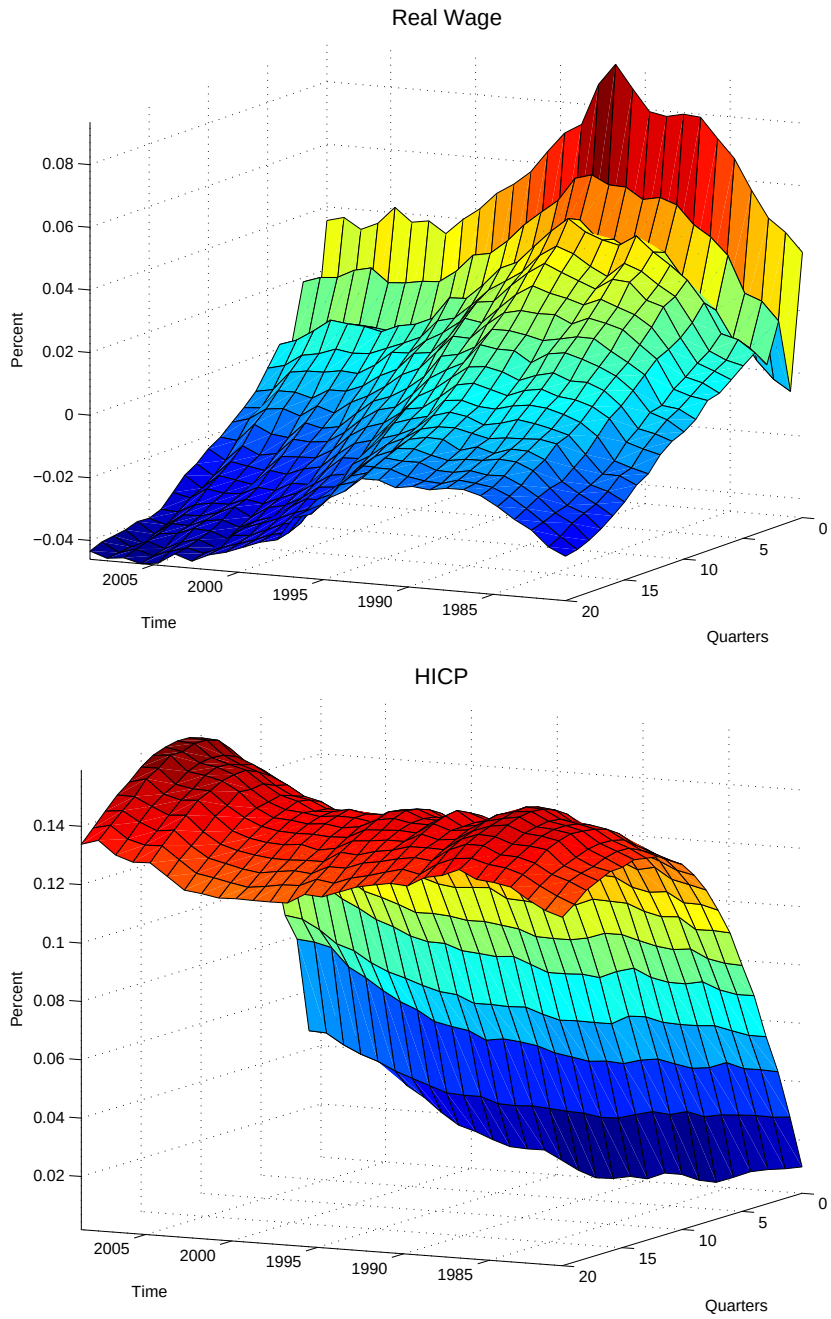

Figure 8: Impulse responses of the real wage and the HICP to a spending shock in each year of the sample, extended TVP-VAR. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the impulse responses are measured in %.

