How do Fiscal and Technology Shocks affect Real Exchange Rates? New Evidence for the United States

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Abstract

Using vector autoregressions on U.S. time series relative to an aggregate of industrialized countries, this paper provides new evidence on the dynamic effects of government spending and technology shocks on the real exchange rate and the terms of trade. To achieve identification, we derive robust restrictions on the sign of several impulse responses from a two-country general equilibrium model. We find that both the real exchange rate and the terms of trade—whose responses are left unrestricted—depreciate in response to expansionary government spending shocks and appreciate in response to positive technology shocks.

Keywords: Real exchange rate, terms of trade, international transmission mechanism, government spending shocks, technology shocks, VAR, sign restrictions

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1 Introduction

How do international relative prices adjust to country-specific fiscal measures and productivity gains? This question is pivotal to understanding the international transmission mechanism; and yet, theoretical and empirical approaches tend to provide conflicting answers. Business cycle models under conventional calibrations predict that government spending raises the relative price of domestic goods, while productivity gains lower it—reflecting, respectively, an increase in relative demand and supply of domestic goods.[1] Recent empirical studies based on estimated vector autoregressive (VAR) models suggest the opposite. Kim and Roubini (2008), Monacelli and Perotti (2006), and Ravn, Schmitt-Grohé and Uribe (2007), among others, find that government spending depreciates the real exchange rate. Corsetti, Dedola and Leduc (2008b), Kim and Lee (2008), and Enders and Müller (2009) document that productivity gains (or 'technology shocks') appreciate real exchange rates, measured by the terms of trade or the relative price of consumption across countries.[2]

In order to reassess these puzzling findings, the present paper employs a new methodological approach: while existing studies identify exogenous structural innovations through either short-run or long-run restrictions, we follow Uhlig (2005) and restrict the sign of the responses to the shocks we seek to identify.[3] Importantly, and in contrast to a closely related study by Corsetti, Dedola and Leduc (2009), we use a quantitative general equilibrium model to formally derive the sign and the time horizon of the identification restrictions. Our model is richly specified and nests distinct transmission mechanisms, once we consider the entire range of plausible parameterizations. Specifically, while the model delivers robust predictions for the behavior of several macroeconomic variables, it does not yield clear-cut predictions for how exchange rates respond to government spending and technology shocks. This result is key to our identification strategy: we derive sign restrictions for several variables from the model, while remaining agnostic about exchange rate dynamics.[4]

[1]See, e.g., Backus, Kehoe and Kydland (1994) and Erceg, Guerrieri and Gust (2005). Assuming debt-finance, textbook versions of the Mundell-Fleming model also predict that an exogenous increase in government spending appreciates exchange rates. In the case of tax finance, results differ as disposable income and money demand fall if money supply is unchanged, see Frenkel and Razin (1987). For similar reasons, government spending depreciates the nominal exchange rate in Obstfeld and Rogoff (1995).

[2]The aforementioned studies on fiscal shocks focus on the real exchange rate and consider data for Australia, Canada, the U.K. and the U.S. Evidence on the effect of government spending shocks on the terms of trade is somewhat mixed, see, for instance, Corsetti and Müller (2006) or Monacelli and Perotti (2008). Regarding technology shocks, evidence for an appreciation is established for U.S. data. Corsetti et al. (2008b) find an appreciation in Japan as well, while Kim and Lee (2008) report a depreciation for the Euro area and Japan.

[3]An increasing number of studies has recently employed sign restrictions. They are used, for instance, to identify government spending and technology shocks in a closed economy context by Mountford and Uhlig (2009) and Peersman and Straub (2009). In an open economy context, with a focus on identifying monetary policy shocks, sign restrictions are employed by Faust and Rogers (2003), Farrant and Peersman (2006), and Scholl and Uhlig (2008) among others.

[4]The validity of our identification assumptions thus rest on the plausibility of our theoretical framework. Yet working with a fully specified general equilibrium model imposes discipline on how to specify sign restrictions. By the same token, it allows for a quantification of the time horizon for which restrictions may be imposed as well as for an explicit treatment of a possible anticipation of government spending shocks. We therefore consider our study complementary to Corsetti, Dedola and Leduc (2009) who employ sign restrictions to identify demand and technology shocks in the manufacturing sector and study their effect on the real exchange rate. Rather than using a fully specified general equilibrium model, they

We establish new evidence on how productivity gains and government spending impact U.S. exchange rates. In contrast to existing studies, which analyze the effect of either government spending or productivity gains in isolation, we assess their effects jointly in order to establish encompassing evidence on the international transmission mechanism. Specifically, we estimate our VAR model on quarterly times series for the U.S. relative to an aggregate of industrialized countries for the post-Bretton-Woods period 1975Q1–2005Q4. The VAR includes data for consumption, output, investment, government spending, the government budget balance, inflation, the short-term interest rate and exchange rates. As a measure for the latter, we consider both the real effective exchange rate and the terms of trade, in order to control for the possibility that exchange rate changes merely reflect fluctuations in the price of non-traded goods.

We find that exogenous expansions of government spending depreciate the real exchange rate as well as the terms of trade. Positive innovations to technology, instead, appreciate the real exchange rate and the terms of trade in the short run. While the terms of trade converge back to their initial value, the real exchange rate depreciates in the medium run after a positive technology shock. Sensitivity analysis, accounting for various complications such as possible anticipation effects of government spending, monetary policy shocks, or variations in the sample period, shows that these results are robust.

Overall, our results corroborate the findings of existing studies regarding the effects of government spending and technology shocks on exchange rates, even though we employ an identification scheme which is conceptually quite distinct. Identification assumptions are, by their very nature, controversial, and evidence on exchange rate dynamics which is robust across identification schemes seems particular relevant in assessing conflicting theoretical accounts of the international transmission mechanism. Specifically, international relative prices play an important role in allocating country-specific risk in the absence of explicit risk-sharing. Cole and Obstfeld (1991) identify conditions under which international price movements fully insure country-specific risk, thereby supporting the efficient allocation. Yet, as shown in a recent theoretical contribution by Corsetti, Dedola and Leduc (2008a), to the extent that technology shocks appreciate the real exchange rate, country-specific consumption risk is actually amplified. The reverse holds for government spending shocks. Our empirical findings are thus consistent with the notion that, in the short run, international price movements tend to amplify rather than to mitigate country-specific consumption risk.

The remainder of the paper is organized as follows. In section 2 we describe our identification strategy and outline a quantitative business cycle model from which we derive sign restrictions. In section 3 we discuss our VAR specification and results. Section 4 discusses sensitivity and section 5 concludes.

use sector-specific information to achieve identification.

2 Identifying government spending and technology shocks

2.1 Sign restrictions

As discussed above, several studies use VAR models to document the effects of government spending and technology shocks on exchange rates. In these studies identification is based on either short-run or long-run restrictions. In the following we propose to bring an alternative identification scheme to bear on the question, because the evidence established to date conflicts with the predictions of business cycle models, at least if standard calibrations are considered. In the following we briefly outline our approach. We start from the following reduced-form VAR model

$$Y_t = \mu + B_{(1)} Y_{t-1} + B_{(2)} Y_{t-2} + ... + B_{(m)} Y_{t-m} + u_t, \quad E[u_t u_t'] = \Sigma, \tag{1}$$

$t = 1, ..., T$, for some ℓ-dimensional vector of variables Y_t, coefficient matrices $B_{(i)}$ of size $\ell \times \ell$ and a variance-covariance matrix for the one-step ahead prediction error Σ. Letting υ_t, with $E[\upsilon_t \upsilon_t'] = I_\ell$, denote the vector of structural shocks, we need to find a matrix A such that $u_t = A\upsilon_t$ in order to achieve identification.

Instead of restricting the matrix A a priori, we follow Uhlig (2005) and Mountford and Uhlig (2009) and identify structural shocks by imposing sign restrictions on impulse-response functions of selected variables for a certain period $k = \underline{k}, ..., \overline{k}$ following the shock. Intuitively, we consider various matrices A and check, for each case, whether the sign restrictions are fulfilled and dismiss the matrix if this is not the case. Below, we derive the sign restrictions on the basis of a quantitative business cycle model. Specifically, we assess—for a wide range of model parameterizations—whether the response of a variable to a particular shock is either robustly negative or positive for a specific time period k after the shock impacts the model economy.

To fix ideas, let n be the number of structural shocks that we seek to identify. Mountford and Uhlig (2009) show that identifying n shocks is equivalent to identifying an impulse matrix of rank n that is a sub-matrix of matrix A satisfying $AA' = \Sigma$. Any impulse matrix can be written as

$$[a^{(1)}, ..., a^{(n)}] = \tilde{A} Q \tag{2}$$

where $\tilde{A}$ is the lower triangular Cholesky factor of Σ and $Q = [q^{(1)}, ..., q^{(n)}]$ is a $n \times \ell$ matrix consisting of orthonormal rows $q^{(s)}$, $s = 1, ..., n$, such that $QQ' = I_n$.

Similarly to Uhlig (2005), one can show that the impulse response to $a^{(s)}$ can be written as linear combination of the impulse responses obtained under a Cholesky decomposition of Σ. Let $c_{ji}(k)$ be the impulse response of the jth variable at horizon k to the ith shock in the Cholesky decomposition of Σ and define $c_i(k) \in \mathbb{R}^\ell$ to be the vector response $[c_{1i}(k), ..., c_{\ell i}(k)]$. Then the impulse response

$r_a^{(s)}(k)$ to the impulse vector $a^{(s)}$ is given by

$$r_a^{(s)}(k) = \sum_{i=1}^{\ell} q_i^{(s)}\ c_i(k). \tag{3}$$

The restrictions we impose to identify an impulse vector characterizing shock s are that $(r_a^{(s)}(k))_j \geq 0, j \in \mathcal{J}_+$ and $(r_a^{(s)}(k))_j \leq 0, j \in \mathcal{J}_-$ for some subsets of variables $\mathcal{J}_+$ and $\mathcal{J}_-$ and some horizon $k = \underline{k}, \ldots, \overline{k}$.

For the actual estimation we employ a Bayesian approach. Specifically, we use a flat Normal-Wishart prior (see Uhlig (1994) for a detailed discussion of the properties), while the numerical implementation follows Rubio-Ramirez, Waggoner and Zha (2005) and can be summarized as follows. We take a draw from the Normal-Wishart posterior for (B, Σ) and construct an arbitrary independent standard normal matrix M. We obtain the orthonormal matrix Q using the QR-decomposition of M such that $QQ' = I$ and $QR = M$. We construct impulse vectors a according to (2) and use (3) to compute the impulse responses.

