Research Division
Federal Reserve Bank of St. Louis
Working Paper Series

Sources of Exchange Rate Fluctuations: Are They Real or Nominal?

Luciana Juvenal

Working Paper 2009-040B
http://research.stlouisfed.org/wp/2009/2009-040.pdf

August 2009
Revised May 2010

FEDERAL RESERVE BANK OF ST. LOUIS
Research Division
P.O. Box 442
St. Louis, MO 63166

Sources of Exchange Rate Fluctuations: Are They Real or Nominal?*

Luciana Juvenal†
Federal Reserve Bank of St. Louis

May 4, 2010

Abstract

I analyze the role of real and monetary shocks on the exchange rate behavior using a structural vector autoregressive model of the US *vis-à-vis* the rest of the world. The shocks are identified using sign restrictions on the responses of the variables to orthogonal disturbances. These restrictions are derived from the predictions of a two-country DSGE model. I find that monetary shocks are unimportant in explaining exchange rate fluctuations. By contrast, demand shocks explain between 21% and 37% of exchange rate variance at 4-quarter and 20-quarter horizons, respectively. The contribution of demand shocks plays an important role but not of the order of magnitude sometimes found in earlier studies. My results, however, support the recent focus of the literature on real shocks to match the empirical properties of real exchange rates.

Keywords: Exchange Rates, Real Shocks, Monetary Shocks, Vector Autoregression, Sign Restrictions.
JEL Classification: F31; F41; C30.

*I am grateful to the Editor and two anonymous referees for their constructive comments; to Renee Fry, Gian Maria Milesi-Ferretti, and Adrian Pagan for helpful suggestions and advice at various stages of this paper; to Robert Vigfusson for a very useful discussion; and to Giacomo Carboni, Mike Clements, Valentina Corradi, Riccardo DiCecio, Liam Graham, Sheheryar Malik, John Rogers, Lucio Sarno, Mark P. Taylor, David Wheelock, and seminar participants at the University of Warwick, the City University of New York, Trinity College Dublin, the Federal Reserve Bank of St. Louis, the Bank of Spain and the System Committee on International Economic Analysis 2008 for comments. I am also grateful to Brett Fawley for research assitance and Judy Ahlers for editorial assistance. The views expressed are those of the author and do not necessarily reflect official positions of the Federal Reserve Bank of St. Louis, the Federal Reserve System, or the Board of Governors.

†Correspondence address: Research Division, Federal Reserve Bank of St. Louis, P.O. Box 442, St. Louis, MO 63166-0442. Email: luciana.juvenal@stls.frb.org

1 Introduction

The explanation of the sources of real exchange rate fluctuations is one of the most challenging issues in international economics. From a theoretical standpoint, the literature has focused on the leading role of monetary policy shocks in accounting for real exchange rate movements. The empirical evidence, however, has not shown much support that monetary policy shocks are important in determining the real exchange rate. This paper combines the predictions of a standard DSGE open-economy macro model with recent econometric developments to examine the impact of real and monetary shocks on real exchange rate behavior.

The theoretical focus on monetary shocks as a main driver of the real exchange rate has a long tradition in international economics. This refers not only to the seminal work by Dornbusch (1976) but also to the large body of literature that developed afterward, including DSGE models of the real exchange rate (see Beaudry and Devereux, 1995, and Chari, Kehoe and McGrattan, 2002, among others). Overall, the belief that monetary policy plays a dominant role in explaining real exchange rate fluctuations has long been an accepted fact in economics. Such is the case that Rogoff (1996, p. 647) highlights:

"Most explanations of short-term exchange rate volatility point to financial factors such as changes in portfolio preferences, short-term asset price bubbles, and monetary shocks."

Although this quote reflects the "consensus" that nominal shocks are the main drivers of exchange rate fluctuations, the empirical evidence has not provided much support for this idea. In a seminal paper, Clarida and Galí (1994) estimate the effect of various shocks on real dollar bilateral exchange rates. They find that the contribution of monetary shocks to the variance of the real exchange rate is less than 3% for the UK and Canada. By contrast, demand shocks explain more than 95% of the movement of the real exchange rate both at short and long horizons. Eichenbaum and Evans (1995) study the effects of monetary policy shocks using three different models. Their results show that the mean contribution of monetary policy shocks is less than 25%.

In a recent study, Steinsson (2008) challenges the focus of the theoretical literature on monetary shocks and shifts the analysis to the role of real shocks as drivers of the real exchange rate. He shows that in response to real shocks the real exchange rate exhibits a dynamic that matches the data in terms of volatility and persistence. By contrast, monetary shocks are unable to match the empirical persistence of real exchange rates.

Motivated by the previous literature, I analyze the impact of both real and monetary shocks on the real exchange rate. To do this, I first study the effects of productivity, demand, and monetary shocks using the richly specified two-country DSGE model developed by Ferrero, Gertler and Svensson (2008). I analyze the predictions of the model for a standard parameterization, which yields robust responses of key macroeconomic variables. These responses allow me to derive identifying assumptions to estimate a VAR model based on sign restrictions.

The selection of a VAR model approach to address this type of question is a natural one. In fact, most empirical work on the sources of exchange rate fluctuations builds on the use of an sVAR method to identify monetary policy shocks. However, consensus is lacking on the results; part of the disagreement originates in the choice of the estimation method. The identification strategies in use vary from study to study, and the ones that rely on conventional estimation techniques are subject to some criticism. Those using zero short-run restrictions (Eichenbaum and Evans, 1995) identify shocks of interest based on some assumptions that may be difficult to reconcile with a broad range of theoretical models. For example, the assumption of zero contemporaneous impact of a monetary policy shock on output is in contrast with some general equilibrium models (see Canova and Pina, 1999). Faust and Rogers (2003) show that when "dubious" assumptions are relaxed, the effects of monetary policy shocks on the exchange rate are small. Although long-run restrictions are often better justified by economic theory, in some cases substantial distortions can arise due to the presence of a small-sample bias (Faust and Leeper, 1997) or a lag-truncation bias (Chari et al., 2007).

The empirical work based on the use of sign restrictions, such as that of Scholl and Uhlig (2006) is not subject to the criticism of relying on arbitrary assumptions given that the identifying restrictions are derived from a theoretical model. In these studies, the identification of the shock of interest – a monetary policy shock – is achieved by imposing sign restrictions on impulse responses while being agnostic about the response of the key variable of interest – in this case, the exchange rate. However, as shown by Paustian (2007), the procedure used by Scholl and Uhlig (2006) does not guarantee that the identification of structural shocks is exact because multiple matrices define the linear mapping from orthogonal structural shocks to VAR residuals.

To more precisely estimate impulse responses, the sign restrictions method can be generalized to identify more than one shock (see Peersman, 2005, and Fry and Pagan, 2007). In this case, the estimation is more precise because the range of reasonable impulse responses is narrowed. To see this, note that when only one shock is identified – for example, a monetary policy shock – impulse responses that satisfy the sign restrictions of the monetary policy shock are accepted even if the responses to other shocks are unreasonable. This issue is avoided when more than one shock is identified because only the set of impulse responses that jointly satisfy all sign restrictions for all shocks is accepted. An example of this method is illustrated in Farrant and Peersman (2006), who analyze whether the real exchange rate is a shock absorber or a source of shocks for a series of dollar bilateral exchange rates using a sign-restriction method that identifies various shocks of interest. They find the real exchange rate is important as a shock absorber, mainly of demand shocks.

In an attempt to overcome the limitations previously mentioned, this paper analyzes the effects of productivity, demand, and monetary shocks on the real exchange rate. This approach represents a departure from most of the empirical literature which focuses on the estimation of monetary shocks that leaves potential sources of exchange rate fluctuations unexplained.[1]

In the baseline specification I estimate a VAR on quarterly data for the US *vis-à-vis* an aggregate of the rest of the world (ROW) and impose a stringent set of sign restrictions. I find that the contribution of monetary policy shocks to the variance of the real exchange rate ranges from 5% to 3% at 4- and 20-quarter horizons, respectively. By contrast, demand shocks explain 21% of the variance of the real exchange rate at a 4-quarter horizon and 37% at 20 quarters.

Sensitivity analysis suggests that my findings are robust to alternative sign restrictions. In particular, I experiment by leaving the response of the real exchange rate to a demand shock unrestricted. I also examine the robustness of my results to a subsample analysis and to different exchange rate measures. Overall, I find that the results are robust to these alternative specifications. I also check the sensitivity of my results to different estimation methods. When the identification strategy is based on a recursive approach, I find that monetary policy shocks explain only 3% of the movement of the real exchange rate at all horizons. Interestingly, this identification strategy yields a significant "puzzle", thus casting doubt on its validity. To evaluate the results using zero long-run restrictions I estimate the Clarida and Galí (1994) model using my data and sample period. I find that monetary shocks are unimportant in explaining exchange rate fluctuations and that demand shocks explain around 87% of exchange rate variance at 4- to 20-quarter horizons. The contribution of demand shocks imposing the zero long-run restrictions approach is significantly larger

[1]Important exceptions include the work of Clarida and Galí (1994) and generally the literature based on estimation using long-run restrictions. A recent contribution by Enders, Müller and Scholl (2008) that focuses on the role of fiscal policy on the real exchange rate also estimates a set of shocks.

than the sign-restriction approach. Indeed, this could be driven by an aggregation of multiple shocks (Faust and Leeper, 1997). These results suggest that model size and estimation method do matter.

The remainder of the paper is organized as follows. Section 2 outlines the theoretical model. Section 3 contains the empirical methodology based on a structural VAR framework with sign restrictions. The results of the baseline model are presented in Section 4, and I describe a battery of robustness tests in Section 5. Section 6 concludes.

2 Theoretical model

I estimate an sVAR using restrictions derived from the theoretical predictions of the model by Ferrero et al. (2008), which is a richly specified two-country DSGE model in the tradition of Obstfeld and Rogoff (2007).

The model consists of two countries, home (H) and foreign (F). Each economy has a representative household that behaves competitively and consumes tradable and nontradable consumption goods. Tradable consumption goods consist of both domestic and foreign-produced goods. Each economy also has a production sector for tradable goods and one for nontradable goods. Each sector has final and intermediate goods firms. Final goods firms behave competitively and produce an homogeneous good using differentiated intermediate goods with a CES production function. Intermediate goods firms are assumed to be monopolistic competitors and set prices on a staggered basis.

Financial markets are incomplete: Home bonds, denominated in home currency, are traded internationally but foreign bonds, denominated in foreign currency, are not. Monetary policy follows a feedback rule with interest rate smoothing.

Finally, each country's system is driven by three exogenous shocks: (i) tradable sector productivity, (ii) preference or demand, and (iii) monetary. In this paper, I refer interchangeably to "preference" shocks and "demand" shocks.

The following subsection outlines the loglinear approximation of the model around a deterministic steady state. A fully microfounded model that yields these equations can be found in Ferrero et al. (2008). I depart from their original formulation in three ways. First, the original model considers technology and preference shocks. My version here allows for tradable sector productivity, preference, and monetary shocks. The addition of a sector-specific shock and a monetary shock is straightforward. Second, for simplicity I assume that the productivity shock follows an autoregressive process rather than the combination of two autoregressive processes studied by Ferrero et al. (2008).[2] Finally, although the focus of Ferrero et al. (2008) is on current account dynamics, I am interested in the theoretical impulse responses of key macroeconomic variables to the three shocks described previously. I use these theoretical restrictions to derive empirical sign restrictions and consequently abstract from my analysis the study of current account adjustment.

2.1 Loglinear model

The world economy consists of two symmetric countries: home (H) and foreign (F). The derivation below is for the home country, and foreign country variables are represented with an asterisk.

Domestic output is a linear combination of home tradable and nontradable output denoted as

$$y_t = \gamma y_{Ht} + (1-\gamma)\, y_{Nt}, \tag{1}$$

where γ is the preference share for tradables.