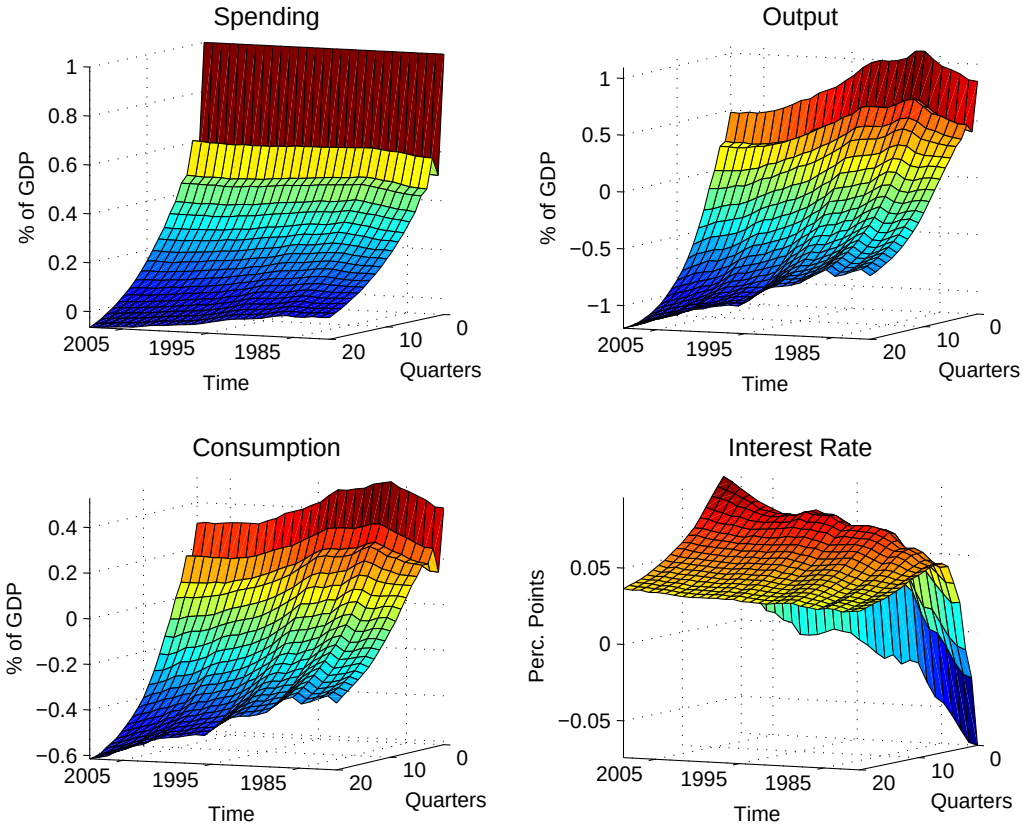

Figure 9: Impulse responses to a spending shock in each year of the sample, baseline TVP-VAR with stationarity conditions imposed. *Notes.* Median accumulated impulse responses are reported with 16% and 84% probability bands; spending increase is normalized to have size 1% of real GDP; the responses of output, consumption and spending are measured in % of GDP and have the interpretation of multipliers, i.e. responses in % of real GDP due to spending increase of size 1% of real GDP; they are computed according to the following formula: multiplier at time t and horizon k = responding variable's response at time t and horizon k/(spending response at time t and horizon $0 \times$ ratio of spending to responding variable at time t); the response of the interest rate is reported in percentage points.