Considering orthogonal structural shocks may result in tight identifying sign restrictions in the sense that many draws from the Normal-Wishart posterior for the VAR parameters (B, Σ) are rejected because they do not permit any impulse matrices that satisfy the sign restrictions. This means that many draws receive zero prior weight, even in cases where only few of the restrictions are mildly violated. This issue gets more severe if the number of orthogonal shocks and the number of variables included in the VAR model increases. To account for this complication, we allow for small deviations ε from the sign restrictions and define

$$(\omega_a^{(s)}(k))_j = \begin{cases} \max\{-(r_a^{(s)}(k))_j, 0\} \text{ for } j \in \mathcal{J}_+,\ k = \underline{k}, ..., \overline{k} \text{ and } s = 1, .., n, \\ \max\{(r_a^{(s)}(k))_j, 0\} \quad \text{for } j \in \mathcal{J}_-,\ k = \underline{k}, ..., \overline{k} \text{ and } s = 1, .., n. \end{cases}$$

We keep the impulse responses if the sum of the squared deviations over all structural shocks, variables and horizons is smaller than ε:

$$\sum_s \sum_j \sum_k \left[(\omega_a^{(s)}(k))_j \right]^2 < \varepsilon, \qquad \varepsilon \geq 0. \tag{4}$$

Inference statements are based on the posterior distribution of those draws for which (4) is satisfied.[5]

2.2 A quantitative business cycle model

We now turn to a quantitative business cycle model from which we derive sign restrictions. The model is a two-country business cycle model featuring various frictions frequently employed in earlier

[5]Alternatively, Mountford and Uhlig (2009) minimize a penalty function for sign restriction violations for each draw from the posterior distribution for the VAR parameters. However, to account for several orthogonal shocks they sequentially determine the optimal impulse vectors such that the ordering of the structural shocks may be important. To avoid this, we simply allow for small deviations and draw the impulse vectors simultaneously. This also implies that, in contrast to Mountford and Uhlig (2009), we simultaneously estimate the reduced-form parameters together with the impulse matrix.

studies, see, e.g., Chari, Kehoe and McGrattan (2002) and Kollmann (2002). Notably, we consider various degrees of price rigidity, since sticky prices potentially alter the transmission of real shocks, as forcefully argued by Galí (1999). In addition, this assumption allows us to study the behavior of international relative prices also in response to monetary policy shocks, once we turn to sensitivity analysis. Moreover, we assume that each country specializes in the production of a particular type of goods. Households in each country consume both types, but to a different extent, such that changes in their relative price govern real exchange rate dynamics. We abstract from non-traded goods, as fluctuations in their relative price are of minor importance in accounting for U.S. real exchange rate changes, see Engel (1999) and Chari et al. (2002).

Before we turn to model simulations in order to derive sign restrictions, we briefly outline the model structure. The world economy consists of two symmetric countries indexed by $i \in \{1, 2\}$. We refer to country 1 as the domestic economy or 'Home', and to country 2 as 'Foreign'.

Households In country i, a representative household allocates resources to consumption goods, C_{it}, and supplies labor, H_{it}, to monopolistic firms. Preferences are given by

$$\begin{aligned} &E_0 \sum_{t=0}^{\infty} \beta_{it} \frac{[C_{it}^{\mu}(1-H_{it})^{1-\mu}]^{1-\gamma}}{1-\gamma}, && \mu < 1, \\ &\beta_{i0} = 1, \quad \beta_{it+1} = \left(1 + \psi[C_{it}^{\mu}(1-H_{it})^{1-\mu}]\right)^{-1} \beta_{it}, && t \geq 0. \end{aligned} \tag{5}$$

Here β_{it} is an endogenous discount factor implying higher discounting if consumption and leisure are above their steady-state values.[6] The positive constants μ and γ specify the preferences of households. Labor and capital are internationally immobile; households in country i own the capital stock, K_{it}, and rent it to intermediate good firms on a period-by-period basis. It may be costly to adjust the level of investment, I_{it}. As in Christiano, Eichenbaum and Evans (2005), the law of motion for capital is given by

$$K_{it+1} = (1-\delta)K_{it} + [1 - \Psi(I_{it}/I_{it-1})]I_{it}, \tag{6}$$

where δ denotes the depreciation rate; restricting $\Psi(1) = \Psi'(1) = 0$, and $\Psi''(1) = \chi > 0$ ensures that the steady-state capital stock is independent of investment adjustment costs captured by the parameter χ. Across countries there is trade in nominal non-contingent bonds, Θ_{it}, denominated in the currency of country i. The budget constraint of the representative household in country i reads as

$$\begin{aligned} &(1-\tau_{it})(W_{it}H_{it} + R_{it}^{K}K_{it} + \Upsilon_{it}) - P_{it}C_{it} - P_{it}I_{it} \\ = &\begin{cases} (\Theta_{1t+1} + D_{1t+1})R_{1t}^{-1} + S_t\Theta_{2t+1}R_{2t}^{-1} - \Theta_{1t} - D_{1t} - S_t\Theta_{2t}, & \text{for } i = 1, \\ (\Theta_{2t+1} + D_{2t+1})R_{2t}^{-1} + \Theta_{1t+1}(R_{1t}S_t)^{-1} - \Theta_{2t} - D_{2t} - S_t^{-1}\Theta_{1t}, & \text{for } i = 2, \end{cases} \end{aligned} \tag{7}$$

[6] We assume that the effect of consumption and leisure on the discount factor is not internalized by the household, see Schmitt-Grohé and Uribe (2003) for a detailed analysis. The parameter ψ determines how strongly the discount factor responds to consumption and leisure; it also pins down the value of the discount factor in steady state.

where W_{it} and R_{it}^K denote the nominal wage rate and the rental rate of capital, and Υ_{it} are nominal profits earned by monopolistic firms and transferred to households. τ_{it} denotes the tax rate levied on households' income; P_{it} is the price of the final good defined below; R_{it} is the gross nominal interest rate, D_{it} denotes debt issued by the government in country i held by domestic residents, and S_t is the nominal exchange rate. In each country households maximize (5) subject to (6), (7) and a non-Ponzi scheme condition.

Final good firms Consumption and investment goods are composite goods which households purchase from final good firms. These firms operate under perfect competition and buy intermediate goods from a continuum of monopolistic competitive firms. We use $j \in [0,1]$ to index those intermediate-good firms as well as their products and prices. Further, let $A_{it}(j)$ and $B_{it}(j)$ denote the amount of good j originally produced in country 1 and 2, respectively, and used in country i to assemble final goods F_{it}. These are produced under the following technology

$$F_{it} = \begin{cases} \left[\begin{array}{c} \omega^{\frac{1}{\sigma}} \left(\left[\int_0^1 A_{1t}(j)^{\frac{\epsilon-1}{\epsilon}} dj \right]^{\frac{\epsilon}{\epsilon-1}} \right)^{\frac{\sigma-1}{\sigma}} \\ + (1-\omega)^{\frac{1}{\sigma}} \left(\left[\int_0^1 B_{1t}(j)^{\frac{\epsilon-1}{\epsilon}} dj \right]^{\frac{\epsilon}{\epsilon-1}} \right)^{\frac{\sigma-1}{\sigma}} \end{array} \right]^{\frac{\sigma}{\sigma-1}}, & \text{for } i = 1 \\ \left[\begin{array}{c} \omega^{\frac{1}{\sigma}} \left(\left[\int_0^1 B_{2t}(j)^{\frac{\epsilon-1}{\epsilon}} dj \right]^{\frac{\epsilon}{\epsilon-1}} \right)^{\frac{\sigma-1}{\sigma}} \\ + (1-\omega)^{\frac{1}{\sigma}} \left(\left[\int_0^1 A_{2t}(j)^{\frac{\epsilon-1}{\epsilon}} dj \right]^{\frac{\epsilon}{\epsilon-1}} \right)^{\frac{\sigma-1}{\sigma}} \end{array} \right]^{\frac{\sigma}{\sigma-1}}, & \text{for } i = 2 \end{cases} \tag{8}$$

where σ denotes the elasticity of substitution between foreign and domestic goods ('trade price elasticity', for short) and ϵ measures the elasticity of substitution between goods produced within the same country. The parameter ω measures the home bias in the composition of final goods. Let $P_{it}^A(j)$ be the price in country i of an intermediate good produced in country 1 and $P_{it}^B(j)$ the price in country i of a good produced in country 2. Assuming that the law of one price holds, we have

$$P_{1t}^B(j) = S_t P_{2t}^B(j); \qquad P_{1t}^A(j) = S_t P_{2t}^A(j). \tag{9}$$

The price for final goods is given by

$$P_{it} = \begin{cases} \left[\omega \left(P_{1t}^A \right)^{1-\sigma} + (1-\omega) \left(P_{1t}^B \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}, & \text{for } i = 1 \\ \left[(1-\omega) \left(P_{2t}^A \right)^{1-\sigma} + \omega \left(P_{2t}^B \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}, & \text{for } i = 2 \end{cases}, \tag{10}$$

where

$$P_{it}^A = \left(\int_0^1 P_{it}^A(j)^{1-\epsilon} dj \right)^{\frac{1}{1-\epsilon}} \text{ and } P_{it}^B = \left(\int_0^1 P_{it}^B(j)^{1-\epsilon} dj \right)^{\frac{1}{1-\epsilon}} \tag{11}$$

denote the GDP deflators in Home and Foreign, respectively.

The problem of final good firms is to minimize expenditures in assembling intermediate goods subject to (8) and the requirement that $F_{it} = C_{it} + I_{it}$. The first-order condition that characterizes the behavior of final good firms in equilibrium implicitly defines the demand for a generic intermediate goods, $Y_{it}^D(j)$.

For future reference, taking the perspective of the home country, we define the real exchange rate as follows

$$RX_t = S_t P_{2t}/P_{1t}, \tag{12}$$

such that an increase corresponds to a depreciation. The terms of trade are defined as the price of imports relative to the price of exports: P_{1t}^B/P_{1t}^A.

Intermediate good firms At the intermediate good level, firms specialize in the production of differentiated goods. A generic firm $j \in [0, 1]$ in country i engages in monopolistic competition facing imperfectly-elastic demand from domestic and foreign final-good producers, as well as domestic governments which are assumed to consume only domestically produced goods, as discussed below. Production of intermediate goods is Cobb-Douglas:

$$Y_{it}(j) = e^{Z_{it}} K_{it}(j)^{\theta} H_{it}(j)^{1-\theta}, \tag{13}$$

where Z_{it} denotes the level of technology common to all firms. It evolves exogenously according to

$$Z_{it} = \rho_z Z_{it-1} + \rho_{zz} Z_{3-i,t-1} + \varepsilon_{it}^z, \tag{14}$$

such that ρ_z captures the degree of autocorrelation and ρ_{zz} possible spillovers across countries. Labor and capital inputs of firm j, $H_{it}(j)$ and $K_{it}(j)$, are adjusted freely in each period. Price setting, however, is constrained exogenously by a discrete time version of the mechanism suggested by Calvo (1983). In a given period, each firm has the opportunity to change its price with probability $1-\xi$ only. When a firm has the opportunity, it sets the new price in order to maximize the expected discounted value of profits; otherwise prices are indexed to past inflation, where the degree of indexation is given by $\iota \in [0, 1]$. When setting the new price $P_{1t}^A(j)$ or $P_{2t}^B(j)$, the problem of a generic intermediate-good firm j in country i is given by

$$\max \sum_{k=0}^{\infty} \xi^k E_t \begin{cases} \rho_{it,t+k} Y_{it+k}(j) \left[P_{it}^A(j) (P_{it+k-1}^A / P_{it-1}^A)^{\iota} - MC_{it+k} \right] / P_{it+k}, & \text{for } i = 1 \\ \rho_{it,t+k} Y_{it+k}(j) \left[P_{it}^B(j) (P_{it+k-1}^B / P_{it-1}^B)^{\iota} - MC_{it+k} \right] / P_{it+k}, & \text{for } i = 2 \end{cases} \tag{15}$$

subject to the production function (13) and the optimal choice of factor inputs which minimizes marginal costs, MC_{it}.[7] As households own firms, profits are discounted with $\rho_{it,t+k}$, which equals the household's marginal rate of substitution between consumption in period t and $t + k$.