The demand for home tradables can be expressed as

$$y_{Ht} = 2\alpha(1-\alpha)\eta\tau_t + (1-\gamma)\left[\alpha x_t + (1-\alpha)x_t^*\right] + \alpha c_t + (1-\alpha)c_t^*, \tag{2}$$

where α represents the home bias in tradables; η is the elasticity of substitution between home and foreign tradables; τ_t is the terms of trade defined as $p_{Ft} - p_{Ht}$; x_t and x_t^* are the relative prices of nontradable goods to tradable goods in the home and foreign countries[3] respectively; c_t is aggregate consumption in the home country; and c_t^* is aggregate consumption in the foreign country.

The demand for nontradables is given by

$$y_{Nt} = -\gamma x_t + c_t. \tag{3}$$

Aggregate consumption evolves according to the following intertemporal Euler equation:

$$c_t = E_t c_{t+1} - (i_t - E_t \pi_{t+1}) - \widehat{\beta}_t, \tag{4}$$

where $E_t c_{t+1}$ is expected future consumption, $i_t - E_t\pi_{t+1}$ is the real interest rate, and $\widehat{\beta}_t$ is the time-varying discount factor, which is defined as:

[2] This does not have an impact on the qualitative responses of the variable but affects the persistence of the productivity shock.

[3] More precisely, $x_t \equiv p_{Nt} - p_{Tt}$ and $x_t^* \equiv p_{Nt}^* - p_{Tt}^*$.

$$\widehat{\beta}_t = \varsigma_t - \psi \beta c_t, \tag{5}$$

where ς_t is a preference shock that follows an AR(1) process of the form

$$\varsigma_t = \rho_\varsigma \varsigma_{t-1} + u_{\varsigma t}, \tag{6}$$

$$u_{\varsigma t} \sim i.i.d.\ N(0, \sigma_\varsigma^2).$$

The evolution of the terms of trade is

$$\tau_t = \tau_{t-1} + \left(\pi_{Ft}^* - \pi_t^* - \Delta q_t\right) - \left(\pi_{Ht} - \pi_t\right), \tag{7}$$

where Δq_t is the rate of change of the real exchange rate, which is defined as $q_t = p_t - e_t - p_t^*$ (price of domestic goods in terms of foreign goods); e_t is the nominal exchange rate; π_{Ft}^* and π_{Ht} denote inflation in foreign and home tradables, respectively.

The relative price of nontradables to tradables evolves according to

$$x_t = x_{t-1} + \pi_{Nt} - \pi_{Ht} - \gamma(1-\alpha)\Delta\tau_t, \tag{8}$$

where π_{Nt} is home inflation in nontradables.

Let home inflation in tradables be

$$\pi_{Ht} = \kappa \left[(y_{Ht} - y_{Ht}^{\circ}) - \frac{1}{1+\varphi} \left(\mathrm{nx}_t - \mathrm{nx}_t^{\circ}\right) \right] + \beta E_t \pi_{H,t+1}. \tag{9}$$

The superscript $^{\circ}$ denotes the flexible price equilibrium of a variable, $y_{Ht}^{\circ} = a_t + \frac{1}{1+\varphi}\mathrm{nx}_t^{\circ}$, $\mathrm{nx}_t^{\circ} = 0$, $\kappa = \frac{(1-\xi)(1-\beta\xi)(1+\varphi)}{[\xi(1+\sigma\varphi)]}$, where φ is the inverse of the Frisch elasticity of labor supply, σ is the elasticity of substitution between intermediate inputs, ξ is the probability that a price does not adjust, and a_t denotes a tradable productivity shock given by

$$a_t = \rho_a a_{t-1} + u_{at}, \tag{10}$$

$$u_{at} \sim i.i.d.\ N(0, \sigma_a^2).$$

Inflation in nontradables is

$$\pi_{Nt} = \kappa \left(y_{Nt} - y_{Nt}^{\circ}\right) + \beta E_t \pi_{N,t+1}, \tag{11}$$

where $y_{Nt}^{\circ} = 0$.

CPI inflation depends on tradables inflation, nontradables inflation, and terms of trade inflation, denoted as

$$\pi_t = \gamma\pi_{Ht} + (1-\gamma)\pi_{Nt} + \gamma(1-\alpha)\Delta\tau_t. \tag{12}$$

The central bank sets nominal interest rates according to a feedback rule with interest rate smoothing:

$$i_t = \rho i_{t-1} + (1-\rho)\phi_\pi \pi_t + u_{mt}. \tag{13}$$

where ϕ_π is the long-run inflation semi-elasticity of short term interest rates and u_{mt} denotes a zero mean *i.i.d.* monetary policy shock.

Uncovered interest parity holds so the expected real exchange rate change must be offset by the real interest rate differential:

$$\left(i_t^* - E_t\pi_{t+1}^*\right) - \left(i_t - E_t\pi_{t+1}\right) = E_t q_{t+1} - q_t. \tag{14}$$

Net exports are defined as

$$\mathrm{nx}_t = \delta\left(\eta - 1\right)\tau_t + \sum_{s=0}^{\infty}\left(1-\alpha\right)E_t\widehat{\beta}_{Rt+s}, \tag{15}$$

where $\delta = 2\alpha\left(1-\alpha\right) > 0$ and $\widehat{\beta}_{Rt}$ is the difference between the home and foreign time varying discount factors.

Net foreign indebtness evolves according to

$$b_t = \frac{1}{\beta}b_{t-1} + \mathrm{nx}_t, \tag{16}$$

where b_t is debt normalized by trend output.

I am interested in the model predictions regarding the sign of the responses of key macroeconomic variables to tradable productivity, preference and monetary shocks. To compute the responses of a set of variables to these three shocks the model is calibrated. The following subsection outlines the parameter values used to simulate the model and describes the theoretical predictions.

2.2 Calibration

The calibration of the model follows Ferrero et al. (2008). Table 1 details the values of the parameters used to simulate the model.

[Insert Table 1 here]

The parameters displayed on the upper panel of the table are in line with earlier work with these type of models. However, the magnitude of the elasticity of substitution between home and foreign tradables η is controversial. Studies have estimated quite a range for this parameter and the uncertainty about its value is large: aggregate estimates are substantially lower than disaggregated estimates (see Bernard, Eaton, Jensen and Kotrum, 2003; and Hooper, Johnson and Marquez, 2000). I calibrate it to be equal to 2 in line with Obstfeld and Rogoff (2007) and Ferrero et al. (2008).[4]

The parameters on the lower panel describe the calibration of shocks. As it was noted by Corsetti, Dedola and Leduc (2008) and will be discussed below, the effects of a productivity shock are sensitive to the persistence of the autoregressive process. Thus, I calibrate the productivity shock for two cases: one in which a_t is a "near" unit root process (ρ_a=0.999) and another case in which it is less persistent ($\rho_a = 0.94$).[5] The preference shock ς_t intends to capture structural factors that influence the differences in consumption propensities across countries. It is a process that may persist over time. I set ρ_ς to equal 0.90 but the results remain robust when I assume that it is more persistent. The monetary policy shock is governed by the feedback coefficient ϕ_π and the smoothness parameter ρ. In line with Clarida, Galí and Gertler (1998), I set ϕ_π=2 and $\rho = 0.75$.

Figure 1 displays the theoretical responses of relative output $(y - y^*)$, relative consumption $(c - c^*)$, relative prices $(p - p^*)$, relative interest rates $(i - i^*)$, the real exchange rate $(reer)$, and the trade balance (tb) to tradables productivity, demand, and monetary shocks.[6] The shocks originate in the home country, but I focus on the effects on relative variables (home with respect to foreign) because I am interested in the behavior of the real exchange rate, which is a relative variable.

The predictions of the model are the following. In response to a tradables productivity shock (both in the case of a "near" unit root process and a less persistent shock), relative output and relative consumption increase. Relative prices decrease because the shock raises the supply of domestic goods and relative interest rates decrease. The productivity shock does not generate unambiguous responses on the real exchange rate. For a particular calibration, when the shock is highly persistent (ρ_a=0.999), the exchange rate appreciates. This result is driven by the increase in the relative price of nontraded goods. By contrast, when the shock is assumed to be

[4] Given the uncertainty regarding the appropriate parameter value and the key role that it may play in open economy models, I simulated the model for values ranging from 0.5 to 2. The simulations suggest that varying the elasticity from 0.5 to 2 does not have a major effect on the results.

[5] The response of some of the variables to a producitivy shock depends on the assumption of the persistence parameter. Thus, I experiment with these two calibrations to derive the sign restrictions.

[6] The shock's size is 1%.

less persistent (ρ_a=0.94), the increase in the relative price of nontraded goods is very small and as a consequence the real exchange rate depreciates. Overall, the effects of a productivity shock on the real exchange rate are ambiguous and depend on the calibration used.[7] The response of the trade balance is sensitive to the behavior of the real exchange rate. In the case of a highly persistent shock, the real exchange rate appreciates and the trade balance deteriorates. In the second case real exchange rate depreciates and the trade balance improves.

The effects of a preference or demand shock are in line with the textbook version of the Mundell-Fleming model: A demand shock leads to an increase in relative consumption and relative output. The increase in relative demand pushes relative prices higher, relative interest rates increase and the real exchange rate appreciates. The exchange rate appreciation and the rise in home demand both for domestic and foreign tradables generate a worsening of the trade balance.

Finally, an expansionary monetary shock induces a reduction in relative interest rates that leads to an increase in relative consumption, relative output, and relative prices. In addition, the interest rate decrease causes a depreciation of the real exchange rate given that UIP holds. Note that in the case of the monetary shock, two forces affect the trade balance in opposite directions. Home output increases as a result of the monetary expansion. As a consequence, import demand rises, which, *ceteris paribus*, would induce a deterioration of the trade balance; this is the *income-absorption effect*. At the same time, the exchange rate depreciation implies a reduction in the price of home tradables which generates an increase in the foreign demand of home tradables and an improvement of the trade balance; this is the *expenditure-switching effect*. Which of the two forces dominates depends on certain parameter values. In particular, the value of the elasticity of substitution between home and foreign tradables is crucial. In the calibration I assumed the value of 2 (as in Ferrero et al., 2008; and Obstfeld and Rogoff, 2007). With this parameter value the expenditure-switching effect dominates the income-absorption effect. Hence, the trade balance improves. [8]

[7] The response of the real exchange rate to a productivity shock is controversial in the literature. In a recent study, Corsetti, Dedola and Leduc (2008) show that for a particular calibration the real exchange rate appreciates after a productivity shock. The authors highlight that for an alternative calibration the model predicts the opposite.

[8] If I assume that η is lower (say, between 0.5 and 1), then the result would be a deterioration of the trade balance. The calibration results for alternative values of η are not presented here to preserve space but are available upon request.

3 Identification using short-run sign restrictions

3.1 Motivation

Researchers have noted that conventional methods to estimate VARs have a series of shortcomings. Estimation based on zero long-run restrictions, for example, suffers from distortions related to small-sample biases and measurement errors (Faust and Leeper, 1997). Chari et al. (2007) show that a lag-truncation bias can be present in VARs with long-run restrictions. This happens because the available data require a VAR with a small number of lags, which is a poor approximation of the infinite-order VAR of the observables from the model. In a related study, Christiano, Eichenbaum, and Vigfusson (2007) find that long-run identified VARs can be useful for discriminating among competing economic models.

Conventional methods involving zero short-run restrictions, such as the Choleski decomposition, have also been questioned on various grounds. Firstly, such restrictions are usually derived from some assumptions that may be difficult to reconcile with theoretical models. For example, the assumption of zero contemporaneous impact of a monetary policy shock on output is in contrast with some general equilibrium models (see Canova and Pina, 1999). Second, they sometimes yield counterintuitive impulse response functions of key endogenous variables that are not easily rationalized on the basis of conventional economic theory. An example is the so-called price puzzle, which refers to the increase in prices after a monetary tightening (see Sims and Zha, 2006; Christiano et al., 1999; Kim and Roubini, 2000). Third, as noted by Sarno and Thornton (2004), the results are often sensitive to the ordering of the variables.

