

Nominal Shocks and Real Exchange Rate Fluctuations*

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Abstract

I analyze the role of nominal and real shocks on the exchange rate behavior using a structural vector autoregressive model (sVAR) for the US vis-à-vis the rest of the world. I estimate the contribution of various shocks to explaining the movement of the real dollar exchange rate, imposing two alternative identification strategies based on theoretical restrictions derived from an open economy macro model. Using zero long-run restrictions I find that nominal shocks are unimportant to explain real exchange rate fluctuations. I compare this result to an identification strategy based on sign restrictions proposed by Fry and Pagan (2007). The results show that although monetary policy shocks are not the main drivers of exchange rate fluctuations, they are nevertheless important. The range for their short-horizon contribution is 37% to 47%.

Keywords: Exchange Rates, Nominal Shocks, Vector Autoregression, Sign-restrictions.

JEL Classification: F31; F41; C30.

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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 for real exchange rate determination. This paper combines a conventional open economy macro model with recent econometric developments to examine the impact of nominal and real shocks on the real exchange rate behavior.

A natural starting point for the link between monetary policy and exchange rates is Dornbusch (1976) model. Dornbusch prediction is that in response to a monetary policy shock the exchange rate would immediately overshoot its long run level and then adjust gradually. This conclusion follows from two cornerstone arbitrage conditions: uncovered interest parity (UIP) and purchasing power parity (PPP). Since Dornbusch's seminal work, the theoretical literature has mainly concentrated on the effects of monetary policy shocks (see Obstfeld and Rogoff, 1995; Beaudry and Devereux, 1995; Chari et al, 2002). Thus, 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 to this. In a seminal paper, Clarida and Gali (1994) estimate the relative contribution of various shocks to 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 Canadian cases. 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%.

Motivated by the previous literature, I concentrate on three issues. Firstly, I analyze the effects of nominal and real shocks on the real dollar exchange rate. Secondly, I quantify the importance of nominal vs real shocks for the variance of the real exchange rate. Thirdly, I evaluate the contribution of each of the shocks to the real exchange rate behavior. The first issue will allow us to study the impact of various shocks on the real exchange rate and to determine if the effects of monetary policy

are consistent with Dornbusch's overshooting. The second one addresses the question of whether monetary policy explains a large share of exchange rate fluctuations, as would be expected from theory. The last one will allow us to link the sources of exchange rate movements with economic factors and thus will shed some light on the nature of real exchange rate fluctuations.

The first two questions of my analysis have been extensively analyzed in the literature. For example, Clarida and Gali (1994) and Eichenbaum and Evans (1995) find that the exchange rate overshoots its long-run value after a monetary policy shock but that the peak occurs from one to three years and not on impact as predicted by Dornbusch (1976). This result, known as delayed overshooting, became a consensus result in international economics. In a related study, Faust and Rogers (2003) find that the delayed overshooting result is sensitive to dubious assumptions of conventional estimation methods. In particular, when simultaneity is allowed between financial market variables (interest rates and exchange rates), they find that overshooting is nearly immediate. In line with Faust and Rogers (2003), Scholl and Uhlig (2006) show that the exchange rate overshooting takes place much quicker than the three years horizon found in Eichenbaum and Evans (1995). The literature has not given a clear cut answer about the proportion of exchange rate variability explained by monetary policy shocks. In fact, results vary considerably depending on the countries under consideration and the identification strategy.

The empirical work cited above builds on the use of a structural VAR method to identify monetary policy shocks. The identification strategy employed varies from study to study and the ones relying 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 arbitrary assumptions which 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). Although long-run restrictions are often better justified by economic theory, there seem to be cases in which substantial distortions can arise due to the presence of a small sample bias (Faust and Leeper, 1997) or a lag-truncation bias (Chari, Kehoe and McGrattan, 2007).

The empirical work based around the use of sign restrictions such as Faust and Rogers (2003) and Scholl and Uhlig (2006) are not subject to the criticism of relying on arbitrary assumptions given that the identifying assumptions 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) this procedure does not guarantee that the identification of structural shocks is exact because there are multiple matrices defining the linear mapping from orthogonal structural shocks to VAR residuals.