[7] In this formulation we impose the constraint that demand is met by actual production at all times: $Y_{it}(j) = Y_{it}^D(j) + Y_{it}^G(j)$, where the last term denotes the demand stemming from government consumption.

Fiscal and monetary policy Government policies are characterized by feedback rules. Turning to fiscal policy first, we assume that government spending, G_{it}, consists of a bundle of intermediate goods. Specifically, we assume an aggregation technology isomorphic to (8), except that only domestically produced goods enter the consumption basket of the government.[8] Government consumption evolves according to the following feedback rule:

$$G_{it} = (1-\rho_g)G_i + \rho_g G_{it-1} + \varphi_y(Y_{it} - Y_i) - \varphi_d(D_{it} - D_i) + \varepsilon^g_{it,t-n}, \tag{16}$$

where $Y_{it} = \left(\int_0^1 Y_{it+k}(j)^{\frac{\epsilon-1}{\epsilon}} dj\right)^{\frac{\epsilon}{\epsilon-1}}$ is an index for aggregate domestic production ('output', for short); letters without time subscript refer to steady-state values; ρ_g captures persistence, while φ_y and φ_d measure to what extent government spending responds to the deviation of output and debt from their steady-state values.[9] $\varepsilon^g_{it,t-n}$ is an i.i.d. innovation to current government spending, which may have been correctly anticipated n periods in advance because, say, of institutional features of the legislative process.

Regarding the tax rate we assume that $\tau_{it} = \tau_i + \varphi_\tau(D_{it} - D_i)/Y_i$, with $\varphi_\tau \geq 0$ measuring how strongly the tax rate adjusts to the level of debt.[10] The budget constraint of the government in country i is given by

$$D_{it} + P^G_{it}G_{it} = \tau_{it}(W_{it}H_{it} + R^K_{it}K_{it} + \Upsilon_{it}) + D_{it+1}R^{-1}_{it}, \tag{17}$$

where P^G_{it} is the price index of government consumption.

Monetary policy is characterized by an interest feedback rule, whereby the policy rate is adjusted in response to domestic (i.e., producer-price) inflation, Π_{it}, and a measure of the output gap, $y_{it} = (Y_{it} - Y_i)/Y_i$ as, for instance, in Galí and Monacelli (2005):

$$R_{it} = \rho_r R_{it-1} + (1-\rho_r)\left(R + \phi_\pi(\Pi_{it} - \Pi) + 0.25\phi_y y_{it}\right) + \varepsilon^r_{it}. \tag{18}$$

Here $\rho_r \geq 0$ captures interest rate smoothing, while ϕ_π and ϕ_y measure the long-run inflation and output gap response of the policy rate; ε^r_{it} is an i.i.d. exogenous monetary policy shock.

2.3 Generating sign restrictions

We approximate the equilibrium conditions of the model around a deterministic, symmetric, zero-debt steady state and compute the model solution numerically. In order to determine parameter values we

[8]Put differently, we assume that government goods are assembled in the same way as in (8), with $\omega = 1$. The evidence discussed in Corsetti and Müller (2006) suggests that the import content in government spending is generally less than half the import content in private spending. As a first approximation it is thus reasonable to assume zero import content in government spending.

[9]Rules of this type have been estimated by Galí and Perotti (2003), among others.

[10]In the simulation of the model we only allow for values of $\{\varphi_\tau, \varphi_d\}$ such that government debt is stationary. It is interesting to observe that in the case of $\varphi_\tau = 0$ all financing of government spending occurs through reduced future spending. As a result, the standard wealth effect of government spending is absent.

focus on steady-state relationships which link particular parameters to first moments of the data or, in case parameters have no bearing on the steady state, turn to empirical studies which report appropriate estimates. We account for uncertainty of measurement by specifying a particular interval of plausible values for each parameter. As our VAR model is estimated on time-series data for the U.S. relative to an aggregate of industrialized countries, we mostly rely on evidence for the U.S., but also account for non-U.S. observations.

In order to generate sign restrictions which are robust across the entire range of plausible model parameterizations, we adopt the following procedure. Given the specified intervals and assuming a uniform and independent distribution, we draw a total of 100,000 realizations of the parameter vector. For each realization we compute impulse responses to a government spending and a technology shock. Finally, we compute confidence bounds containing 99 percent of the responses and analyze which variables respond unambiguously either positively or negatively to a particular shock for a specific number of periods after the shock.[11] In addition, we compute impulse responses to monetary policy shocks and anticipated government spending shocks, because in our empirical analysis we also account for these shocks once we assess the robustness of the results obtained under the baseline specification.

Table 1 about here

Table 1 summarizes the range of parameter values used in the model simulations. A period in the model is one quarter. As the discount factor relates, via the Euler equation, to observed after-tax returns, we chose the interval for the steady-state value of the discount factor to be consistent with annual after-tax returns between 4.2 and 7.5 percent, see Gomme and Rupert (2007) and references therein. The elasticity of substitution between varieties determines the markup, for which Rotemberg and Woodford (1993) find values between 20 and 40 percent. Correcting for a potential bias due to intermediate inputs reduces the lower bound to 7 percent.[12] For parameters governing the capital share and the depreciation rate, we allow for values consistent with a range of observations for the labor share and annual depreciation rates in various sectors of the economy, see Rotemberg and Woodford (1999) and Gomme and Rupert (2007), respectively. Government spending on domestic goods in steady state, g, is assumed to vary between 14 and 23 percent of output. Regarding the degree of home bias, we consider an interval which accounts for an export share between 7 and 12.5 percent.[13]

[11] Computing the impulse responses for a large number of realizations of the parameter vector ensures the robustness of our sign restrictions. Assuming a uniform distribution over the specified interval, we consider the entire range of parameter values, while dismissing all values outside the interval as implausible on a priori grounds. In order to dismiss very unlikely implications of realizations of the parameter vector, we consider 99 percent coverage bands.

[12] Here and in the following we draw parameters sequentially, and treat the earlier realizations as given when calculating the implied target values.

[13] The values for the export and government shares correspond to observations for the U.S. over the sample period used in our estimation (data sources are discussed below).

The parameters μ and γ jointly determine the Frisch elasticity and the intertemporal elasticity of substitution (IES). Our parameter intervals account for the uncertainty regarding appropriate values for these elasticities, see Basu and Kimball (2002) and Domeij and Flodén (2006). We also allow for higher values, in line with the early RBC literature.

The parameter governing the average price duration is somewhat controversial in the literature. Here we rely on international evidence reported by Dhyne et al. (2006). As a lower bound we employ the value for the U.S., yielding an average price duration of 6.7 months. The upper bound of 13 months is set in line with observations for the Euro Area. Regarding χ, the parameter capturing investment costs, we consider an interval which is centered around the point estimate of Christiano et al. (2005) and accounts for two standard errors. Concerning price indexation, we allow for the whole range from zero to full indexation. In order to specify monetary policy rules we consider an interval given by estimates for the U.S. and the Euro Area, reported by Clarida, Galí and Gertler (2000) and Enders, Jung and Müller (2009), respectively. We specify an admissible range of parameter values for the coefficients in the government spending feedback rule drawing on estimates reported by Galí and Perotti (2003) for a sample of OECD countries. The tax rule coefficient is chosen such that deficits display considerable persistence as in the data, see Corsetti and Müller (2008). Finally, for the persistence of technology shocks we sample from a range of two standard deviations around point estimates reported by Backus, Kehoe and Kydland (1992) and Heathcote and Perri (2002). Possible cross-country spillovers are admitted to take any value which does not lead to explosive behavior of the model.

Regarding the trade price elasticity σ, we consider two distinct intervals of empirically plausible values. One interval allows for low values, in line with evidence reported in a number of recent macroeconometric studies, see Enders and Müller (2009) for further discussion. A second interval allows for higher values of up to 2.5, the highest value which Backus et al. (1994) consider on the basis of surveying model-independent evidence. In our simulations we consider both intervals (each for 50,000 draws), but omit the middle range, as this will allow us to highlight the importance of the trade price elasticity for the international transmission mechanism and, in particular, for the sign of the exchange-rate response.[14]

Turning to our results, we display in figure 1 and 2 the impulse responses to an unanticipated innovation in government spending and technology, respectively. We consider both intervals for the trade price elasticity: the shaded area covers 99 percent of the responses in case values are drawn from the low interval (pointwise), the dashed lines cover 99 percent of the responses if values are drawn

[14]To be precise, the trade price elasticity interacts with all model parameters, but most importantly with the degree of home bias and the persistence of shocks in determining the sign of the exchange-rate response. As a result, there is no single threshold value for σ, which is independent of the values assumed for the other model parameters. Hence, we exclude a sizeable range between the two intervals. Note that we do not judge parameter values in this range as empirically implausible, but simply omit them in the simulation to illustrate that the trade price elasticity governs the sign of the exchange-rate response. Note also that we do not restrict the exchange-rate response in order to identify shocks in our estimated VAR model.

from the high interval. On the horizontal axis we measure periods after the shock (in quarters), on the vertical axis we measure responses in percentage deviation from steady-state values. We display the response of relative variables to a domestic shock, i.e., the difference in the response of a domestic variable and its foreign counterpart, because we are concerned with the behavior of the real exchange rate, which is determined by these relative variables.[15] For the real exchange rate and net exports we report the response of the domestic variable. We do not show responses for the terms of trade, because they move proportionally to the real exchange rate, as we assume home bias throughout.

Figure 1 about here

Figure 1 shows how the economy adjusts to unanticipated government spending shocks. Relative government spending increases for at least four quarters and possibly falls below its steady-state level thereafter, as it is systematically cut in response to higher public debt.[16] Output increases for at least two quarters, but a decline below steady state cannot be ruled out for later periods. The government budget deteriorates for at least four quarters. Importantly, the sign of the response of private consumption is ambiguous, once we consider the entire range of plausible model parameterizations.[17] Government spending crowds out investment for at least six quarters and triggers a rise in inflation. The nominal interest rate rises for at least four quarters.

The responses of net exports and the real exchange rate are ambiguous at all horizons. The sign of the response of net exports is determined by the trade price elasticity: if we sample from the low (high) interval, the response is positive (negative).[18] The real exchange rate appreciates, i.e., falls, in case we assume a high trade price elasticity—at least as long as government spending expands. If, instead, we assume a low trade price elasticity, the sign of the exchange-rate response is ambiguous. Corsetti et al. (2008a) show that, absent explicit risk-sharing across countries, the trade price elasticity determines the sign of the exchange-rate response to technology shocks, thereby possibly amplifying consumption risk. Our simulation results suggests that this is the case for government spending shocks, too.

Figure 2 about here

[15] Given the symmetry of the model, results are unchanged if we consider relative innovations, e.g., an exogenous increase in domestic government spending relative to foreign government spending.

[16] For ease of exposition, we do not always explicitly refer to the fact that we are dealing with variables in relative terms in the following discussion.

[17] Galí, López-Salido and Vallés (2007) analyze the transmission of government spending shocks in a new Keynesian closed economy model. They show that government spending raises private consumption only in the presence of labor market frictions and if a considerable fraction of households consumes disposable rather than permanent income. Our simulation results show that, although our model does not feature these frictions, we nevertheless cannot rule out a positive response of consumption to government spending. This is because we allow for a wide range of preference specifications while assuming infrequent price adjustment, see Bilbiie (2009) for a detailed analysis.

[18] See Müller (2008) for a detailed analysis of how the trade price elasticity (in relation to the IES) determines whether net exports rise or fall in response to government spending shocks.