To overcome the potential problems of the previous methods, I use an alternative identification procedure based on sign restrictions. Faust (1998), Canova and De Nicolò (2002), and Uhlig (2005) use sign restrictions to identify one shock– a monetary policy shock. Since I am interested in identifying a full set of shocks, I employ a methodology that extends the sign restriction approach to identify more than one shock. In particular, I apply the method described in Peersman (2005).

3.2 VAR model with sign restrictions

Consider the reduced-form VAR

$$Y_t = c + B(L)Y_{t-1} + A\epsilon_t, \tag{17}$$

where c is an $N \times 2$ matrix of constants and linear trends, Y_t is the $N \times 1$ vector of endogenous variables; $B(L)$ is a matrix polynomial in the lag operator L; ϵ_t is an $N \times 1$ vector of structural innovations; and $u_t = A\epsilon_t$ are the residuals. The six endogenous variables that I include in the VAR are the same as the ones analyzed in the previous section.

The usual problem of the VAR is the decomposition of the residuals u_t to obtain meaningful structural innovations. I identify the structural shocks using a sign-restriction approach. Because the shocks are assumed to be orthogonal, so that $E[\epsilon_t \epsilon_t'] = I$, the variance-covariance matrix of Eq. (17) is equal to $\Sigma = AA'$. For any orthogonal decomposition of A, we can find an infinite number of possible orthogonal decompositions of Σ, such that $\Sigma = AQQ'A'$, where Q is any orthonormal matrix ($QQ' = I$). A Choleski decomposition, for example, would assume a recursive structure on A so that A is a lower triangular matrix. Another candidate for A is the eigenvalue-eigenvector decomposition, $\Sigma = PDP' = AA'$, where P is a matrix of eigenvectors, D is a diagonal matrix of eigenvalues, and $A = PD^{1/2}$. This decomposition generates orthonormal shocks, making the value of P unique for each variance-covariance matrix decomposition without imposing zero restrictions. Following Canova and De Nicolò (2002), I consider $P = \prod_{m,n} Q_{m,n}(\theta)$, where $Q_{m,n}(\theta)$ is an orthonormal rotational matrix of the following form:

$$Q_{m,n} = \begin{bmatrix} 1 & 0 & ... & 0 & 0 \\ 0 & \cos(\theta) & ... & -\sin(\theta) & 0 \\ ... & ... & 1 & ... & ... \\ 0 & \sin(\theta) & ... & \cos(\theta) & 0 \\ 0 & 0 & ... & 0 & 1 \end{bmatrix} \tag{18}$$

where (m,n) indicates that the rows m and n are being rotated by the angle θ.

In a 6-variable model we have a 6×6 rotational matrix Q and 15 bivariate rotations.[9] The angles $\theta = \theta_1, ..., \theta_{15}$, and the rows m and n are rotated in Eq. (18).

My estimation is performed as follows. Firstly, all possible rotations are produced by varying the rotation angles θ in the range $[0, \pi]$. For practical purposes, I grid the interval $[0, \pi]$ into M points.[10] After estimating the coefficients of the $B(L)$ matrix using ordinary least squares (OLS), the impulse responses of N variables up to K horizons can be calculated for the contemporaneous impact matrix, $A_j (j = 1, ..., M^{15})$, as follows:

$$R_{j,t+k} = [I - B(L)]^{-1} A_j \epsilon_t \tag{19}$$

where $R_{j,t+k}$ is the matrix of impulse responses at horizon k. To identify the shock v of interest, sign restrictions can be imposed on $p \leqq n$ variables over the horizon

[9] In general terms, we have a total of $[N(N\text{-}1)]/2$, where N is the number of variables.

[10] In this case M=12, which implies 12^{15} possible rotations.

$0, ..., K$ in the following form:

$$R^{p,\, v}_{j,t+k} \lesseqgtr 0 \tag{20}$$

The sign restrictions are imposed based on the open-economy model of Ferrero et al. (2008). Table 2 summarizes the restrictions imposed on the data. Numbers on the table refer to the quarters for which the restrictions are binding and a question mark (?) denotes that the response of the variable is left unrestricted.[11]

Productivity shocks are identified by imposing that relative output and relative consumption do not fall for 4 quarters, relative prices do not increase for 4 quarters, and relative interest rates do not increase for 1 quarter. The response of the real exchange rate and the trade balance is left unrestricted.[12] Shocks to demand are identified by assuming that relative output, relative consumption, and relative prices do not decrease for 4 quarters, relative interest rates do not decrease for 1 quarter, the real exchange rate does not depreciate for 1 quarter, and no restriction is imposed on the response of the trade balance. To identify monetary policy shocks, I restrict the response of relative output, relative consumption, and relative prices to be nonnegative for 4 quarters, the response of the relative interest rate to be nonpositive for 1 quarter, and the exchange rate not to appreciate for 1 quarter.

[Insert Table 2 here]

Impulse responses are constructed using a Monte Carlo experiment. From all possible rotations (12^{15}), I select those that jointly satisfy the sign restrictions of the impulse responses for the three shocks. The restrictions imposed allow me to uniquely identify the three shocks. Solutions that satisfy all the restrictions are kept and the others are discarded. In practice, I repeat this procedure until 1000 draws that satisfy the restrictions are found. I show the median of the impulse responses and the 16th and 84th percentile error bands. The next section presents the results.

4 Empirical Results

4.1 Data

I use quarterly data over the period 1976-2007. The ROW series includes an aggregate of the other G7 countries (except the US).

[11] The response of the variables for which theory yields a range of predictions is left unrestricted.
[12] The theoretical model does not give unambiguous identifying restrictions for these two variables.

All the series are from the *International Financial Statistics* (IFS) of the International Monetary Fund (IMF). The data on real GDPs and real consumption are seasonally adjusted in local currencies at year 2000 price levels. I convert the GDP and consumption series in local currencies to US dollars using the average market exchange rate in 2000 (I do this to preserve consistency with the price's base year and to avoid mixing changes in real GDP with changes in the value of the dollar). As explained in the previous section, the log of US real GDP (y) and the log of US real consumption (c) are measured in deviation from the log GDP in the ROW (y^*) and the log of consumption in the ROW (c^*), respectively. y^* is the log of the sum of GDP in the other G7 countries and c^* is the log of the sum of consumption in the other G7 countries. The price series (p and p^*) are based on the CPI, and are presented in logs. Interest rates correspond to the US treasury bills (i) and an aggregate of 3-month money market rates for the other countries (i^*). The series p^* and i^* are calculated, respectively, as an average of prices and interest rates in the ROW weighted according to their respective (time-varying) GDP shares at purchasing power parity (PPP) values. The GDPs used for calculating the weights are at price levels and PPP values for the year 2000 and obtained from the OECD. The log of the real effective exchange rate ($reer$) corresponds to the REU series of the IFS. Finally, the US trade balance is expressed as a ratio of the GDP (tb). Figure 2 contains plots of the series (in levels) used in the VAR.

Table A1 in the appendix reports the results of the augmented Dickey-Fuller (ADF) and Kwiatkowski et al. (1992; KPSS) unit root tests. The ADF test fails to reject the unit root null hypothesis and the KPSS test rejects the stationary null for all the series except the interest rate in levels. By contrast, all variables show evidence of stationarity in first differences.[13]

Table A2 in the appendix shows the results of the Johansen (1991) test for the number of cointegrating vectors. According to the trace test, the null of no cointegration vectors cannot be rejected. Overall, these results suggest estimating the VARs in first differences.

4.2 Estimates of the baseline model

I now turn to the empirical findings by presenting the benchmark results from implementing the VAR described in Section 3.

Figure 3 shows the impulse responses of relative output, relative consumption, relative inflation, relative interest rates, the real exchange rate, and the trade balance to

[13] Note that the KPSS test rejects the null of stationarity for the first difference of relative consumption at the 10% level.

the three shocks of interest. The impulse responses suggest that a productivity shock generates a persistent increase in relative output and relative consumption. By contrast, relative prices and relative interest rates decline and the trade balance exhibits a continuous deterioration. The response of the real exchange rate is insignificant.

After a demand shock, there is a rise of around 0.4% in relative output and an increase in relative consumption. Relative prices go up by 0.1% on impact and relative interest rates increase persistently. The real exchange rate appreciates 1% on impact and it continues to rise. The trade balance worsens for about 8 quarters and thereafter the response is not statistically significant.

Finally, a monetary policy shock induces a decrease in relative interest rates and a temporary depreciation of the US dollar. After an initial depreciation of about 0.5%, the real exchange rate reaches its minimum value after 4 quarters and then reverts to equilibrium (consistent with PPP), showing no statistically significant reaction after 7 quarters. This result supports the delayed overshooting conclusion given that the peak is not immediate as predicted in Dornbusch (1976). A monetary shock also leads to a temporary positive effect on relative output and relative consumption. By contrast, relative prices exhibit a persistent rise.

Table 3 reports the variance decomposition of the real exchange rate. The results suggest that the contribution of monetary policy shocks to the variance of the real exchange rate is very small. Indeed, at a 4-quarter horizon the contribution of monetary policy shocks to real exchange rate fluctuations is 5% and at a 20- quarter horizon it is only 3%. By contrast, demand shocks explain a substantial proportion of the variance of the real exchange rate both at short and long horizons. Their contribution is 21% and 37% at horizons of 4 quarters and 20 quarters, respectively. The results also show that supply shocks play a moderate role in explaining real exchange rate fluctuations. In particular, I find that supply shocks explain 14%, 10%, and 9% of the movement in the real exchange rate at 4 quarters, 12 quarters, and 20 quarters, respectively.

[Insert Table 3 here]

In summary, when sign restrictions are used the findings indicate that demand shocks have been an important determinant of real exchange rate fluctuations both at short and long horizons. Supply shocks play a moderate role and monetary shocks are unimportant in explaining real exchange rate fluctuations.

5 Robustness and extensions

Empirical results often depend on modeling assumptions and variable definitions. Thus, in this section I assess the robustness of my results to different VAR specifications, variable definitions, and estimation methods.

5.1 Alternative sign restrictions

In this subsection I assess the robustness of my results to estimating the VAR without imposing a sign on the response of the real exchange rate to a demand shock. This allows the data to "speak" and assessment of whether the exchange rate responds in line with the theoretical model when it is left unrestricted.

Figure 4 compares the impulse responses of the model estimated using the baseline sign restrictions and shown in Figure 3 (solid lines) with the ones obtained using the alternative sign restrictions (hatched and dashed lines). Overall the impulse responses mirror those obtained for the baseline VAR. Interestingly, the response of the real exchange rate is only marginally modified.

Table 3 presents the variance decomposition of the real exchange rate using the alternative sign restrictions. Again, the results are in line with those of the baseline model. When the alternative sign restrictions are used, the contribution of productivity and demand shocks to explaining the real exchange rate variance is reduced.