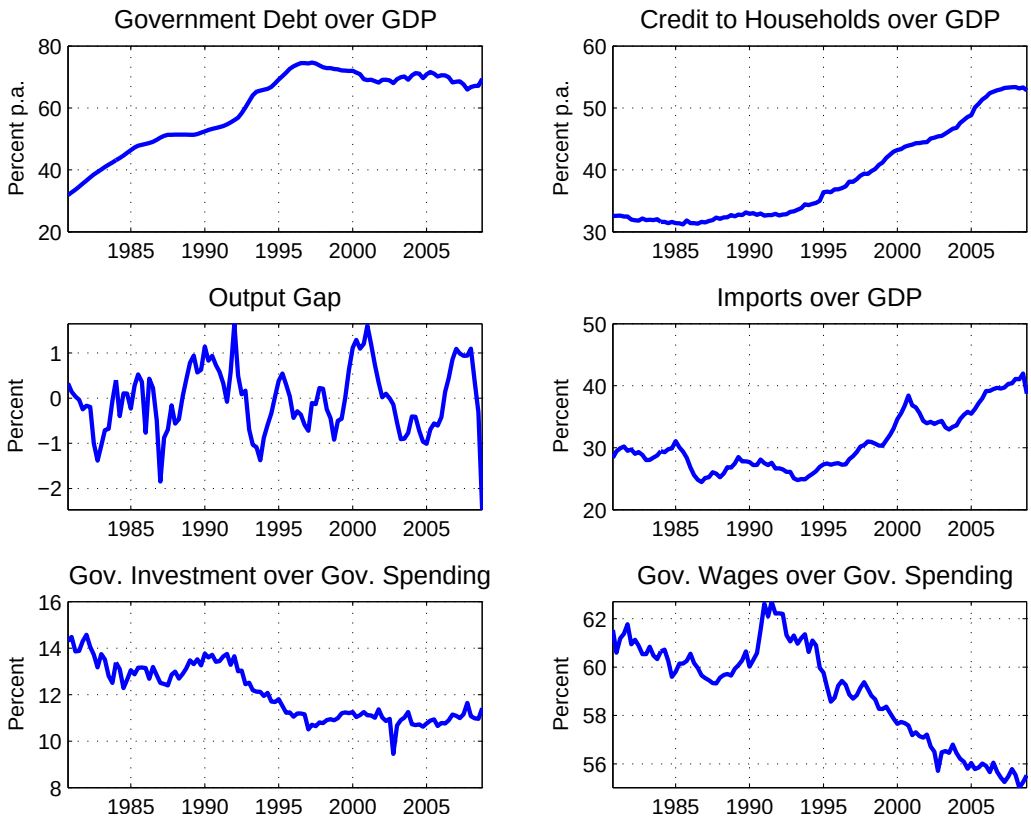

Figure 10: Potential determinants of spending multipliers. *Notes.* Debt-to-GDP ratio is in nominal annual terms; ratio of credit to households over GDP is outstanding (end-of-period) loans to households divided by the sum of nominal GDP of the last four consecutive quarters; output gap $outgap_t$ is measured as quarterly percentage deviation from trend real GDP, trend is based on HP-filter with smoothing parameter 1600; ratio of imports over GDP and shares of government investment and wage expenditures in total spending are based on quarterly nominal data; source of fiscal data: Paredes, Pérez, and Pedregal (2009); source of remaining data: ECB's Area-Wide Model database and Bank of International Settlements macroeconomic series (data on loans).