Figure 2 shows how the economy adjusts to a positive technology innovation in the home country. The response of government spending is ambiguous as we allow for a positive and negative adjustment of government spending to output, which increases unambiguously for more than eight quarters. The budget does not fall on impact, because tax revenues are procyclical. The response of consumption is distinctly positive from quarter two to eight, after an initial period when a drop in consumption cannot be ruled out. The response of investment is positive during the first four quarters after the shock, inflation falls for at least two quarters and the interest rate drops for at least six quarters.

As with government spending shocks, the responses of net exports and the real exchange rate are governed by the trade price elasticity. If we consider only low values, net exports tend to fall, while they increase if a high trade price elasticity is assumed. Similarly, if a high elasticity is assumed the real exchange rate depreciates robustly, while its response is not clear-cut if a low trade price elasticity is assumed, in line with the results of Corsetti et al. (2008a). In sum, the model delivers sign restrictions for a number of variables, as their responses are qualitatively robust with respect to the entire range of plausible model parameterizations. Yet, at the same time, the model does not deliver unambiguous predictions as to how net exports and the real exchange rate respond to government spending and technology shocks.

Table 2 about here

Table 2 summarizes the sign restrictions implied by the model simulations. Specifically, the length and sign of the restrictions are given by the maximum number of quarters for which the simulations provide robust predictions for the sign of the response of a particular variable. Table entries indicate whether a variable is restricted to respond non-negatively (+), non-positively (-) to a specific shock, or whether it is left unrestricted (ø). Numbers indicate the quarters for which the response is restricted, with '0' indicating the impact period of the shock. In addition to unanticipated government spending shocks (column 1) and technology shocks (column 3), table 2 reports sign restrictions for monetary policy shocks (column 4) and for possibly anticipated government spending shocks (column 2). In this case, we compute for each realization of the parameter vector impulse responses to both unanticipated as well as to anticipated shocks considering an anticipation horizon of up to two quarters, i.e., we allow for $n \in \{0, 1, 2\}$. If a government spending shock is anticipated, the date at which it becomes known defines the impact period.[19]

Our set of sign restrictions ensures that productivity and fiscal shocks are distinguishable along several dimensions. The same holds for monetary policy shocks and potentially anticipated spending shocks. Nevertheless, without explicitly analyzing the responses to other shocks we cannot rule out that these

[19] In principle, recovering structural shocks from estimated VAR models is complicated by 'fiscal foresight', see Leeper, Walker and Yang (2009). For our setup, however, we find on the basis of the test suggested by Fernández-Villaverde, Rubio-Ramírez, Sargent and Watson (2007) that the mapping from VAR shocks to structural shocks is typically invertible.

shocks satisfy a particular set of sign restrictions, too. In this respect, however, including a relatively large number of variables provides some assurance. For instance, considering preference shocks which trigger an exogenous increase in private consumption demand, we find that responses are fairly similar to those of government spending shocks—except, that is, for the response of the government budget balance, which improves in response to preference shocks.

Before turning to our empirical specification, we note that several studies employ sign restrictions to identify fiscal and technology shocks. In particular, Corsetti, Dedola and Leduc (2009) investigate the effects of productivity and demand shocks on the real exchange rate using a six variable VAR estimated on U.S. times series relative to an aggregate of industrialized countries. They impose sign restrictions on labor productivity, manufacturing production (in country differences) and manufacturing production relative to GDP, for a horizon of 20 quarters. In order to discriminate between demand and productivity shocks they restrict the price of traded goods relative to the price of non-traded goods. In other words, their identification assumptions relate to a number of variables not included in our analysis. Interestingly, regarding the restrictions on output, we find that our model simulations provide no justification for a priori restricting the response over such a long horizon. Yet, as we will show below, the output restriction imposed by Corsetti, Dedola and Leduc (2009) is in fact satisfied by our estimated impulse response function. This observation may explain why we find similar effects of technology shocks on real exchange rates.[20]

3 New evidence on the behavior of U.S. real exchange rates

3.1 Data and baseline specification

We estimate the VAR model (1) on time-series data for the U.S. relative to an aggregate of industrialized countries consisting of the Euro Area, Japan, Canada and the U.K. ('rest of the world', for short). We include a constant and 4 lags of endogenous variables in the VAR model. The vector of endogenous variables consists of, in logs and real terms, private consumption, GDP, private investment, government spending as well as the primary budget balance scaled by GDP, inflation (measured using the GDP deflator), the nominal short-term interest rate and the log of the real exchange rate. We also report results for a specification which includes the log of the terms of trade instead of the real exchange rate. We focus on a post-Bretton-Woods sample, with data ranging from 1975Q1 to 2005Q4, as we omit the first two turbulent years after the collapse of the Bretton-Woods system.

For all variables we consider time-series data for the U.S., which we express, except for the real

[20]Mountford and Uhlig (2009) take a closed economy perspective and restrict output, consumption, and investment to increase in response to a business cycle shock. Government spending shocks are assumed to be orthogonal to business cycle shocks and to raise government spending. Fratzscher and Straub (2009) analyze the effects of asset price shocks on the current account and also discriminate shocks to technology and government spending through their differential impact on inflation. Both studies restrict responses for one year, rather than considering a variable-specific horizon.

exchange rate and the terms of trade, relative to the rest of the world. Data for real output, private consumption, government spending, the GDP deflator, and private fixed investment (excluding stock-building) are taken from the OECD Economic Outlook database. Government spending includes government spending on goods and services (government consumption), but neither investment nor transfers.[21] In addition, except for the Euro Area, we obtain from the same source data for the short-term interest rate, the primary government balance (measured in percent of GDP), exports of goods and services (value, local currency), imports of goods and services (value, local currency), and GDP (market prices). Net exports, as a fraction of GDP, are computed on the basis of these series. We use the series for the export price of goods and services (local currency) and the import price of goods and services (local currency) to measure the terms of trade. For the Euro area we obtain several series from the ECB's AWM database, see Fagan, Henry and Mestre (2001).[22] We obtain the CPI-based real effective exchange rate for the U.S. from the Main Economic Indicators of the OECD. In constructing the rest-of-the-world aggregate, we eliminate national basis effects by aggregating quarterly growth rates, weighted by each currency area's GDP.[23]

Under our baseline specification, we jointly identify unanticipated government spending shocks and technology shocks on the basis of the sign restrictions summarized in table 2. Since the search for impulse responses that fulfill all sign restrictions exactly is very cumbersome, we allow for small deviations using criterion (4) introduced above, setting $\varepsilon = 0.005$. Inference is based on 1000 draws satisfying the identification restrictions. The results reported below are robust with respect to assuming lower values of ε. In these cases, however, many draws from the Normal-Wishart posterior for the VAR parameters (B, Σ) receive zero prior weight, even if only few restrictions are mildly violated.

3.2 Effects of government spending and technology shocks

Given the estimated VAR model and the identified shocks to government spending and technology, we compute and display the corresponding impulse responses in figure 3 and 4. In all panels we plot the median as well as the 16 and 84 percent quantiles of the posterior distribution of impulse responses. In our discussion of the results we will use the term 'significance' whenever both quantiles are either above or below zero at a particular point in time. Shaded areas indicate that the sign of a response has been restricted over the corresponding horizon.

Figure 3 about here

[21] We do not consider government investment, because of an accounting problem in the U.K. in 2005Q2. We neither consider transfers to ensure consistency with our business cycle model.

[22] Specifically, we use the short-term interest rate (STN), the deflator of exports of goods and services (XTD), the deflator of imports of goods and services (MTD), and the government primary surplus (GPN_YEN).

[23] Euro area growth rates include West-Germany until 1990Q4, and unified Germany from 1991Q1 onwards (in case OECD data is used, similar adjustments have been applied in constructing the AWM database). Weights are based on PPP-adjusted values for the year 2000, as reported in the World Economic Outlook database (2007) of the IMF.

Figure 3 shows the effects of an exogenous innovation in relative government spending.[24] Government spending, displayed in the upper left panel, increases persistently. In line with the evidence reported by Perotti (2005) for a post-1980 sample as well as by Mountford and Uhlig (2009), we find a very short-lived increase in output in response to government spending shocks. The increase is limited to the period for which we restrict output to respond non-negatively. In fact, we find that significance bands cross the zero line while the response is still restricted to be non-negative, as a result of admitting small violations of our sign restrictions, see equation (4). The budget deteriorates persistently. Consumption, the response of which is left unrestricted, does not display a significant response, in line with evidence reported by Mountford and Uhlig (2009).[25] Investment shows a protracted decline, while inflation increases considerably. Interest rates, in turn, increase initially as long as they are restricted to respond non-negatively, but fall thereafter.

The middle and right panel of the last row show the response of the terms of trade and the real exchange rate. As discussed in the introduction, business cycle models under standard calibrations predict that government spending appreciates exchange rates, as do textbook versions of the Mundell-Fleming model. Yet, as shown above, our quantitative business cycle model fails to deliver robust predictions for how government spending impacts the real exchange rate (and the terms of trade), if one considers the entire range of plausible model parameterizations. Consequently, we do not restrict their responses and the obtained estimates constitute fresh evidence: government spending depreciates (raises) both the real exchange rate and the terms of trade.

This finding largely confirms evidence obtained under alternative identification schemes. Following Blanchard and Perotti (2002), several authors assume that government spending is predetermined in order to achieve identification. Kim and Roubini (2008) analyze U.S. times series for the period 1973–2002 and find that government spending shocks depreciate the real exchange rate; Monacelli and Perotti (2006) report similar results for Australia, the U.S. and the U.K., but not for Canada. Ravn et al. (2007) pool the data of all four countries, reporting a depreciation, too. Results for the terms of trade on the basis of this identification scheme are less clear-cut, see Corsetti and Müller (2006), Müller (2008), and Monacelli and Perotti (2008).[26]

Figure 4 about here

[24]Given that identification is based on sign restrictions derived from a symmetric business cycle model, we are agnostic as to whether relative government spending rises because domestic government spending rises in absolute terms or merely relative to foreign government spending. The same applies to technology shocks.

[25]The response of consumption to government spending shocks has been the subject of a considerable debate with different results emerging from different identification schemes based on short-run restrictions and narrative identification schemes, see Perotti (2007) and Ramey (2009), respectively.

[26]Beetsma, Giuliodori and Klaasen (2008) consider a panel of European countries and find that government spending appreciates the real exchange rate. The narrative approach to the identification of government spending shocks, suggested by Ramey and Shapiro (1998), is a widely considered alternative to the Blanchard-Perotti approach. It is employed by Monacelli and Perotti (2006) who find that government spending falls in response to the Carter-Reagan military build-up, while the real exchange rate depreciates.

Figure 4 shows the effects of a positive innovation to relative productivity. There is no effect on government spending, but output rises for an extended period, beyond the horizon that is restricted to be characterized by a non-negative response. There is also a beneficial and persistent effect on the government budget. The response of consumption and investment is also strongly positive, at horizons before and after restrictions are imposed. Inflation shows a persistent decline, as do interest rates. In the later case, the fall is limited to the period for which we impose restrictions.

The middle and right panel of the last row of figure 4 show the response of the terms of trade and the real exchange rate. As discussed in the introduction, business cycle models under standard calibrations predict that gains in relative productivity depreciate international relative prices. A notable exception is known as the Balassa-Samuleson effect, according to which productivity gains in the production of traded goods may appreciate the real exchange rate via their effect on the price of non-traded goods. However, even within our model, which does not allow for non-traded goods, we fail to detect a robust depreciation of the real exchange rate in response to positive technology shocks, since we consider a wide range of plausible model parameterizations. Consequently, we do not restrict the response of the real exchange rate and the terms of trade.