5.2 Subsample analysis

Financial markets in the G7 countries have witnessed substantial changes over the sample period. For example, capital controls were gradually eliminated during the 1980s. These changes may have affected the way monetary policy shocks are transmitted into the economy. Thus, I divide the period into two subsamples (1976–1989 and 1990–2007) and estimate the impulse responses for each to check whether regime shifts change the results. The advantage of dividing the sample is that it avoids mixing periods with different structural characteristics. However, this comes with a cost. The estimation of the impulse responses is more likely to be imprecise and the shocks more difficult to detect. I choose 1990 as the date to split the two samples because it could be defined as the starting point for the recent wave of financial globalization.[14]

Impulse responses are shown in figures 5A (1976–1989) and 5B (1990-2007). Some interesting differences emerge when dividing the sample period. The effect of monetary shocks on the real exchange rate is larger in the second subsample. In fact, the

[14] Some authors have chosen 1982 as the split between subsamples (see, e.g., Kim, 1999, and Canova and De Nicoló, 2002). I do not analyze the results based on this break because the sample size becomes too small for the first subperiod.

contribution of monetary shocks to the exchange rate variance ranges from 10% to 4% at 4- quarter and 20- quarter horizons, respectively. By contrast, for the 1976-1989 period the contribution of monetary shocks is no larger than 4% for the entire forecast horizon. Interestingly, in the first subsample monetary shocks lead to an insignificant response of the trade balance, but in the second subsample the trade balance deteriorates. This implies that in the second subsample the expenditure-switching effect dominates.

5.3 Alternative exchange rate measures

I test for the sensitivity of the results by using the real effective exchange rate from the US Federal Reserve Board statistics instead of the one from the IFS. This index is CPI based and includes a wider set of countries. Figure 6 compares the impulse responses of the real effective exchange rate for the baseline model shown in Figure 3 (solid line) with the ones obtained using the alternative exchange rate measure (hatched and dashed lines). The impact of each shock on the real exchange rate is only marginally affected. In particular, the response of the real exchange rate is slightly attenuated when using the alternative real effective exchange rate index. The variance decomposition of the real effective exchange rate (not presented but available upon request) is very similar to that of the baseline specification.

In the analysis of my benchmark model, I used the real effective exchange rate of the United States *vis-à-vis* its trading partners while the ROW series are an aggregate of the other G7. One caveat is that the main trading partners of the United States are highly heterogeneous in many regards, in particular in their exchange rate regimes. About 50% of US trade today takes place with G6 economies and the other 50% with emerging market economies. While the former have had mostly freely floating exchange rates against the US dollar, the latter largely have had fixed or managed regimes over the sample period. Therefore, as a robustness check I estimate the VAR using the real effective exchange rate of the US dollar *vis-à-vis* the other G7 economies. Figure 7 shows the impulse response functions for the three structural shocks. The results are highly robust and little changed.

5.4 Other methods

To gain a further understanding of the sources of real exchange rate fluctuations, it is informative to identify the shocks using other methods. In particular, I examine the impact of monetary shocks using the Choleski decomposition in the same fashion as Eichenbaum and Evans (1995) and zero long-run restrictions as in Clarida and Galí (1994).

5.4.1 Choleski decomposition

Figure 8 shows the impulse responses using the Choleski decomposition. I identify the monetary policy shock with innovations in the interest rate differential. The order of the variables in the VAR is the same as in the figure.

The results show that in response to a monetary expansion all the variables except prices respond in line with the predictions of the model. In particular, the real exchange rate exhibits a temporary depreciation, and relative output and relative consumption increase. Interestingly, prices decrease for nine quarters and increase afterward. This response of prices to a monetary expansion resembles the so-called price puzzle noted by Sims (1992).

In terms of variance decomposition, one point to highlight is that according to the recursive approach, monetary policy shocks explain only around 3% of the movement of the real exchange rate at all horizons. These results are in line with those of the baseline model.

5.5 Clarida and Galí analysis revisited

In this subsection I estimate the standard 3-variable VAR of Clarida and Galí (1994). Their VAR model contains $\Delta(y - y^*)$, $\Delta reer$, and $\Delta(p - p^*)$ and the disturbances consist of supply, demand, and monetary shocks. Based on the theoretical restrictions of the Clarida-Galí model I use two identification strategies. The first consists of estimating the VAR applying the zero long-run restrictions approach popularized by Blanchard and Quah (1989). This replicates the results of Clarida and Galí on my data and sample period.[15] In addition, I estimate the 3-variable VAR using the sign-restriction method. Interestingly, the short-run predictions of the Clarida-Galí model match those of the Ferrero et al. (2008) DSGE model. Note that the supply shock of the Clarida-Galí model can be understood as the productivity shock analyzed before. Hence, I estimate the baseline VAR for a subset of variables.

Impulse responses are presented in Figures 9A (long-run restrictions) and 9B (sign restrictions) and Table 4 shows the variance decomposition. In line with the findings of Clarida and Galí (1994), using long-run restrictions I find that monetary shocks are unimportant and demand shocks explain almost all of the variance of the real exchange rate. By contrast, using sign restrictions on the 3-variable VAR I find that although monetary policy shocks are not the main drivers of the real exchange rate,

[15] The model is lower triangular in the long run. The restrictions are based on the predictions of the model, which are that in the long run (i) only supply shocks lead to increases in the level of relative output, (ii) supply and demand shocks have an impact on the long-run level of the real exchange rate, and (iii) the three shocks have an impact on relative prices in the long run.

they are nevertheless important. The range for their short-horizon contribution is 38% to 48%.

There is some contrast among the baseline sign restrictions, the Clarida-Galí approach and the 3-variable sign restriction results. Both the baseline sign restrictions in Section 4 and long-run restrictions imply that monetary policy shocks account for very little of the exchange rate fluctuations. However, it is relevant to emphasize one difference between these two estimations. Long-run restrictions imply that demand shocks account for 87% of real exchange rate variance, but when the baseline sign restrictions are used the contribution ranges from 21% to 37%. This difference is somewhat explained by construction. Note that VARs estimated with the zero long-run restrictions method are exactly identified. Hence, in a 3-variable model with 3 shocks, 100 % of the variance decomposition is explained by the shocks of the model. A potential drawback of this method is that the shocks identified may be the result of a multiple aggregation of shocks, as discussed by Faust and Leeper (1997), which could lead to an estimation bias.[16] Thus, model size may matter. By contrast, the baseline sign-restriction method does not provide an exact identification scheme and allows for the presence of unexplained shocks (i.e., shocks that have a pattern that does not match the sign restrictions imposed, originated, for example, in risk premia or other shocks). The problem of the multiple aggregation of shocks still remains in the 3-variable VAR model estimated with sign restrictions. Some of the impulse responses that are accepted as demand or monetary shocks may, in fact, contradict the responses of relative interest rates, relative consumption, or the trade balance. As a consequence, model size and estimation method matter.[17]

[Insert Table 4 here]

6 Conclusion

The explanation of the sources of real exchange rate fluctuations is still an open area. There has been a widespread belief that monetary policy is the main driver of exchange rate movements. Much of the theoretical literature has focused on confirming this belief. However, the empirical evidence on the role of monetary policy

[16] An example of the multiple aggregation of shocks is described in Rogers (1999).

[17] Ideally, the researcher would like to estimate a VAR with more variables. However, as more variables are added, the researcher faces the well-known problem of the course of dimensionality. In addition, the estimation outcome becomes more imprecise with a higher-order VAR using sign restrictions. This happens because the shocks are more difficult to identify. The 6-variable VAR may not be capturing some important aspect of the economy analyzed but seems to provide a good summary of the key macroeconomic variables.

shocks has not provided clear-cut answers on the link between monetary policy and exchange rate movements. In addition, such work often has been criticized for its lack of credible identifying assumptions.

This paper has focused on one specific question: How important are real and nominal shocks as drivers of the US real exchange rate? To address this question, I begin by analyzing the effects of productivity, demand, and monetary shocks on a set of macroeconomic variables using a two-country DSGE model. The predictions of this model are used to derive empirical sign restrictions that are then applied to the estimation of a VAR model. I find that real shocks play a dominant role as drivers of exchange rate fluctuations and that monetary shocks are unimportant. This conclusion is robust to a battery of sensitivity tests.

These findings have important implications. First, they reveal that, in contrast to results based on zero long-run restrictions, the contribution of demand shocks plays a key role but not of the order of magnitude sometimes found in the earlier literature. This implies that the aggregation of multiple shocks can have an important effect. Second, the finding that monetary policy shocks are unimportant suggests that the recent focus of the literature on real shocks to match the empirical properties of real exchange rates is well founded. A next step could be to examine the potential contribution of risk premia as a driver of exchange rate fluctuations.

A Appendix

Table A1. Tests for unit roots

Test	$y-y^*$	$c-c^*$	$p-p^*$	$i-i^*$	q	tb
	Levels					
ADF^{AIC}	-1.33	-0.51	0.87	-2.75^*	-1.41	-1.00
	(0.618)	(0.884)	(0.995)	(0.069)	(0.577)	(0.753)
ADF^{BIC}	-0.24	1.21	0.41	-2.86^{**}	-1.05	-0.89
	(0.929)	(0.998)	(0.983)	(0.053)	(0.734)	(0.788)
KPSS	0.49^{**}	0.88^{***}	1.31^{***}	0.22	0.76^{***}	1.023^{***}
	First Differences					
ADF^{AIC}	-3.50^{***}	-3.13^{**}	-5.96^{***}	-4.14^{***}	-8.69^{***}	-3.87^{***}
	(0.009)	(0.027)	(0.000)	(0.001)	(0.000)	(0.003)
ADF^{BIC}	-8.39^{***}	-5.92^{***}	-4.07^{***}	-9.30^{***}	-8.69^{***}	-9.45^{***}
	(0.000)	(0.000)	(0.002)	(0.000)	(0.000)	(0.000)
KPSS	0.31	0.37^*	0.14	0.04	0.08	0.09

Notes: The table shows the augmented Dickey-Fuller (ADF) and the Kwiatkowski et al. (1992) (KPSS) test statistics. The former tests the null of unit root against a stationary alternative. The latter tests the null of stationarity. The critical values of the ADF test are -2.58, -2.88, and -3.48 for the 10%, 5%, and 1% significance levels respectively. The critical values for the KPSS test are 0.74, 0.46, and 0.35 for the 10%, 5% and 1% significance levels respectively. AIC denotes that the lag length was selected according to the Akaike infromation criterion and BIC denotes that it was selected based on the Schwartz criterion. p-values are listed in parenthesis. The sample period is 1976-2007. *, **, *** indicate rejection of the null at 10%, 5%, and 1% levels respectively.

Table A2. Test of cointegrating rank

Rank=r	Trace	95% critical value	p-value
	Baseline Model		
r=0	106.221	117.451	0.212
r=1	60.355	88.554	0.844
r=2	47.102	63.659	0.551
r=3	28.270	42.770	0.610
r=4	15.643	25.731	0.529
r=5	5.272	12.448	0.566

Notes: The table shows the trace statistic corresponding to the Johansen (1991) test for the number of cointegrating vectors. The statistics apply a small-sample correction. The sample period is 1976-2007. The VAR model is estimated with 4 lags.

References

Beaudry P. and Devereux M.B., 1995. Money and the real exchange rate with sticky prices and increasing returns. Carnegie-Rochester Conference Series on Public Policy 43, 55-102.

Bernard, A.B., Eaton, J., Jensen, J.B., and Kotrum, S., 2003. Plants and Productivity in International Trade. American Economic Review 93, 1268-1290.

Blanchard, O. and Quah, D., 1989. The Dynamic Effects of Aggregate Demand and Supply Disturbances. American Economic Review 79, 655-673.

Canova, F. and De Nicolò , G., 2002. Monetary Disturbances Matter for Business Fluctuations in the G-7. Journal of Monetary Economics 49, 1131-1159.

Canova, F. and Pina J., 1999. Monetary Policy Misspecification in VAR Models. CEPR Discussion Paper No 2333.

Chari, V., P., Kehoe, P. and McGrattan, E., 2002. Can Sticky Price Models Generate Volatile and Persistent Real Exchange Rates? Review of Economic Studies 69, 533-563.

Chari, V., Kehoe, P. and McGrattan, E., 2007. Are Structural VARs with Long-Run Restrictions Useful in Developing Business Cycle Theory? Staff Report 364, Federal Reserve Bank of Minneapolis .