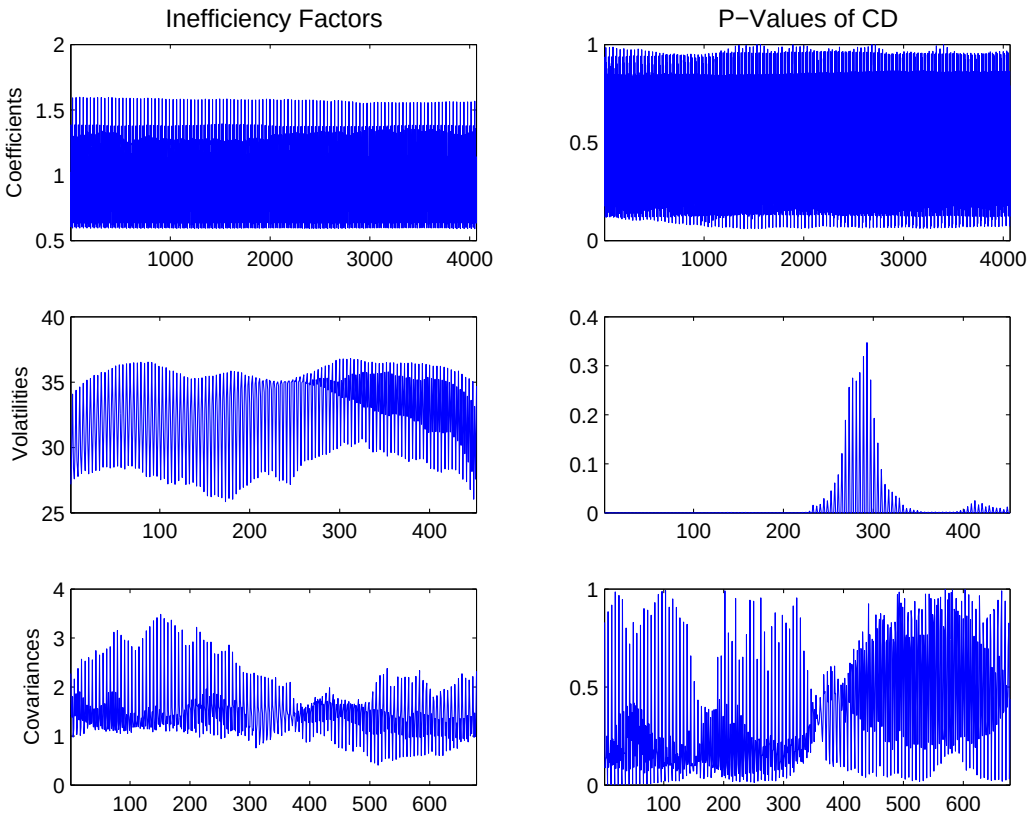

Figure 11: Convergence diagnostics for state vectors. *Notes.* Horizontal axes refer to vectors of time-varying parameters with one point representing one parameter at a given time (e.g. volatilities $h_{i,t}$); left panels: inefficiency factors, i.e. inverse of Geweke's (1992) relative numerical efficiency measure; computed as $IF = 1 + 2\sum_{k=1}^{\infty} \rho_k$, where ρ_k is the k-th order autocorrelation of the Markov chain; right panels: P-values of Geweke's (1992) convergence diagnostic; computed as $CD = (\widehat{\theta}_1 - \widehat{\theta}_2)/(\widehat{\sigma}_1/\sqrt{N_1} + \widehat{\sigma}_2/\sqrt{N_2}) \rightarrow \mathrm{N}(0,1)$, where $N_1 = 100$, $N_2 = 400$, middle 500 draws dropped.

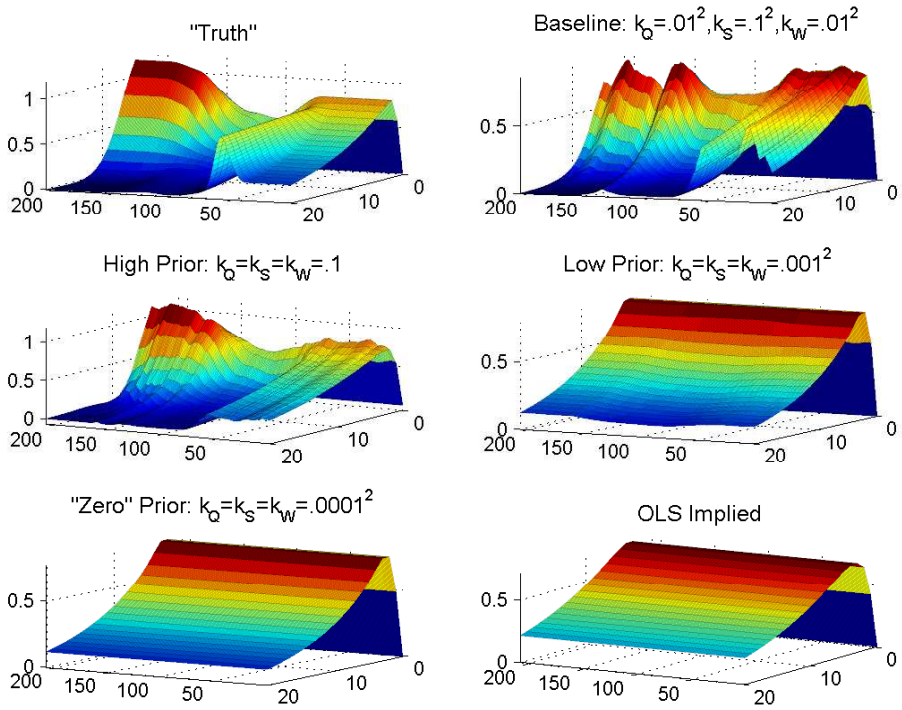

Figure 12: Impulse responses from Monte Carlo exercise, Q unrestricted and S block diagonal. *Notes.* Impulse responses are estimated based on a simulated bivariate data set with 1050 observations; every 5th reduced-form response of the second variable to an innovation in the first variable is plotted; left axes: observations/time; right axes: horizon of impulse response.