We find that exchange rates appreciate (fall) significantly during the first few quarters after a technology shock. To the extent that technology shocks not only appreciate the real exchange rate, but also the terms of trade (computed on the basis of import and export price indices), the exchange-rate response does not merely reflect a Balassa-Samuelson effect, but a general increase in the relative price of domestically produced goods. Our findings are in line with results reported by Corsetti et al. (2008b), Kim and Lee (2008), and Enders and Müller (2009), who, drawing on Galí (1999), use long-restrictions to identify technology shocks.[27] Moreover, Corsetti, Dedola and Leduc (2009), identifying demand and productivity shocks on the basis of a set of sign restrictions which differs from ours, also find that productivity shocks appreciate U.S. exchange rates.[28]

A new finding relative to earlier studies, however, concerns the medium-term dynamics of the real exchange rate. We find an appreciation only for the first couple of quarters. Afterwards, the exchange rates starts to rise above its pre-shock level and shows a significant depreciation for an extended period. This strikes us an interesting finding, notably because there is no evidence for a reversal of the sign of the terms of trade response.[29]

[27] Corsetti et al. (2008b) and Kim and Lee (2008) specify their VAR model in relative terms and identify technology shocks by assuming that only these shocks affect relative labor productivity in the long run. Corsetti et al. (2008b) report that relative technology shocks appreciate the real exchange rate and the terms of trade in the U.S. and Japan. Kim and Lee (2008) find an exchange-rate appreciation for the U.S., but not for Japan and the Euro area. Enders and Müller (2009) assume instead that only technology shocks affect the level of U.S. labor productivity in the long run. They also find an appreciation of the U.S. terms of trade and the exchange rate.

[28] They find that demand shocks appreciate exchange rates, too. This result does not conflict with our result on government spending shocks, which, in equilibrium, simultaneously affect the supply and demand of domestic goods.

[29] Corsetti, Dedola and Leduc (2009) also detect some signs of a long-run depreciation in response to productivity shocks, but only at about 35 quarters after impact.

3.3 Exchange rate dynamics: further analysis

In the following we take up a number of issues to shed further light on our results. First, to give a more systematic account of the uncertainty surrounding the median responses, we follow Scholl and Uhlig (2008) and report in figure 5 the posterior joint distribution of the timing and the size of the peak responses of the real exchange rate and the terms of trade. The distribution of peak responses to government spending and technology shocks are displayed in the left and right column, respectively, against the size of the response and the quarter when the peak response occurs. Overall, the distribution of peaks is fairly well behaved, with almost the entire mass of the distribution leaning towards the median response.

Figure 5 about here

Note, however, that the posterior distribution reflects both sampling and model uncertainty. In order to gauge the extent of model uncertainty, we rule out sampling uncertainty by holding the coefficients fixed at the OLS point estimates when computing the posterior distribution of impulse response functions. The solid lines in figure 6 display the median as well as the 16 and 84 percent quantiles of the posterior distribution of the responses of the real exchange rate and the terms of trade. While considerable model uncertainty is apparent, the posterior distribution of the responses is tighter relative to the results reported in figure 3 and 4.

Figure 6 about here

Finally, we note that focusing on the median of the posterior distribution of impulse responses might be problematic, particularly if several structural shocks are identified. Fry and Pagan (2007) point out that the posterior distribution of impulse responses is a distribution across different identified models such that the median impulse response functions to two shocks are generated by two different impulse matrices. Consequently, the identified shocks associated with the median responses are not necessarily orthogonal. To explore this issue further, we generate impulse responses on the basis of a single model, by computing an impulse matrix which implies responses as close to the median responses as possible, as suggested by Fry and Pagan (2007). The responses, displayed by the dashed line in figure 6, show that the results obtained under our baseline specification are not very sensitive to this adjustment. In particular, we still find a short-run appreciation and the medium-run depreciation of the real exchange rate in response to technology shocks.

3.4 Accounting for fluctuations

We now turn to a brief analysis of the actual incidence of government spending and technology shocks as suggested by our identification scheme. In figure 7 we plot four-quarter moving averages of the estimated innovations. In the left panels, the solid lines display the median estimates, while the dashed lines display 16 and 84 percent quantiles. However, since the median shock series result from different identified models, they are potentially correlated. We therefore apply once more the procedure suggested by Fry and Pagan (2007) and compute innovations from a single model. The results are displayed in the middle column of figure 7. The solid line refers to the median innovations holding the VAR parameters fixed at the OLS point estimates, while the dashed lines display the innovations obtained under the single model.

Figure 7 about here

Despite considering four-quarter moving averages, the picture of the entire history of identified shocks is fairly complex. Yet a few episodes stand out and may be related to familiar narratives concerning important macroeconomic episodes during the last three decades. Focusing on the results obtained under the single model (middle panels, dashed line) and turning first to government spending shocks, we observe spikes during the Carter-Reagan military build-up in the early 1980s as well as after 9/11. Regarding technology shocks, strong positive innovations during the late 1990s can be detected, in line with the notion of a distinct productivity driven upturn in the U.S. at that time.

Table 3 about here

The right panels of figure 7 plot a historical decomposition of the real exchange rate, i.e., a comparison of the time series predicted by the VAR model assuming that all shocks occurred (solid line) and that either only technology or government spending shocks occurred (dashed line). In both scenarios we assume zero as the starting value of the vector of endogenous variables in order to abstract from initial conditions. A casual inspection suggests that technology shocks—more than government spending innovations—account for a considerable fraction of actual exchange rate dynamics.

A similar picture emerges once we compute a business cycle variance decomposition as in Altig, Christiano, Eichenbaum and Lindé (2005). Specifically, we compute, on the basis of counterfactual simulations using the single model, the fraction of the variance of each time series that is accounted for by either government spending or technology shocks. In table 3, we report the fraction of the variance of the corresponding counterfactual time series relative to the variance of the actual data, after applying the HP-filter with a smoothing parameter of 1600 to each series (in brackets, we report the corresponding statistics based on unfiltered data).

According to this measure, technology shocks account for 21, 11, and 10 percent of the short-run fluctuations of consumption, output, and investment, while government spending shocks account for 7, 12, and 12 percent, respectively. Both shocks have a large impact on the cyclical volatility of inflation: 29 and 21 percent of the variation is due to technology and spending shocks, respectively.

Turning to the real exchange rate, we find that 18 percent of its business cycle variance is due to shocks to technology, while government spending shocks account for 9 percent of fluctuations. Yet the role of both shocks in accounting for fluctuations appears to be much smaller once we consider unfiltered times series, suggesting that both shocks are of minor importance at median and long-run frequencies.

4 Sensitivity analysis

4.1 Anticipation of government spending shocks

As a result of the institutional features of the budget process, innovations in government spending may become known before they are actually implemented. Blanchard and Perotti (2002) discuss this issue and show how accounting for anticipation requires stronger identification assumptions within their framework. They explicitly investigate the possibility that shocks are known one quarter in advance and find somewhat stronger output effects.[30] Mountford and Uhlig (2009) argue that the sign restriction approach to identification is particularly well-suited to address the issue of announcement effects. In order to identify anticipated spending shocks, they restrict government spending not to respond for four quarters and find, in comparison with unanticipated shocks, more persistent and stronger effects of government spending, notably on output and consumption; the latter increasing significantly only in response to anticipated shocks.

Figure 8 about here

In order to allow for the possibility that spending shocks are anticipated, we impose the sign restrictions reported in the second column of table 2. In other words, we rely on our model simulations which explicitly allow for implementation lags in government spending innovations of up to two quarters. Recall that the sign restrictions reported in the second column are satisfied by anticipated and unanticipated spending shocks, such that we are agnostic as to whether shocks have been anticipated or not. In our view, working with a fully specified general equilibrium model to derive sign restrictions ensures capturing non-trivial feedback effects of anticipated innovations. For instance, we find that government spending, if anticipated to rise exogenously, may nevertheless adjust instantaneously—to the extent that it responds endogenously to the state of the economy.

As in our baseline specification we identify (possibly anticipated) government spending and technology shocks jointly. In fact, the only difference relative to the baseline specification is the set of sign restrictions: we use, in addition to column 3, column 2 of table 2, rather than column 1. Figure 8

[30] See Mertens and Ravn (2009) for a general treatment of how to account for anticipation effects while employing the Blanchard-Perotti identification scheme.

shows the resulting impulse responses of the terms of trade and the real exchange rate. We detect only minor differences in the effects of government spending and technology shocks.[31]

4.2 Monetary policy shocks

Monetary policy shocks may contribute considerably to real exchange rate fluctuations, see, e.g., Clarida and Galí (1994) and Eichenbaum and Evans (1995). We investigate whether results obtained under our baseline specification are sensitive to a modification where we explicitly identify monetary policy shocks, in addition to government spending and technology shocks. To do so, we rely on sign restrictions implied by our model simulations and reported in the right column of table 2.[32] We restrict all three structural shocks to be orthogonal to each other.

Figure 9 displays the impulse responses of the real exchange rate and the terms of trade obtained under this three-shock identification scheme. We find the responses to government spending and technology shocks hardly altered relative to the baseline specification.

Figure 9 about here

4.3 Further sensitivity analysis

In the following we take up additional complications to assess the robustness of the results obtained under our baseline specification. First, we investigate whether our results are sensitive to the inclusion of net exports as an additional variable in the VAR model. Our quantitative business cycle model does not deliver clear-cut prediction for how net exports respond to any of the shocks we seek to identify. We thus leave their response unrestricted and impose the same restrictions as in the baseline specification.

Figure 10 shows that our results are not much affected by the inclusion of net exports. The initial exchange-rate appreciation after technology shocks, however, is no longer significant, most likely reflecting the fact that the number of variables in the VAR model increased, while the number of restrictions did not. The response of the trade balance displays significant dynamics only if the terms of trade are included in the VAR.[33] Government spending shocks tend to improve U.S. net exports, albeit by a limited amount, see Kim and Roubini (2008) and Corsetti and Müller (2006). For technology shocks, we find a hump-shaped decline of net exports, after an initial increase; Enders and Müller (2009) report similar adjustment dynamics.

[31]We also find responses very similar to those obtained under the baseline specification, once we impose sign restrictions which are generated by model simulations while allowing for an anticipation horizon of up to four quarters. Results are available on request.

[32]Note that we restrict the impact response of exchange rates to monetary policy shocks, in line with the predictions of our model.

[33]These are shown in figure 10. Results for the real exchange rate specification are very similar, but insignificant.

Figure 10 about here

Finally, we consider two alternative starting dates for our sample. First, we use data from the end of the Bretton-Woods system onwards, i.e., we set 1973Q1 as a starting date. Second, we use 1980Q1 as a starting date, as U.S. fiscal policy transmission arguably changed afterwards, see Perotti (2005). Figure 11 displays the median responses of the exchange rate and the terms of trade to government spending and technology shocks, contrasting results for the baseline sample (solid line) with those for the earlier (dashed line) and the later (dashed-dotted line) starting period. We find that results are fairly robust across sample periods.

Figure 11 about here

5 Conclusion

In this paper we establish evidence on the response of international relative prices to government spending and technology shocks. We start from three observations. First, the behavior of the real exchange rate and the terms of trade carries important information regarding the international transmission mechanism. Second, the existing evidence on the behavior of international relative prices in response to technology and government spending shocks conflicts with the predictions of international business cycle models under standard calibrations. Third, this evidence is mostly based on estimated VAR models where identification is achieved either through short-run or long-run restrictions.