Christiano, L.J., Eichenbaum, M. and Evans, C.L., 1999. Monetary Policy Shocks: What Have We Learned and To What End? In Taylor, J.B., Woodford, M. (Eds), Handbook of Macroeconomics, Vol. 1A, Ch. 2, Elsevier, Amsterdam, 65-148.

Christiano, L.J., Eichenbaum, M. and Vigfusson, M., 2007. Assessing Structural VARs. In Acemoglu, D., Rogoff, K., Woodford, M. (Eds.), NBER Macroeconomics Annual 2006, Volume 21, 1-72.

Clarida, R. and Galí, J., 1994. Sources of Real Exchange-Rate Fluctuations: How Important are Nominal Shocks? Carnegie-Rochester Conference Series on Public Policy 41, 1-56.

Clarida, R., Galí, J. and Gertler, M., 1998. Monetary policy rules in practice Some international evidence. European Economic Review 42, 1033-1067.

Corsetti, G., Dedola, L. and Leduc, S. 2007. International Risk Sharing and the Transmission of Producitvity Shocks. Review of Economic Studies 75, 443-473.

Dornbusch, R., 1976. Expectations and Exchange Rate Dynamics. Journal of Political Economy 84, 1161-76.

Eichenbaum, M. and Evans., C.L., 1995. Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates. Quarterly Journal of Economics 110, 975-1009.

Enders, Z., Müller, G.J. and Scholl, A., 2008. How Do Fiscal and Technology Shocks Affect Real Exchange Rates? New Evidence for the United States. CFS Working Paper Series 22.

Farrant K. and Peersman, G., 2006. Is the Exchange Rate a Shock Absorber or Source of Shocks? New Empirical Evidence. Journal of Money, Credit and Banking 38, 939-962.

Faust, J., 1998. The Robustness of Identified VAR Conclusions About Money. Carnegie-Rochester Conference Series in Public Policy 49, 207-244.

Faust, J. and Leeper, E.M., 1997. When Do Long-Run Identifying Restrictions Give Reliable Results? Journal of Business and Economic Statistics 15, 345-353.

Faust, J. and Rogers J.H., 2003. Monetary Policy's Role in Exchange Rate Behavior. Journal of Monetary Economics 50, 1403-1424.

Ferrero, A., Gertler, M. and Svensson, L., 2008. Current Account Dynamics and Monetary Policy. NBER Working Paper 13906.

Fry R. and Pagan, A., 2007. Some Issues in Using Sign Restrictions for Identifying Structural VARs. NCER Working Paper Series 14, National Centre for Econometric Research.

Hooper, P., Johnson, K and Marquez, J., Trade Elasticities for the G-7 Countries. Princeton Studies in International Economics No. 87.

Johansen, S., 1991. Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. Econometrica 59, 1551-1580.

Kim, S., 1999. Do Monetary Policy Shocks Matter in the G-7 Countries? Using Common Identifying Assumptions about Monetary Policy Across Countries. Journal of International Economics 48, 387–412.

Kim, S. and Roubini, N., 2000. Exchange Rate Anomalies in the Industrial Countries: A Solution with a Structural VAR Approach. Journal of Monetary Economics 45, 561-586.

Kwiatkowski, D., Phillips, P.C.B, Schmidt, P. and Shin, Y., 1992. Testing the Null Hypothesis of Stationarity Against the Alternative of Unit Root: How Sure Are We that Economic Time Series Have a Unit Root? Journal of Econometrics 54, 159-178.

Obstfeld, M. and Rogoff, K., 2007. The Unsustainable U.S. Current Account Position Revisited. In R. Clarida (Ed.), G7 Current Account Imbalances: Sustainability and Adjustment. Chicago: Chicago University Press.

Paustian, M., 2007. Assessing Sign Restrictions. B.E. Journal of Macroeconomics 7, 1-31.

Peersman, G., 2005. What Caused the Early Millennium Slowdown? Evidence Based on Vector Autoregressions. Journal of Applied Econometrics 20, 185-207.

Rogers, J. H., 1999. Monetary Shocks and Real Exchange Rates. Journal of International Economics 49, 269-288.

Rogoff, K., 1996. The Purchasing Power Parity Puzzle. Journal of Economic Literature 34, 647-668.

Sarno, L. and Thornton, D.L., 2004. The Efficient Market Hypothesis and Identification in Structural VARs. Federal Reserve Bank of St. Louis Review 86, 49-60.

Scholl, A. and Uhlig, H., 2006. New Evidence on the Puzzles. Results from Agnostic Identification on Monetary Policy and Exchange Rates. SFB 649 Discussion Papers, Humboldt University Berlin.

Sims, C.A., 1992. Interpreting the Macroeconomic Time Series Facts: the Effects of Monetary Policy. European Economic Review 36, 975–1011.

Sims, C.A., Zha, T., 2006. Does Monetary Policy Generate Recessions? Macroeconomic Dynamics 10, 231-272.

Steinsson, J., 2008. The Dynamic Behavior of the Real Exchange Rate in Sticky Price Models. American Economic Review 98, 519-533.

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.

Table 1. Parameter Values

Model	
Preference share for tradables	γ =0.25
Preference share for home tradables	α =0.7
Elasticity of substitution between home and foreign tradables	η =2
Elasticity of substitution between intermediate inputs	σ =11
Steady-state discount factor	β =0.99
Spillover effect of aggregate consumption on discount factor	ψ =7.2361$\times10^{-6}$
Inverse of Frisch elasticity of labor supply	φ =2
Probability that the price does not adjust	ξ =2/3
Productivity shock	
Productivity persistence	ρ_a =[0.94;0.999]
Preference shock	
Preference shock persistence	ρ_ς =0.90
Monetary shock	
Inflation elasticity of interest rate (feedback coefficient)	ϕ_π =2
Interest rate smoothing	ρ =0.75

Table 2. Sign Restrictions: Baseline VAR

Shock	$y-y^*$	$c-c^*$	$p-p^*$	$i-i^*$	$reer$	tb
Productivity	$\uparrow 1-4$	$\uparrow 1-4$	$\downarrow 1-4$	$\downarrow 1$	?	?
Demand	$\uparrow 1-4$	$\uparrow 1-4$	$\uparrow 1-4$	$\uparrow 1$	$\uparrow 1$	?
Monetary	$\uparrow 1-4$	$\uparrow 1-4$	$\uparrow 1-4$	$\downarrow 1$	$\downarrow 1$	?

Table 3. Variance Decomposition of the Real Effective Exchange Rate

(Sign restrictions)

		Shocks		
Horizons		Productivity	Demand	Monetary
4 quarters	Baseline	0.14 [0.02 ; 0.43]	0.21 [0.06 ; 0.50]	0.05 [0.01 ; 0.19]
	Alternative sign	0.12 [0.02 ; 0.41]	0.12 [0.02 ; 0.43]	0.04 [0.01 ; 0.17]
	1976-1989	0.10 [0.02 ; 0.34]	0.25 [0.10 ; 0.51]	0.04 [0.01 ; 0.15]
	1990-2007	0.10 [0.02 ; 0.33]	0.18 [0.04 ; 0.45]	0.10 [0.02 ; 0.34]
8 quarters	Baseline	0.11 [0.02 ; 0.36]	0.29 [0.10 ; 0.59]	0.04 [0.01 ; 0.16]
	Alternative sign	0.11 [0.01 ; 0.35]	0.19 [0.04 ; 0.49]	0.03 [0.01 ; 0.14]
	1976-1989	0.08 [0.02 ; 0.28]	0.36 [0.14 ; 0.60]	0.03 [0.01 ; 0.11]
	1990-2007	0.11 [0.02 ; 0.33]	0.23 [0.07 ; 0.51]	0.08 [0.02 ; 0.28]
12 quarters	Baseline	0.10 [0.01 ; 0.35]	0.33 [0.13 ; 0.62]	0.03 [0.01 ; 0.14]
	Alternative sign	0.10 [0.01 ; 0.34]	0.23 [0.05 ; 0.55]	0.03 [0.01 ; 0.12]
	1976-1989	0.07 [0.02 ; 0.27]	0.40 [0.14 ; 0.59]	0.03 [0.01 ; 0.11]
	1990-2007	0.11 [0.02 ; 0.34]	0.28 [0.09 ; 0.54]	0.06 [0.01 ; 0.22]
16 quarters	Baseline	0.09 [0.01 ; 0.35]	0.35 [0.14 ; 0.64]	0.03 [0.01 ; 0.12]
	Alternative sign	0.10 [0.01 ; 0.35]	0.25 [0.06 ; 0.58]	0.03 [0.01 ; 0.11]
	1976-1989	0.06 [0.02 ; 0.25]	0.39 [0.14 ; 0.59]	0.03 [0.01 ; 0.11]
	1990-2007	0.12 [0.02 ; 0.33]	0.31 [0.10 ; 0.56]	0.05 [0.01 ; 0.17]
20 quarters	Baseline	0.09 [0.01 ; 0.34]	0.37 [0.14 ; 0.65]	0.03 [0.01 ; 0.11]
	Alternative sign	0.09 [0.01 ; 0.33]	0.26 [0.05 ; 0.60]	0.03 [0.01 ; 0.11]
	1976-1989	0.06 [0.02 ; 0.25]	0.40 [0.17 ; 0.62]	0.03 [0.01 ; 0.12]
	1990-2007	0.12 [0.02 ; 0.34]	0.32 [0.11 ; 0.55]	0.04 [0.01 ; 0.16]

Notes: The table shows the percentage of the error variance of the real effective exchange rate due to each shock at 4-, 8-, 12-, 16-, and 20- quarters horizon. The lag length is 3. The 16th and 84th percentile error bands are listed in brackets.

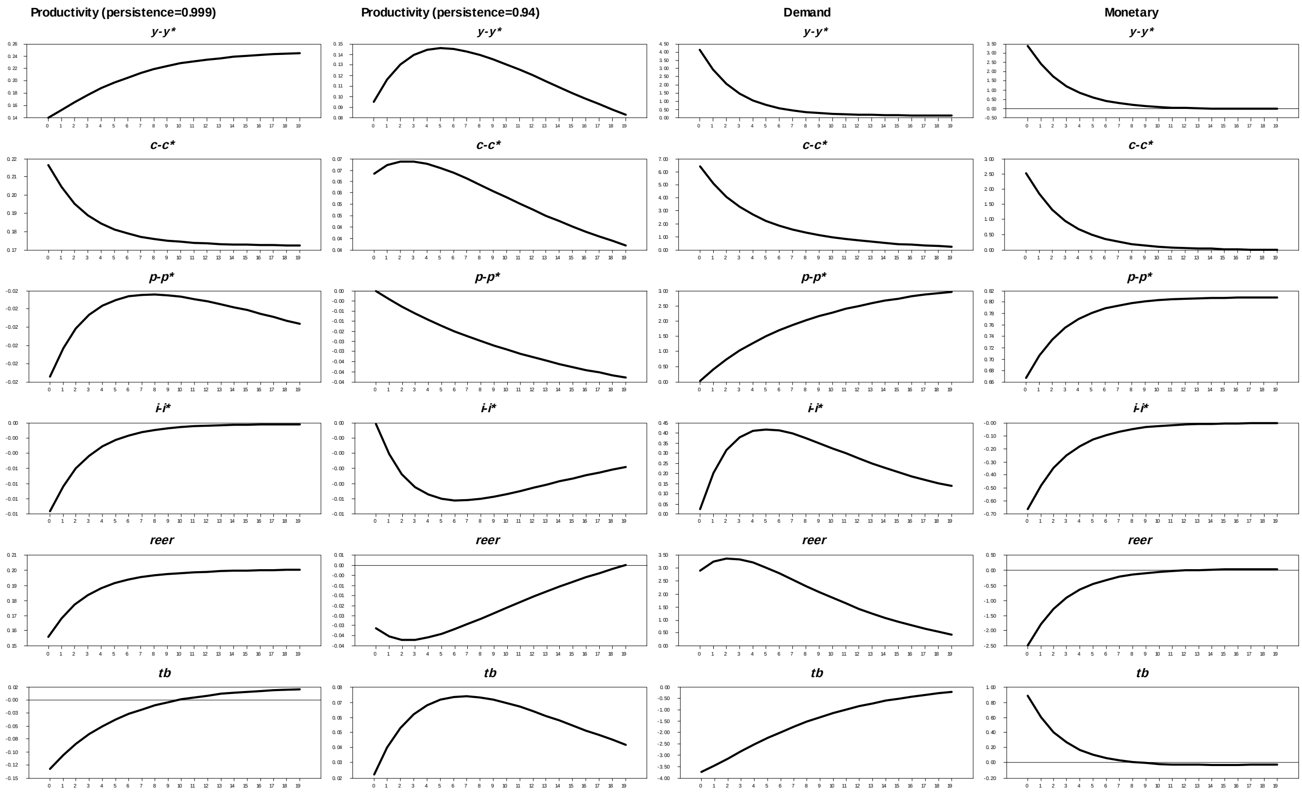
Table 4. Variance Decomposition of the Real Effective Exchange Rate