As a way to obtain more precise estimates of impulse responses, the sign restrictions method can be generalized to identify more than one shock. In this case the estimation is more precise because the range of reasonable impulse responses is narrowed down. 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 are 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 restrictions method that identifies various shocks of interest. They find an important role of the real exchange rate as a shock absorber, mainly of demand shocks. In a related study, Fry and Pagan (2007) highlight some conceptual problems of this method, which result from the multiplicity of impulse vectors and suggest a rule to pin down unique impulse responses using sign restrictions.

In an attempt to overcome the limitations previously mentioned, this paper builds on Farrant and Peersman (2006) but refines the estimation strategy by applying the method developed by Fry and Pagan (2007). The results based on this approach yield an important addition to the literature. I provide evidence that monetary policy shocks are a relevant driver of the real exchange rate. This reconciles the focus of the theoretical literature on monetary policy with the empirical evidence.

More precisely, I estimate the standard 3 variable VAR of Clarida and Gali (1994) composed of relative output, relative prices and the real exchange rate and identify supply, demand and nominal shocks. In the paper I will refer interchangeably to nominal and monetary shocks. Based on the theoretical restrictions derived from the Clarida-Gali model, I use two alternative identification strategies. The first one is based on the Blanchard and Quah (1989) method and is derived from the long-run dynamics of the model. The second one employs sign-restrictions derived from the theoretical short-run predictions of the model.

Using long-run restrictions I get the conventional empirical finding that nominal shocks are unimportant to explain real exchange rate fluctuations. The evidence

I present in this paper indicates that, in contrast to this result, when using sign restrictions monetary (or nominal) shocks are an important driver of exchange rate movements at short horizons. I find that at a 4 quarters horizon, the contribution of monetary policy shocks is 47%. Their relative importance is reduced at longer horizons. Indeed, at a horizon of 20 quarters only 20% of the variation of the real exchange rate is explained by monetary policy shocks. Our results differ from what is typically found in the literature.¹

I then examine the sensitivity of my results to the use of alternative sign restrictions, to a sub-sample analysis, to a different exchange rate measure and to an extended model. Overall, I find that the results are robust to these alternative specifications. I also examine the robustness of my results to an extended model where the identification strategy is based on zero short-run restrictions. When using this method monetary policy shocks explain only 10% 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.

The remainder of the paper is organized as follows. Section 2 outlines the Clarida and Gali (1994) model. Section 3 describes the data and also presents unit root and cointegration tests. The results using the long-run identification strategy are presented in Section 4. Section 5 contains the empirical methodology based on a structural VAR framework with sign restrictions. The results of the baseline model are presented in Section 6 while I report a battery of robustness tests in Section 7. Section 8 concludes.

2 Theoretical model

I estimate a structural VAR using two alternative identification strategies based on theoretical restrictions derived from the Clarida-Gali model. The first identification strategy is the same as the one used by Clarida-Gali and relies on imposing zero long-run restrictions based on the long-run dynamics of the model. The second one employs sign-restrictions derived from the theoretical short-run predictions of the model.

The Clarida-Gali model is a stochastic version of the two-country rational expectations model by Obstfeld (1985). The model captures the short-run dynamics of the Mundell-Fleming-Dornbusch approach when prices adjust sluggishly to demand, supply and monetary shocks as well as the long-run dynamics, when the economy is in equilibrium and prices fully adjust to all shocks. The following set of equations

¹An important exception is Rogers (1999), where he presents evidence that monetary policy shocks are generally important for real exchange rate movements.

describe the model. All variables are in logs except interest rates and represent home (US) relative to foreign (rest of the world) levels.

$$y_t^d = d_t + \eta(s_t - p_t) - \sigma[i_t - E_t(p_{t+1} - p_t)] \quad (1)$$

$$p_t = (1 - \varphi)E_{t-1}p_t^e + \varphi p_t^e \quad (2)$$

$$m_t^s - p_t = y_t - \lambda i_t \quad (3)$$

$$i_t = E_t(s_{t+1} - s_t) \quad (4)$$

Equation (1) is an open-economy IS equation, in which relative output demand (y_t^d) depends positively on a relative demand shock (d_t) and on the real exchange rate ($q_t = s_t - p_t$), and is decreasing in the real interest rate differential. In this model d_t captures any shock to home absorption relative to foreign absorption (e.g. fiscal shocks). Equation (2) is a price setting equation in which the relative price level in period t (p_t) is an average of the market-clearing price in $t - 1$, ($E_{t-1}p_t^e$) and the price that would actually clear the market in time t , (p_t^e). The parameter φ is a measure of price sluggishness. When $\varphi = 1$ prices are fully flexible and output is supply determined and when $\varphi = 0$ prices are fixed and predetermined one period in advance. Equation (3) is an LM equation in which real money balances are increasing in relative output (y_t) and decreasing in the relative interest rate (i_t). Equation (4) is an interest parity condition in which s_t denotes the nominal exchange rate.