We provide new evidence by employing an alternative identification scheme based on sign restrictions. In order to generate robust sign restrictions, we simulate a quantitative business cycle for a wide range of parameter values. Moreover, we document that the model does not deliver clear-cut predictions for the sign of the exchange-rate response to government spending and technology shocks. To identify these shocks, we thus restrict the responses of several variables, but leave the response of exchange rates unrestricted.

Estimating a VAR model on time-series data for the U.S. relative to an aggregate of industrialized countries, we find that expansionary government spending shocks depreciate the real exchange rate and the terms of trade. Positive technology shocks, in contrast, appreciate the real exchange rate and the terms of trade in the short run. These results are robust with respect to several variations of our baseline specification, such as accounting for monetary policy shocks or anticipation of innovations in government spending.

In principle, our empirical results can be rationalized within standard business cycle models. After all, our model simulations illustrate that the exchange-rate response to both shocks can go either way, depending on the parameterization. As conventionally calibrated models predict that government spending (technology) shocks appreciate (depreciate) exchange rates, one might interpret our findings as evidence in support of an alternative calibration, characterized, in particular, by a low trade price elasticity. For this case, Corsetti et al. (2008a) show that strong wealth effects in response to technology shocks raise the demand for domestic goods above supply and thus appreciate the exchange rate. Our simulation results suggest that the reverse holds for government spending shocks. However, a more fundamental reassessment of the international transmission holds considerable promise, too. Results by Ravn et al. (2007) and Corsetti, Meier and Müller (2009), for instance, suggest that 'deep habits' or 'spending reversals' may account for an exchange-rate depreciation following an exogenous increase in government spending. Further research into the international transmission mechanism along these lines seems warranted.

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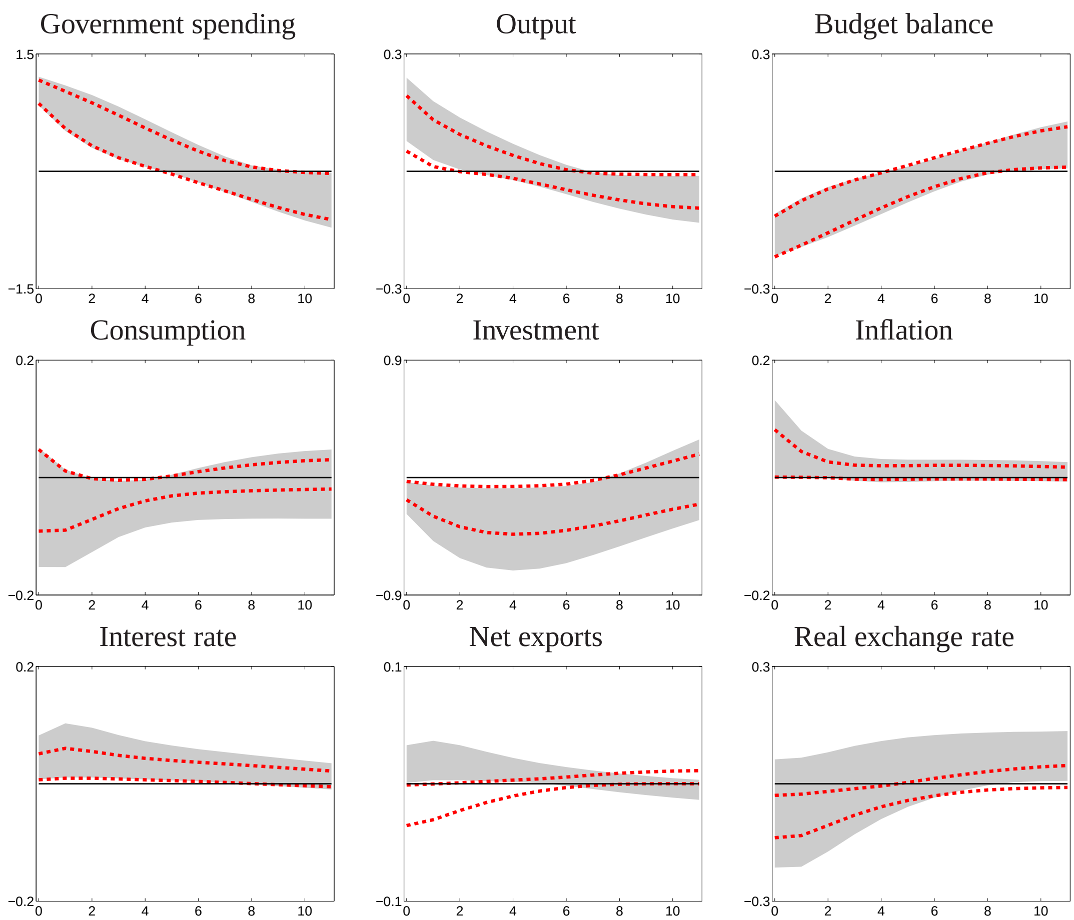

Figure 1: Model responses to exogenous and unanticipated increase in domestic government spending under various parameterizations. Notes: Responses are measured in relative terms, i.e., Home less Foreign, except for net exports and the real exchange rate (Home); shaded area covers 99 percent of responses (pointwise) assuming a low trade price elasticity; dashed lines display the same statistic assuming a high trade price elasticity; number of draws: 100,000. Horizontal axes: quarters; vertical axis: percentage deviation from steady state.

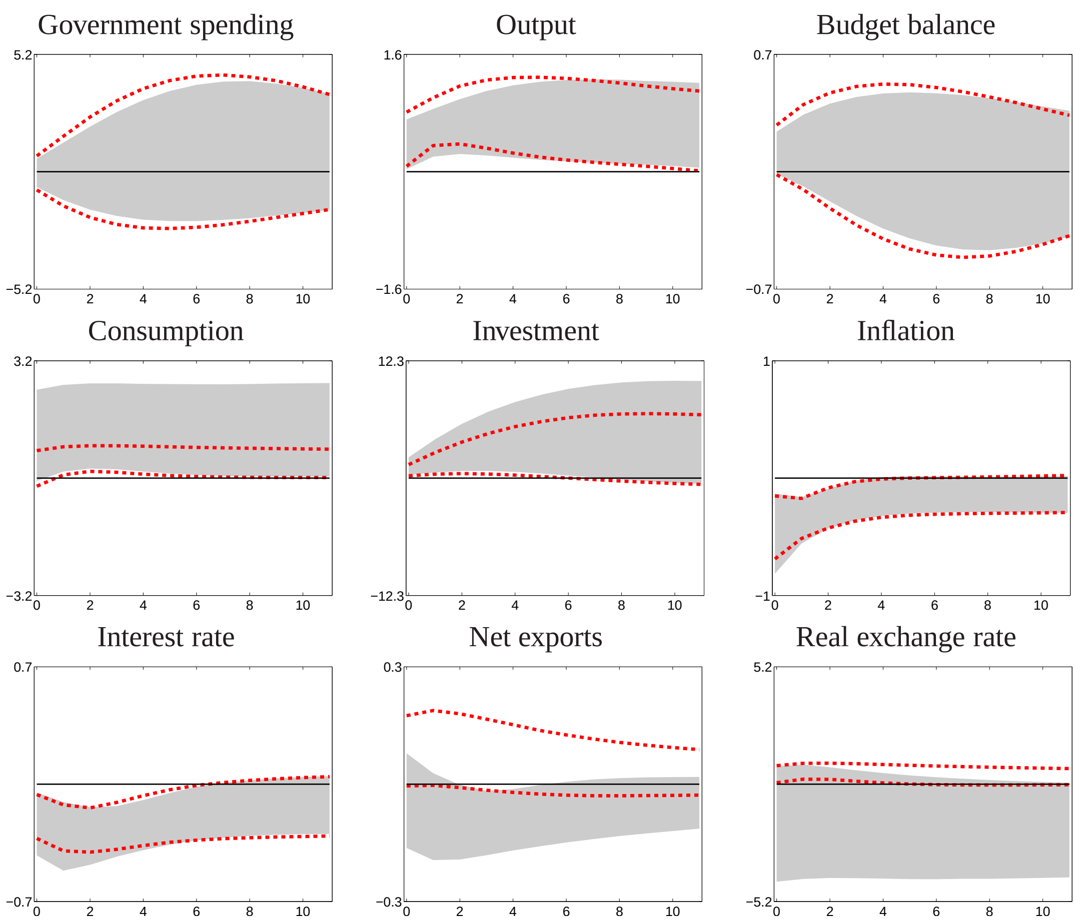

Figure 2: Model responses to positive technology innovation under various parameterizations. Notes: see figure 1.

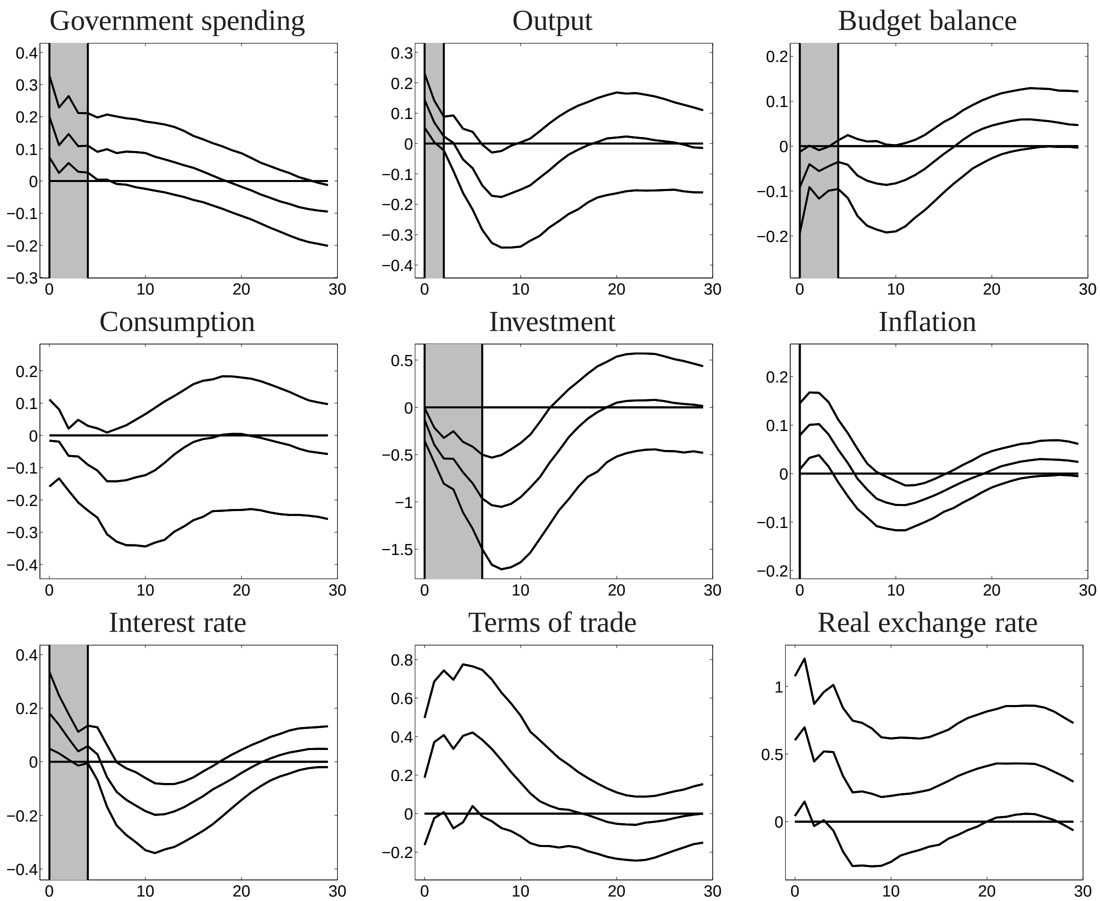

Figure 3: Responses to identified government spending shock (baseline specification). Notes: solid lines display the median response and the 16 and 84 percent quantiles. Shaded areas indicate sign restrictions. Horizontal axes: quarters; vertical axis: percent. All variables, except for the real exchange rate and the terms of trade, are expressed in relative terms (U.S. vs. ROW).