(3-variable VAR)

		Shocks		
Horizons		Supply	Demand	Monetary
4 quarters	CG	0.05 [0.01 ; 0.15]	0.86 [0.73 ; 0.94]	0.07 [0.02 ; 0.15]
	SR	0.04 [0.01 ; 0.13]	0.45 [0.16 ; 0.74]	0.48 [0.19 ; 0.76]
8 quarters	CG	0.09 [0.02 ; 0.20]	0.87 [0.76 ; 0.95]	0.03 [0.01 ; 0.06]
	SR	0.05 [0.01 ; 0.15]	0.54 [0.24 ; 0.80]	0.38 [0.12 ; 0.68]
12 quarters	CG	0.10 [0.02 ; 0.24]	0.86 [0.74 ; 0.96]	0.02 [0.01 ; 0.04]
	SR	0.05 [0.01 ; 0.18]	0.58 [0.26 ; 0.83]	0.32 [0.08 ; 0.64]
16 quarters	CG	0.11 [0.02 ; 0.27]	0.87 [0.71 ; 0.96]	0.01 [0.00 ; 0.03]
	SR	0.06 [0.01 ; 0.19]	0.60 [0.28 ; 0.84]	0.29 [0.07 ; 0.62]
20 quarters	CG	0.12 [0.02 ; 0.28]	0.86 [0.70 ; 0.96]	0.01 [0.00 ; 0.02]
	SR	0.06 [0.01 ; 0.20]	0.61 [0.22 ; 0.84]	0.27 [0.06 ; 0.61]

Notes: The table shows the percentage of the error variance of the real effective exchange rate due to each shock at 4-, 8-, 12-, 16-, and 20- quarters horizon using the Clarida-Galí (CG) and sign-restriction (SR) methods. The lag length is 3. The 16th and 84th percentile error bands are listed in brackets.

Figure 1. Theoretical Impulse Responses



Notes: The figure shows the theoretical impulse responses to productivity, demand, and monetary shocks derived from the DSGE model presented in Section 2.

Figure 2. Data

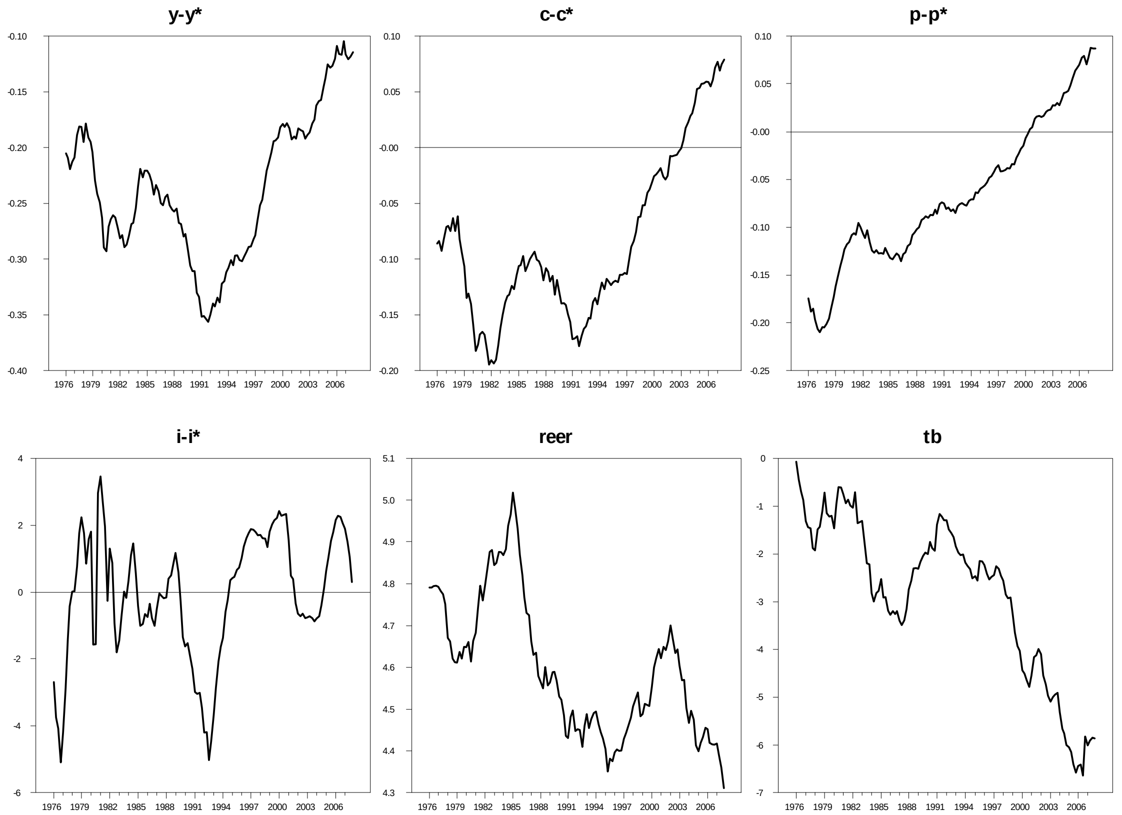
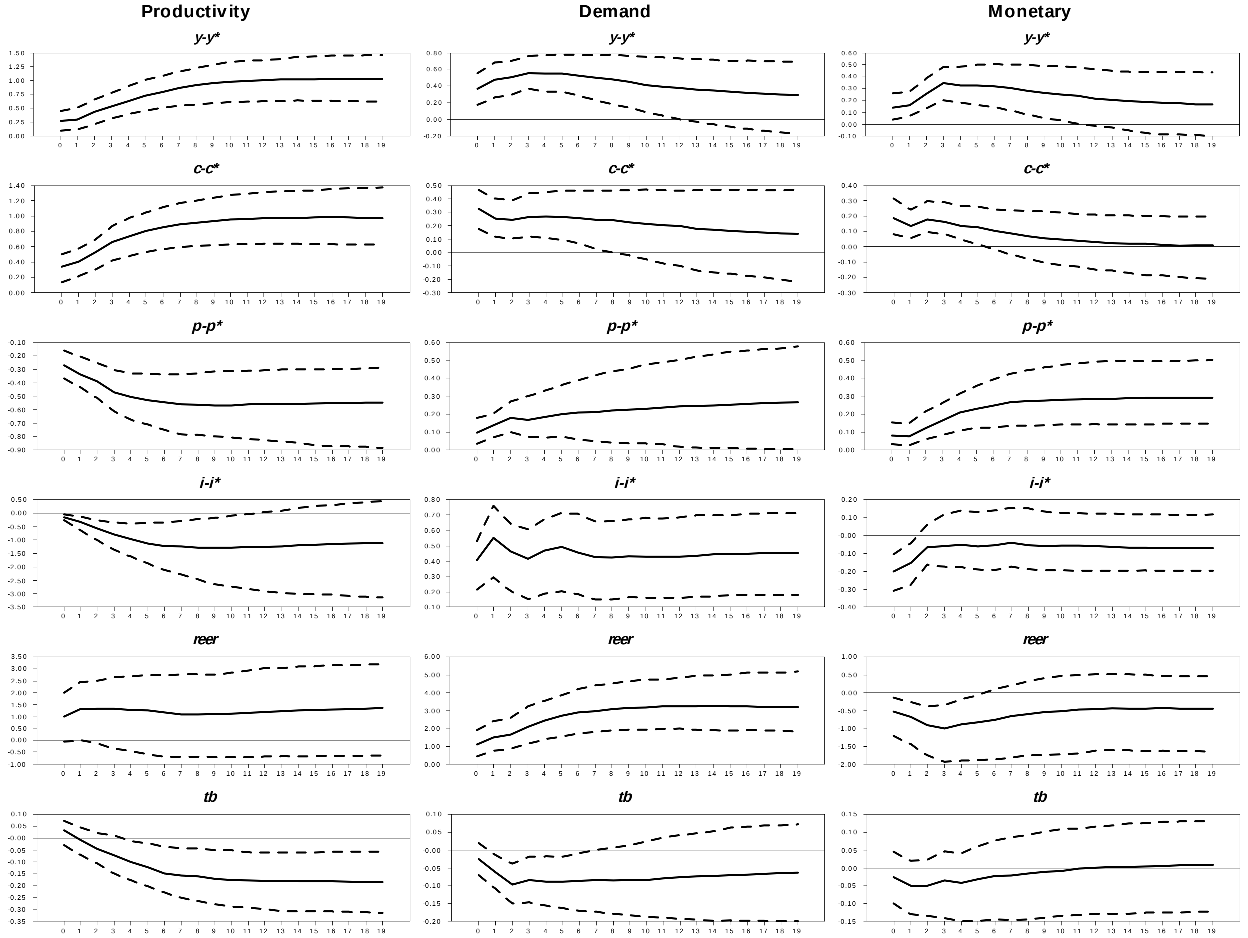
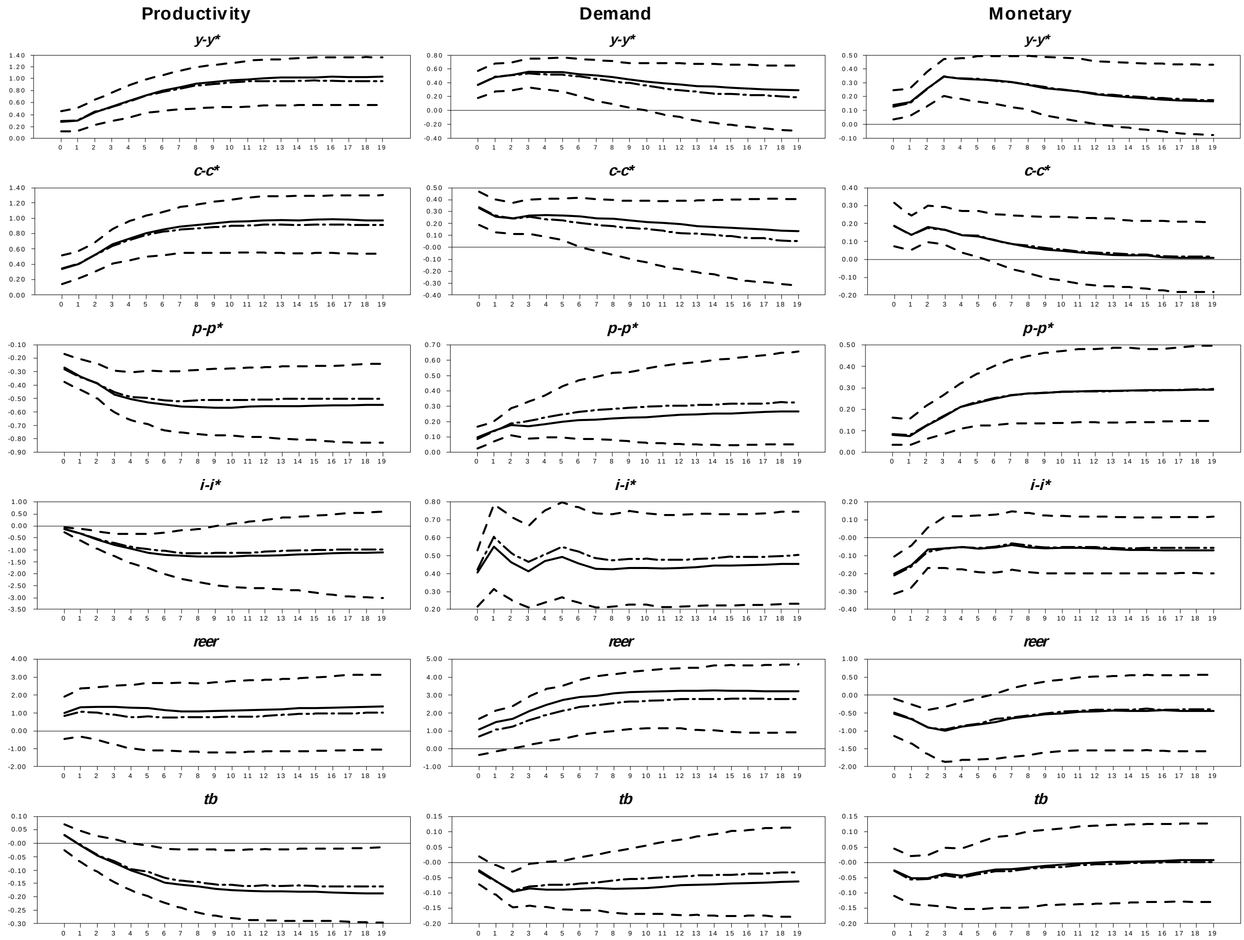