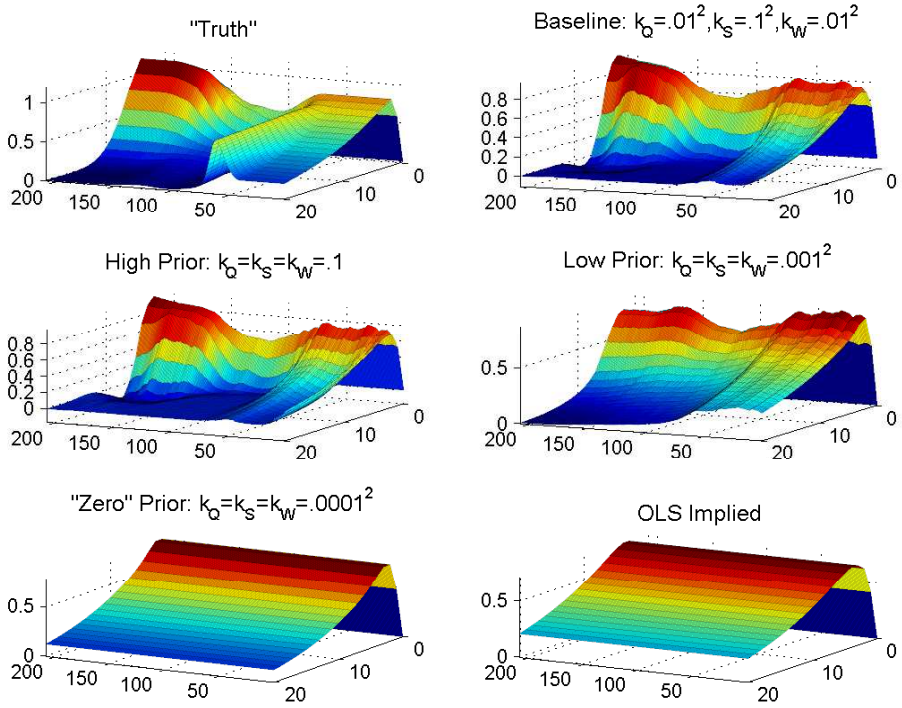

Figure 13: Impulse responses from Monte Carlo exercise, Q diagonal and S block diagonal. *Notes.* Impulse responses are estimated based on a simulated bivariate data set with 1050 observations; every 5th reduced-form response of the second variable to an innovation in the first variable is plotted; left axes: observations/time; right axes: horizon of impulse response.

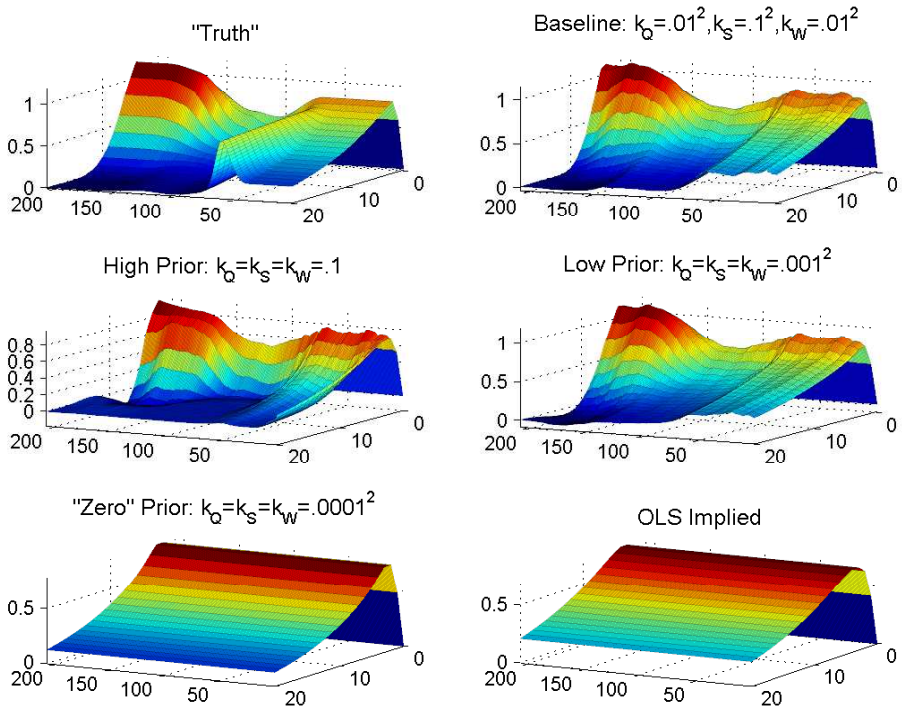

Figure 14: Impulse responses from Monte Carlo exercise, Q and S diagonal. *Notes.* Impulse responses are estimated based on a simulated bivariate data set with 1050 observations; every 5th reduced-form response of the second variable to an innovation in the first variable is plotted; left axes: observations/time; right axes: horizon of impulse response.