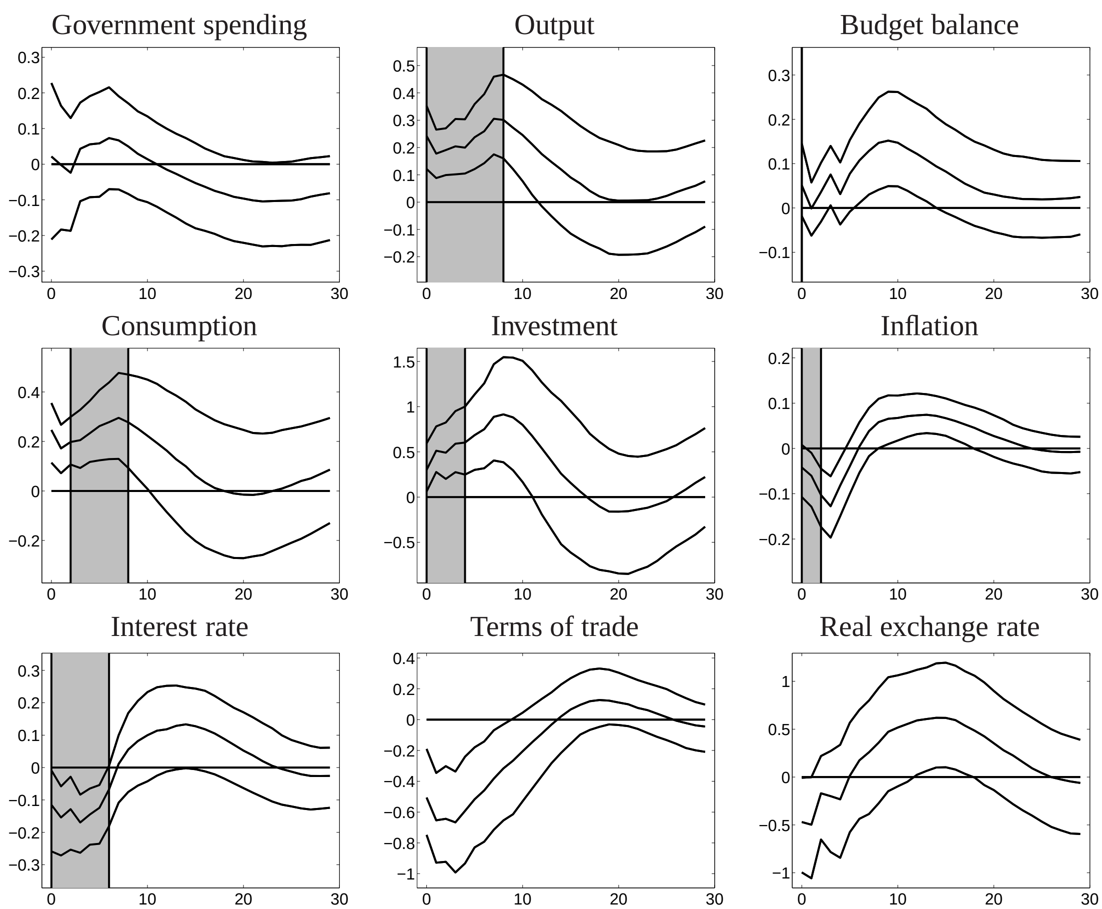

Figure 4: Responses to identified technology shock (baseline specification). Notes: see figure 3.

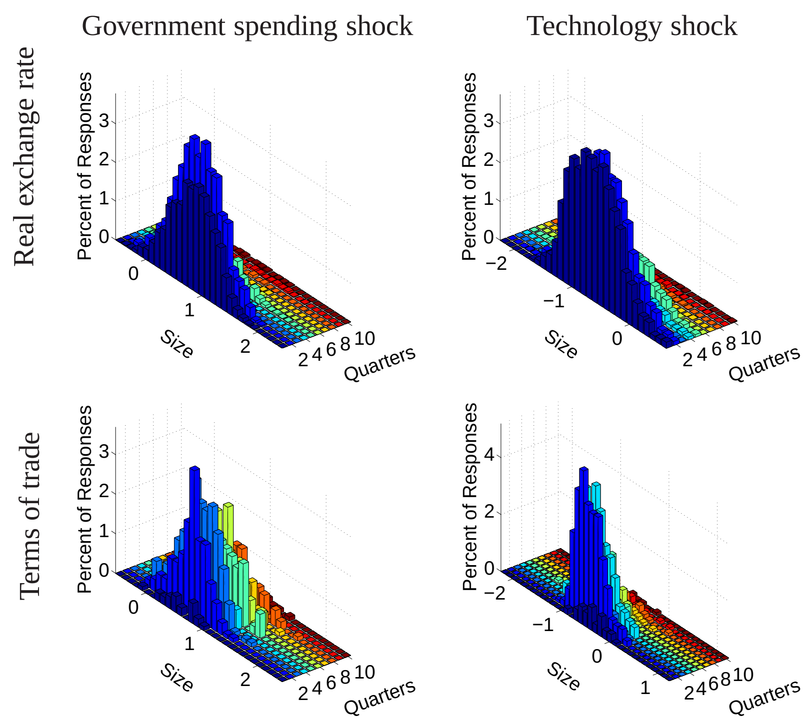

Figure 5: The posterior distribution of the peak responses. Notes: posterior joint distribution of quarter and size of the maximal absolute value of the response to a government spending shock (left) and technology shock (right) within the first 10 quarters.

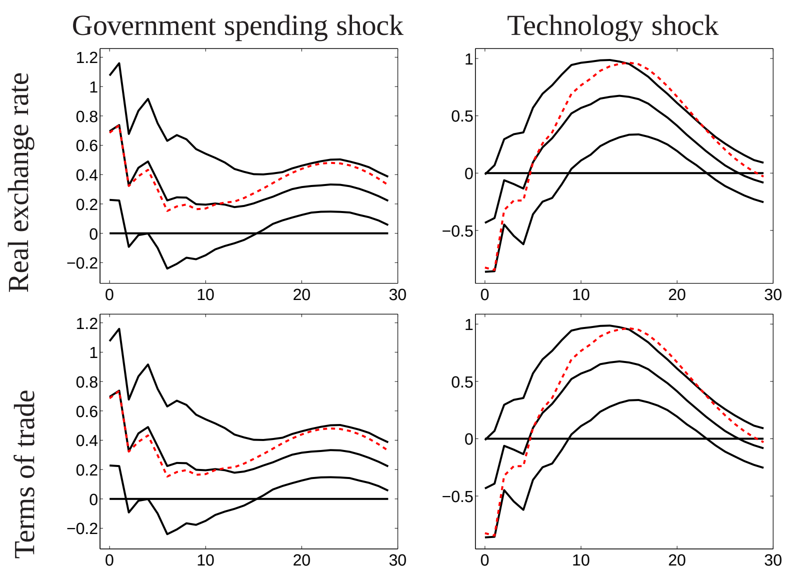

Figure 6: Impulse responses of real exchange rate and terms of trade. Notes: VAR coefficients are fixed at their OLS estimates. Solid lines: median response and 16 and 84 percent quantiles; dashed line: impulse responses implied by a single model as proposed by Fry and Pagan (2007). Horizontal axes: quarters; vertical axes: percent.

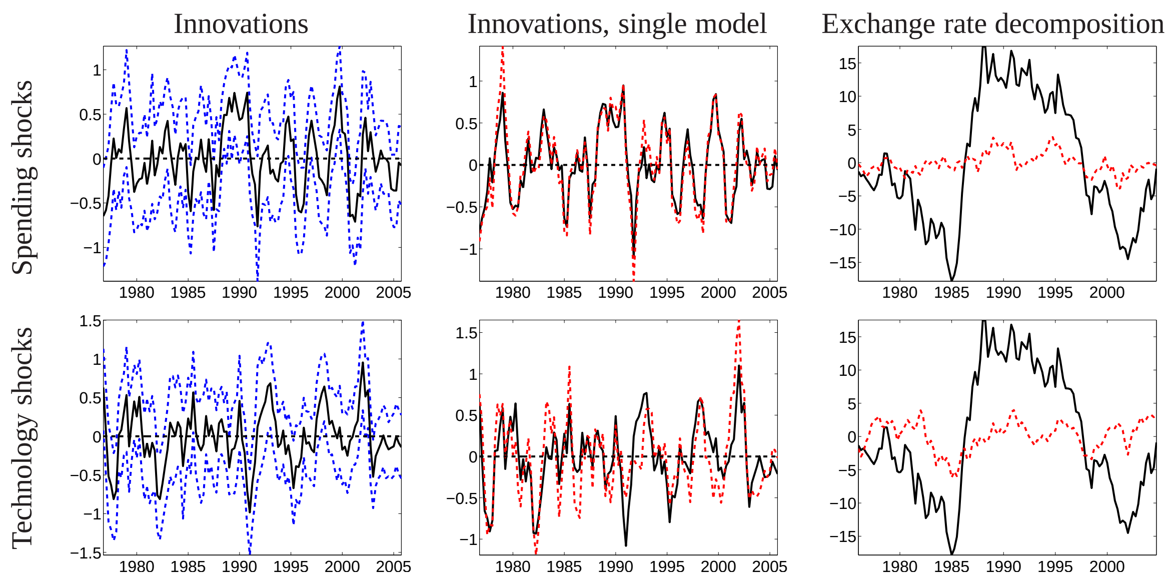

Figure 7: Estimated innovations and historical decomposition of real exchange rate fluctuations. Notes: left panels show four-quarter moving average of estimated innovations (median, 16 and 84 percent quantiles); middle panels show innovations for VAR parameters fixed at OLS point estimates (solid line) and innovations implied by single model (dashed line); right panels show historical decomposition of real exchange rate predicted by VAR assuming zero as starting value for the vector of endogenous variables: all shocks (solid line) vs. spending or technology shocks only (dashed line).