In the Clarida and Gali (1994) model, relative output, relative prices and the real exchange rate are driven by three shocks: a relative supply shock, a relative demand shock and a relative monetary shock.² Both y_t^s and m_t are assumed to follow a random walk process and d_t is characterized by a transitory and permanent component:

$$y_t^s = y_{t-1}^s + \epsilon_t^s \quad (5)$$

$$d_t = d_{t-1} + \epsilon_t^d - \gamma \epsilon_{t-1}^d \quad (6)$$

$$m_t = m_{t-1} + \epsilon_t^m \quad (7)$$

The model can be solved for the long-run flexible-price equilibrium, described by the following equations:³

²Note that the monetary shock reflects both shocks to relative money supply and to relative money demand.

³For details on the derivation of the model see Clarida and Gali (1994).

$$y_t^e = y_t^s \quad (8)$$

$$q_t^e = (y_t^s - d_t) / \eta + [\eta(\eta + \sigma)]^{-1} \sigma \gamma \varepsilon_t^d \quad (9)$$

$$p_t^e = m_t - y_t^s + \lambda(1 + \lambda)^{-1}(\eta + \sigma)^{-1} \gamma \varepsilon_t^d \quad (10)$$

Thus, in the long run, only supply shocks lead to increases in the level of relative output; supply and demand shocks have an impact in the long-run level of the real exchange rate; and the three shocks have an impact on relative prices in the long run. The theoretical long-run predictions of the model are summarized in Table 1, where \times denotes that the shocks have an impact on the variables in the long-run and 0 implies that they do not.

Table 1. Theoretical Predictions

Shock	Long-run predictions			Short-run predictions		
	$y - y^*$	q	$p - p^*$	$y - y^*$	q	$p - p^*$
Supply	\times	\times	\times	+	-	-
Demand	0	\times	\times	+	+	+
Nominal	0	0	\times	+	-	+

The zero long-run restrictions are used by Clarida and Gali (1994) to identify supply, demand and nominal shocks. I will replicate their identification strategy but I will also use the theoretical short-run predictions of their model to estimate the impact of the shocks of interest employing a method based on sign restrictions.

Solving the system for the case of a short-run open economy equilibrium with sluggish price adjustment, we obtain:

$$p_t = p_t^e - (1 - \varphi) (\epsilon_t^m - \epsilon_t^s + \alpha \gamma \varepsilon_t^d) \quad (11)$$

$$q_t = q_t^e + v(1 - \varphi) (\epsilon_t^m - \epsilon_t^s + \alpha \gamma \varepsilon_t^d) \quad (12)$$

$$y_t = y_t^s + (\eta + \sigma) v(1 - \varphi) (\epsilon_t^m - \epsilon_t^s + \gamma \varepsilon_t^d) \quad (13)$$

where $\alpha \equiv \lambda(1 + \lambda)^{-1}(\eta + \sigma)^{-1}$, $v \equiv (1 + \lambda)(\lambda + \sigma + \eta)^{-1}$.

Equations (11), (12) and (13) describe the evolution of relative price levels, the real exchange rate and relative output in the short run, when the economy is characterized by a certain degree of price sluggishness, such that $1 < \varphi < 0$.

The predictions of the model are the following. In response to a positive monetary shock, relative output and relative prices increase and the exchange rate depreciates.

A demand shock leads to an increase in output and relative prices and an appreciation of the real exchange rate. In response to a supply shock relative output increases, relative prices decrease and the real exchange rate depreciates. These restrictions, summarized as well in Table 1 above, refer to the short-run impact of the three shocks on the variables of interest.⁴ I will incorporate these short-run sign restrictions into an empirical model and identify the shocks of interest in an alternative way with respect to the original Clarida-Gali approach.