Figure 3. Impulse Responses: Baseline Model



Notes: The figure shows the impulse responses to productivity, demand, and monetary shocks using sign restrictions. The solid lines are the median impulse responses and dashed lines represent the 16th and 84th percentile error bands.

Figure 4. Impulse Responses: Alternative Sign Restrictions



Notes: The figure compares the impulse responses to productivity, demand, and monetary shocks using the baseline sign restrictions of Figure 3 (solid lines) with the ones obtained using alternative sign restrictions (hatched and dashed lines). The alternative specification relaxes the restriction on the real exchange rate in the case of productivity and demand shocks.

Figure 5A. Impulse Responses: Subsample 1976–1989

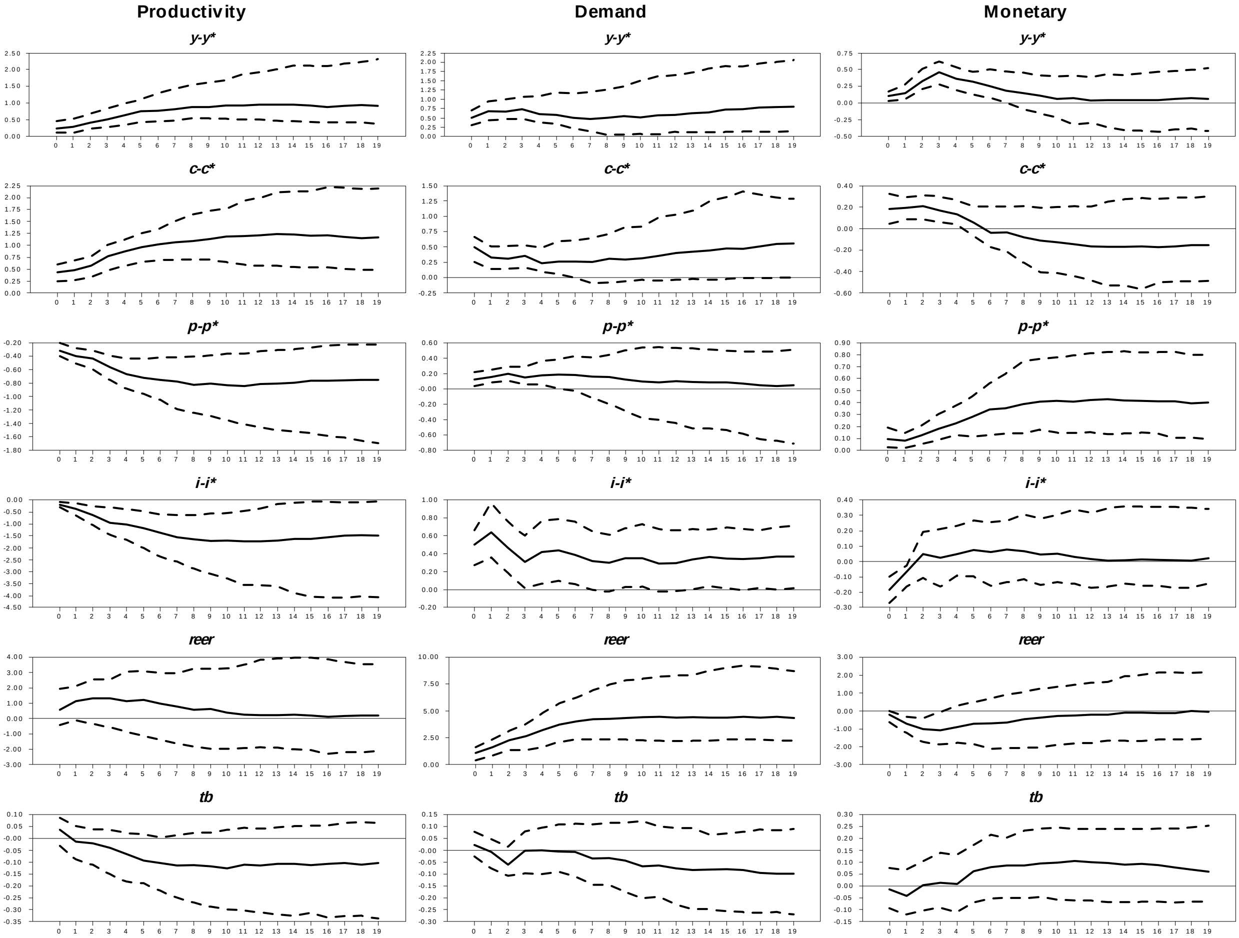
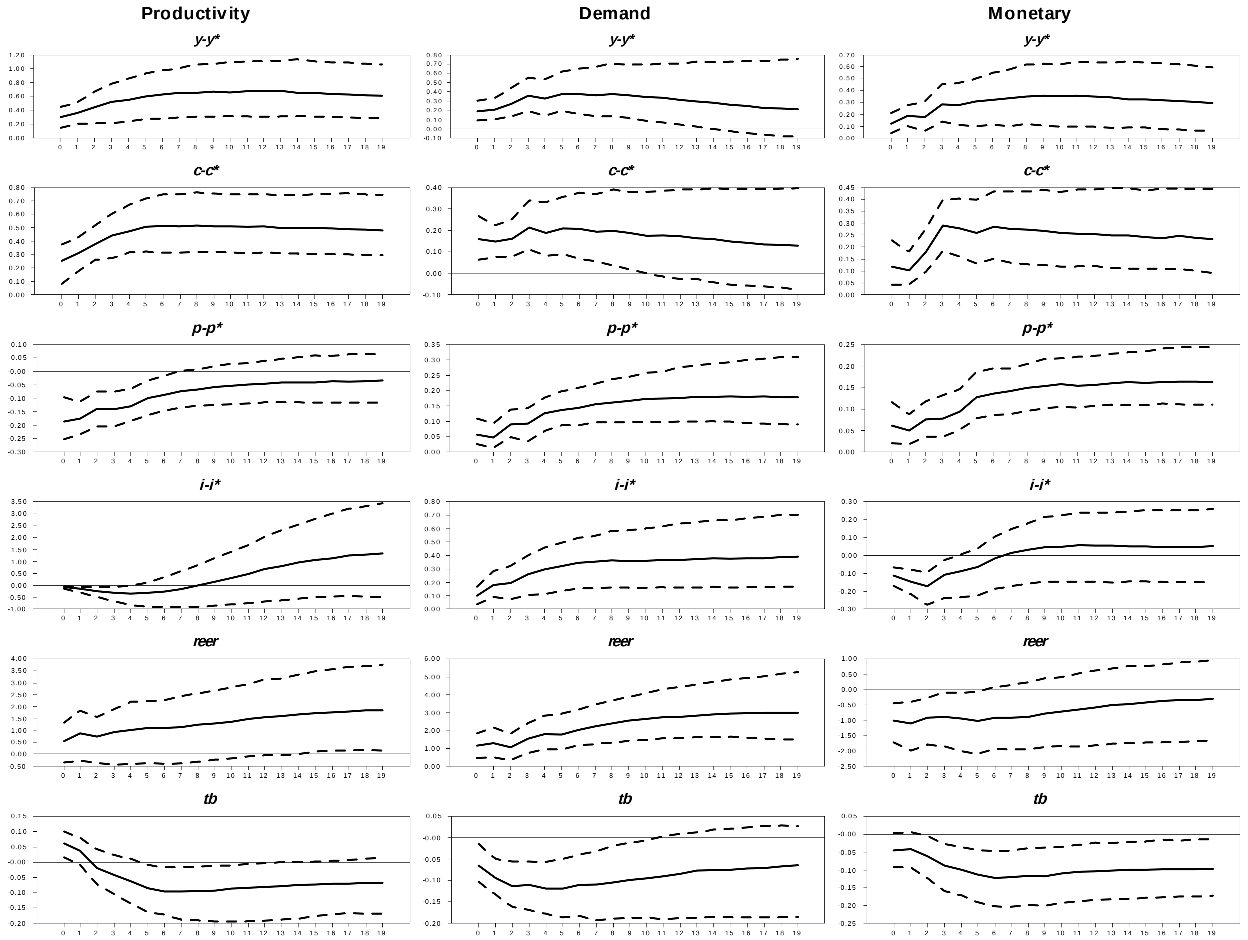
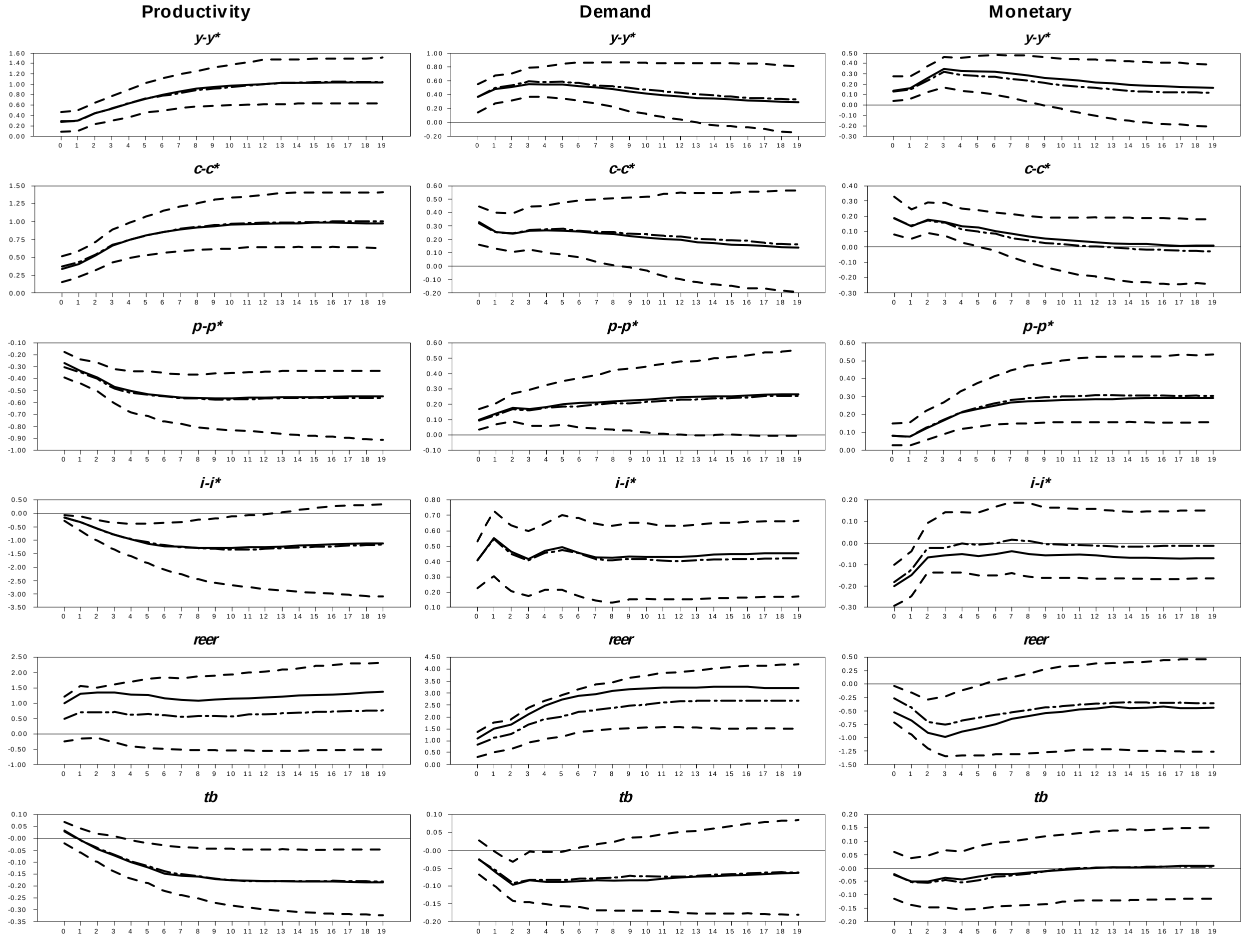