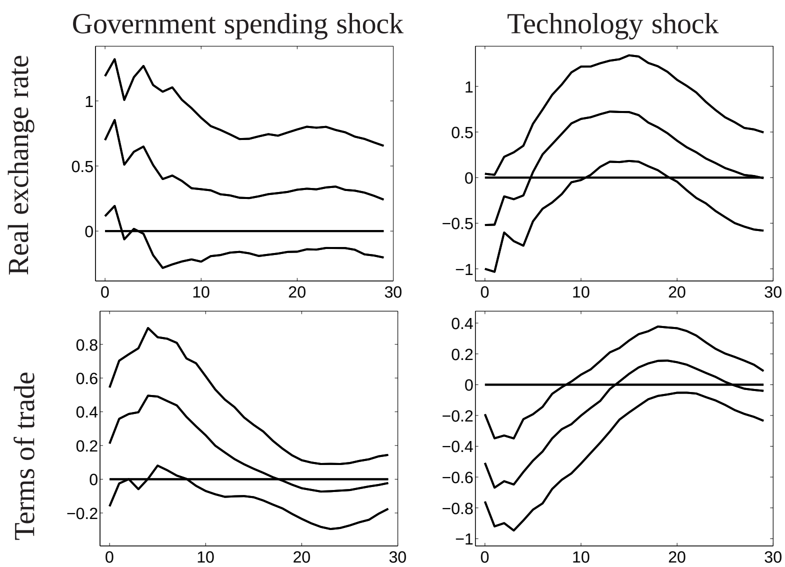

Figure 8: Responses to technology and possibly anticipated government spending shocks; Notes: see figure 3.

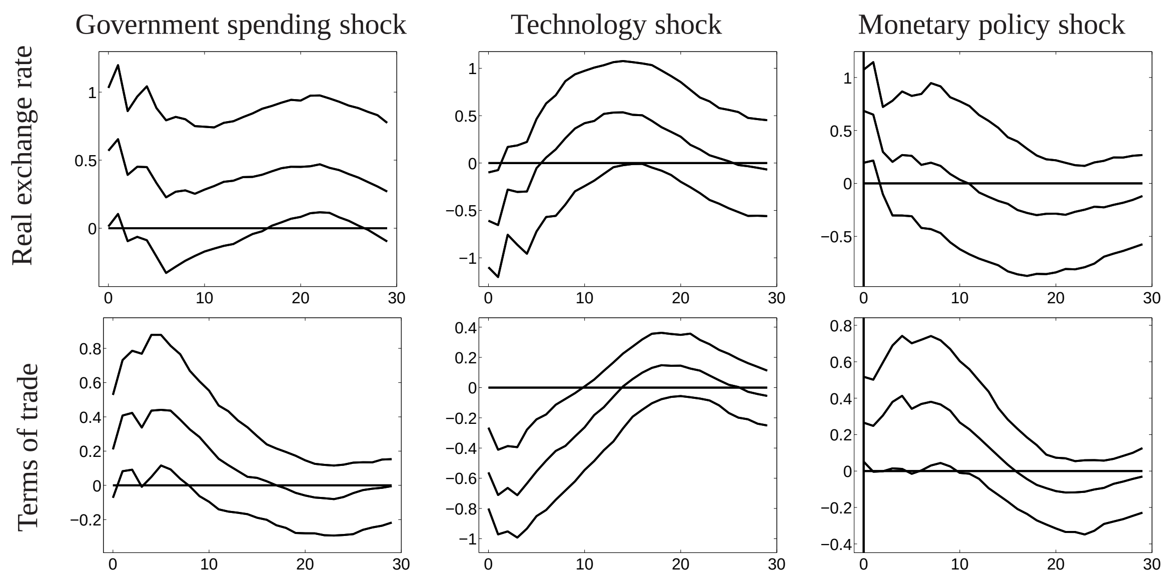

Figure 9: Responses under three-shock identification scheme. Notes: see figure 3.

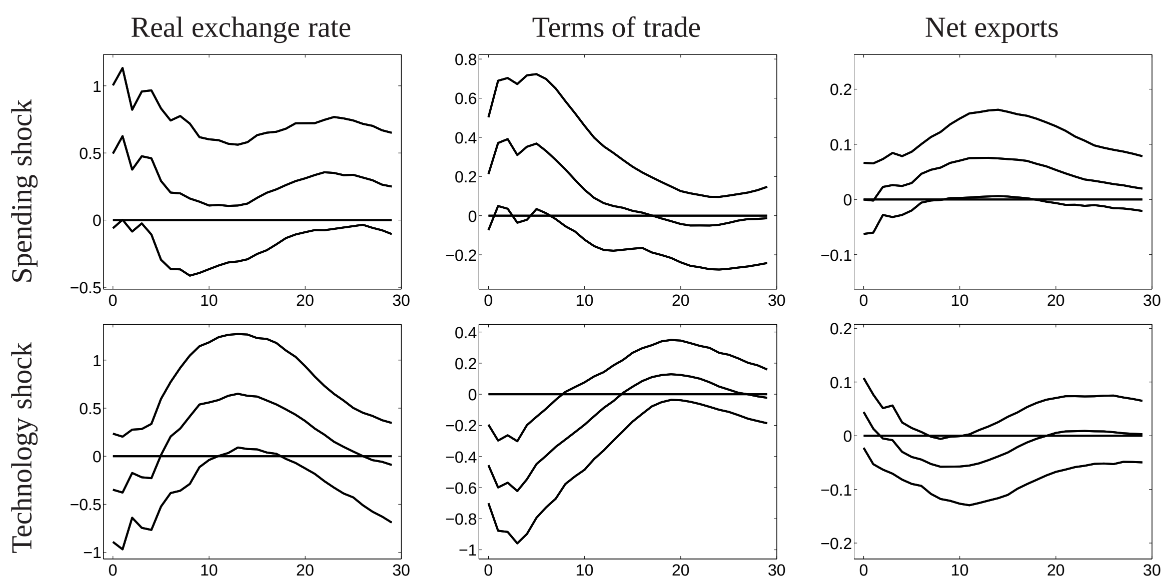

Figure 10: Responses of VAR model which includes net exports. Notes: see figure 3.

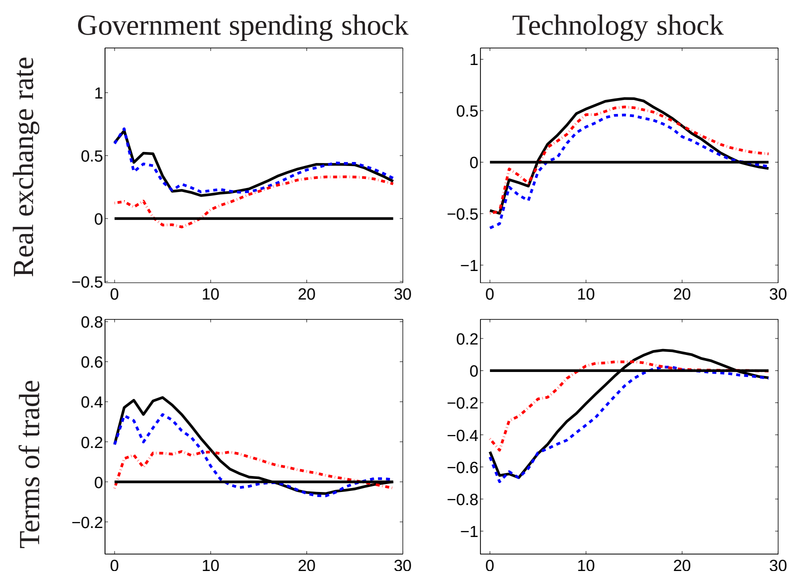

Figure 11: Responses for different samples. Notes: see figure 3; solid line: baseline sample; dashed line: sample starts in 1973Q1; dashed-dotted line: sample starts in 1980Q1.

Table 1: Parameter values used in simulation

	Parameter description	Range	Target / Source	
β	Discount factor (steady state)	$[.982, 0.99]$	After-tax return	$[.042, .075]$
ϵ	Elasticity of substitution	$[3.50, 15.0]$	Markup	$[1.07, 1.40]$
θ	Capital share	$[0.15, 0.39]$	Labor share	$[0.57, 0.78]$
δ	Depreciation rate ($\times 100$)	$[0.38, 6.02]$	Various sectors	$[.015, 0.22]$
g	Government spending (steady state)	$[0.14, 0.23]$	Government share	$[0.14, 0.23]$
ω	Home bias in final goods	$[0.84, 0.92]$	Export share	$[0.07, .125]$
μ	Consumption weight in utility	$[0.02, 0.89]$	Frisch elasticity	$[0.18, 4.00]$
γ	Risk aversion	$[1.00, 20.0]$	IES	$[0.48, 1.00]$
ξ	Fraction of prices kept unchanged	$[0.55, 0.77]$	Price duration (months)	$[6.70, 13.0]$
χ	Investment adjustment costs	$[2.05, 2.91]$	Christiano et al. ’05	
ι	Indexation of prices	$[0.00, 1.00]$	All admissible values	
ϕ_π	Inflation response of interest rate	$[1.68, 2.15]$	Clarida et al. ’98 / Enders et al. ’09	
ϕ_y	Output response of interest rate	$[0.78, 0.93]$	”	
ρ_r	Interest rate smoothing	$[0.62, 0.79]$	”	
ρ_g	Government spending persistence	$[0.70, 0.85]$	Galí & Perotti ’03	
φ_y	Output gap response of G-spending	$[-0.2, 0.20]$	”	
φ_d	Debt response of G-spending	$[1-\beta, 0.04]$	”	
φ_τ	Debt response of tax rate	$[0.00, 0.04]$	Corsetti & Müller ’08	
ρ_z	Technology persistence	$[0.83, 0.98]$	Backus et al. ’92 / Heathcote & Perri ’02	
ρ_{zz}	Technology spillover	$[0, 0.99 - \rho_z]$	Remaining admissible values	
σ	Trade price elasticity	$[0.10, 0.33]$ or $[1.00, 2.50]$	Enders & Müller ’09 / Backus et al. ’94	

Notes: Parameter values used in simulation of the model. ‘Range’ specifies interval from which values of the parameter vector are drawn for each simulation of the model.

Table 2: Sign restrictions implied by model simulations

	Expansionary shock to			
	Government spending		Technology	Monetary policy
Response of	No anticipation	+Anticipation		
Private consumption	ø	ø	$+_{2-8}$	$+_{0}$
Output	$+_{0-2}$	$+_{2}$	$+_{0-8}$	$+_{0}$
Investment	$-_{0-6}$	$-_{0-6}$	$+_{0-4}$	$+_{0}$
Government spending	$+_{0-4}$	$+_{2-4}$	ø	ø
Government budget	$-_{0-4}$	$-_{2-4}$	$+_{0}$	$+_{0}$
Net exports	ø	ø	ø	ø
Nominal interest rate	$+_{0-4}$	$+_{2-4}$	$-_{0-6}$	$-_{0}$
Inflation	$+_{0}$	ø	$-_{0-2}$	$+_{0}$
Real exchange rate/terms of trade	ø	ø	ø	$+_{0}$

Notes: responses of variables (in relative terms) are restricted to be non-negative (+), non-positive (-) or unrestricted (ø). Numbers indicate the time periods after the shock for which the responses are restricted. The column '+Anticipation' refers to potential anticipation of up to two quarters.

Table 3: Business Cycle Variance Decomposition

	Business cycle variation due to shocks to	
Variable	Government spending	Technology
Private consumption	0.07 [0.02]	0.21 [0.08]
Output	0.12 [0.05]	0.11 [0.08]
Investment	0.12 [0.05]	0.10 [0.04]
Government spending	0.08 [0.04]	0.06 [0.04]
Government budget	0.06 [0.03]	0.10 [0.04]
Nominal interest rate	0.17 [0.08]	0.17 [0.08]
Inflation	0.21 [0.17]	0.29 [0.26]
Real exchange rate	0.09 [0.03]	0.18 [0.06]

Notes: Variance of counterfactual relative to actual time series (after applying HP-filter with smoothing parameter of 1600 [without filtering in brackets]); counterfactual time series are computed on the basis of VAR model and identified shocks (single model).