Figure 5B. Impulse Responses: Subsample 1990–2007



Notes: Figures 5A. and 5B. show the impulse responses to productivity, demand, and monetary shocks using sign restrictions for two subperiods. The solid lines are the median impulse responses and dashed lines represent the 16th and 84th percentile error bands.

Figure 6. Impulse Responses: Alternative REER



Notes: The figure compares the impulse responses of the real exchange rate to productivity, demand, and monetary shocks using the baseline model presented in Figure 3 (solid lines) with the ones obtained when the model is estimated using the real effective exchange rate with data from the Federal Reserve Board of Governors (hatched and dashed lines).

Figure 7 . Impulse Responses: REER G6

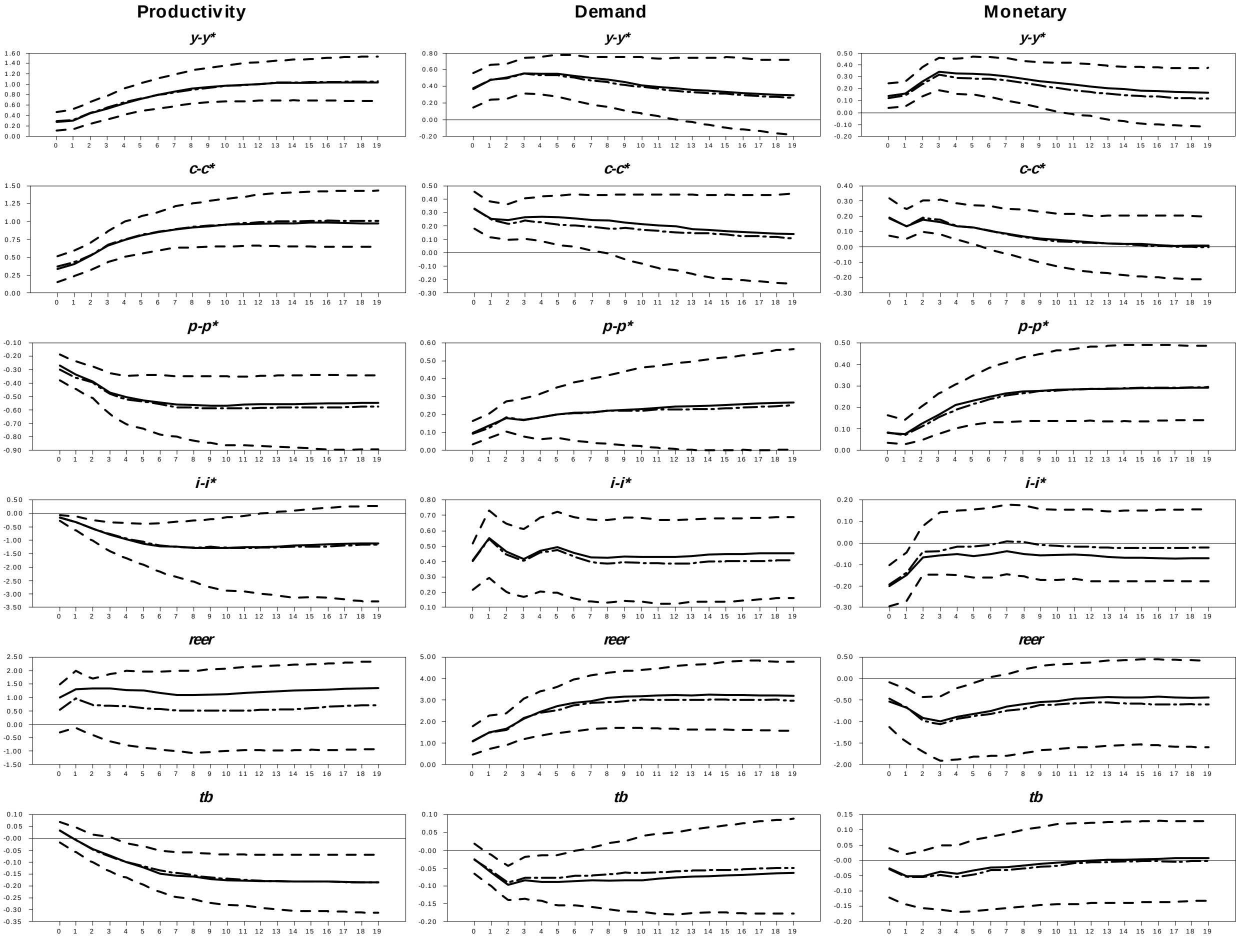
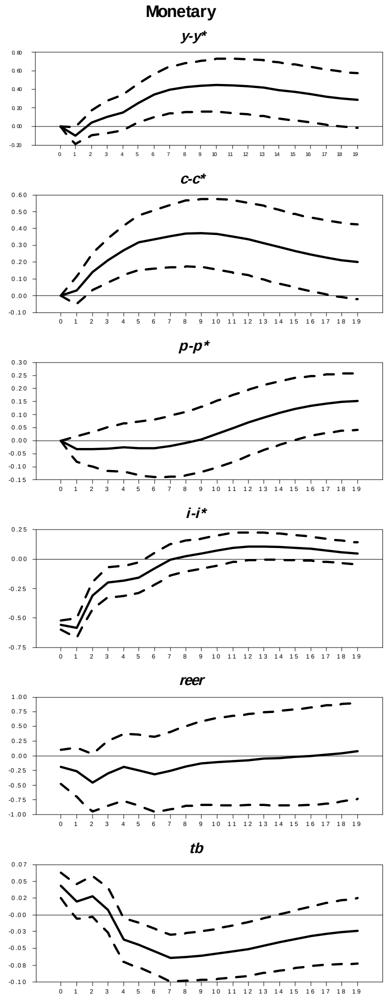
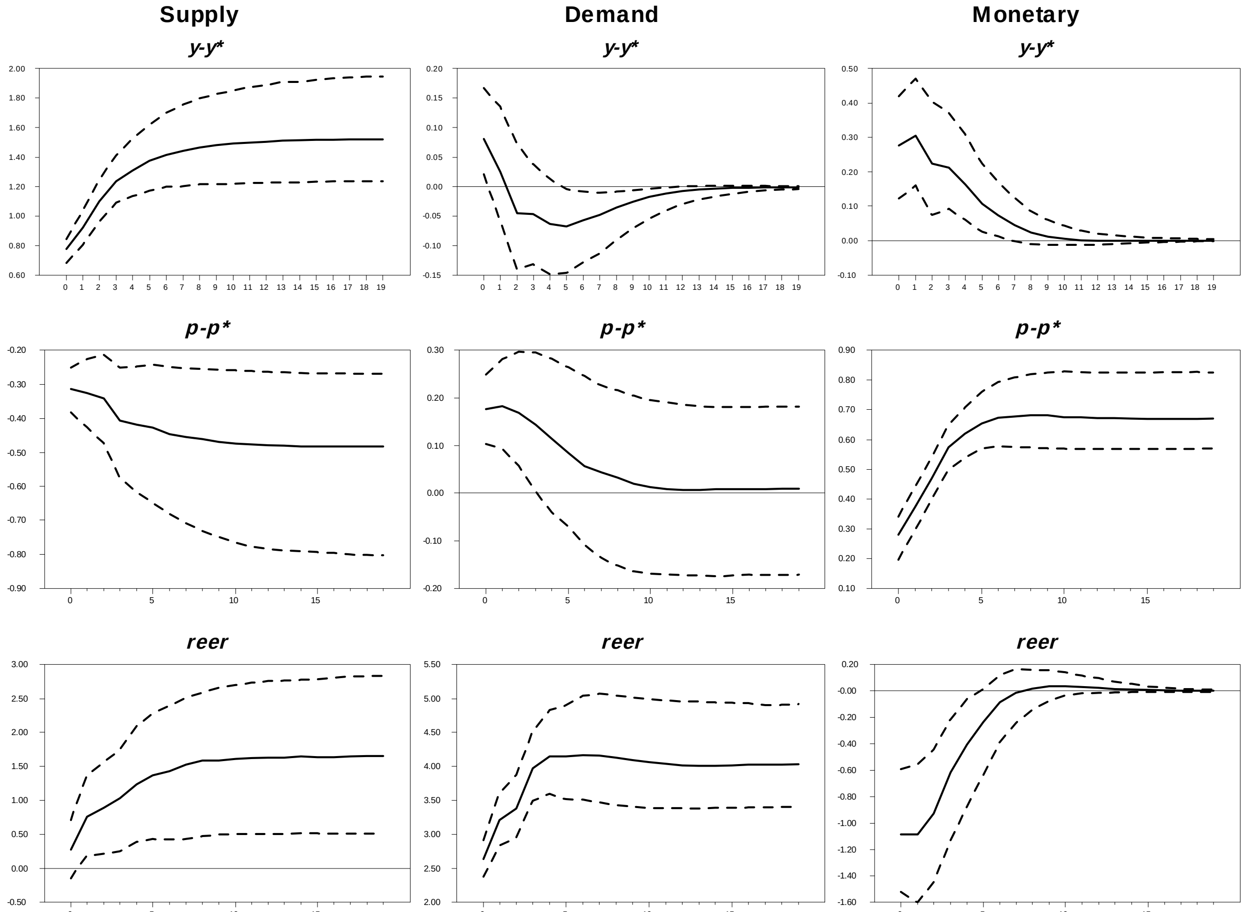


Figure 8. Impulse Responses: Choleski Decompostition



Notes: The figure shows the impulse responses to a monetary policy shock using the Choleski decomposition. Solid lines are point estimates and the dashed lines represent the 16th and 84th percentile error bands.

Figure 9. Impulse Responses: Long-Run Restrictions



Notes: The figure shows the impulse responses to supply, demand, and nominal shocks using zero long-run restrictions. The solid lines are point estimates and the dashed lines represent the 16th and 84th percentile error bands.