3 Data

I use quarterly data over the period 1976-2006. The “rest of the world” (hereafter ROW) series include an aggregate of the other G7 countries (except the US) and another OECD economy (Australia) and are identified by an asterisk in our notation.⁵

All the series are taken from the *International Financial Statistics* (IFS) of the International Monetary Fund (IMF). The data on real GDPs are seasonally adjusted in local currencies at 2000 price levels. I convert the GDP series in local currencies to US dollars using the average market exchange rate of 2000 (I do this to preserve consistency with the prices 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 US real GDP (y) is measured in deviation from GDP in the ROW (y^*), which is the sum of output in the other G7 plus Australia. Price series correspond to the consumer price index (CPI). In section 7 I estimate an extended model which includes the interest rate differential. Federal funds rates are used for the US and 3-month money market rates for the other countries.⁶ The series p^* and i^* are calculated respectively, as an aggregate of prices and interest rates in the ROW weighted according to their respective (time-varying) GDP shares at PPP values. The GDPs used for calculating the weights are at price levels and PPP values of 2000 and obtained from the OECD. The real effective exchange rate (q) corresponds to the REU series of the IFS. I check robustness to an alternative measure of the real exchange rate, labeled $q1$, taken from the US Federal Reserve Board Statistics. Figure 1 contains plots of the series (in levels) used in the VAR.

⁴In the data the real effective exchange rate is expressed as the number of foreign currency per unit of domestic currency. Thus, an appreciation (depreciation) of the dollar implies an increase (reduction) of the real exchange rate.

⁵These 8 countries add up to roughly half of world GDP at PPP values, so they represent a substantial sample of the global economy. Moreover, trade flows among them also amount to over a half of their respective total trade, on average.

⁶I prefer to use the federal funds rates for the US given that it is the one commonly used in previous studies (see Uhlig, 2005). The results do not vary when I consider the 3-month treasury bill rate for the US.

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 trend-stationary null for the levels of relative output, relative prices and real effective exchange rate. Inference is mixed for the level of relative interest rates given that there is evidence of trend-stationarity from one of the ADF tests and the KPSS test. By contrast, all variables are trend-stationary in first differences.

The Clarida-Gali model implies that there are no cointegration relationships among the levels of the variables. Table A2 in the appendix attends to this. It 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 both for the baseline and the extended model. Overall, these results suggest estimating the VARs in first differences.

4 Long-run identification

In this section I will reproduce the results of Clarida and Gali (1994) on my data and sample period. I will first present a brief description of the estimation method. Consider the following structural VAR model

$$Y_t = C(L)\epsilon_t \quad (14)$$

where Y_t is an $N \times 1$ vector of endogenous variables, C is a matrix of coefficients in the lag operator L and ϵ_t is an $N \times 1$ vector of structural disturbances. The endogenous variables Y_t that we include in the VAR are the first difference of relative output (Δy_t), the first difference of the real exchange rate (Δq_t) and the first difference of relative prices (Δp_t). Thus, $Y_t' = [\Delta y_t \quad \Delta q_t \quad \Delta p_t]$. My aim is to identify three types of innovations: a supply, a demand and a monetary shock, $\epsilon_t' = [\epsilon^s \quad \epsilon^d \quad \epsilon^m]$. Identification requires $N \times (N - 1) / 2$ restrictions on the coefficients of the long-run moving average coefficient matrix, $C(1)$. The open economy macro model described in section 2 is triangular in the long-run. The zero long-run restrictions summarized in Table 1 are represented in (15)

$$C(1) = \begin{bmatrix} c_{11} & 0 & 0 \\ c_{21} & c_{22} & 0 \\ c_{31} & c_{32} & c_{33} \end{bmatrix} \quad (15)$$

The restrictions originate in the fact that supply shocks are expected to influence output in the long run, while both supply and demand shocks have an impact on the real exchange rate in the long run.

The restriction $c_{12} = 0$ implies that demand shocks, ϵ^d , have no effect on output in the long run. The restriction that nominal shocks have no impact on relative output and the real exchange rate in the long run justifies that $c_{13} = 0$ and $c_{23} = 0$. Note that by only imposing zero long-run restrictions we leave the short-run dynamics of the model unrestricted. Thus, we can assess whether the shocks that this approach identifies as due to supply, demand and money are consistent with the short-run dynamics of the model presented in section 2.

Figure 2 shows the impulse responses to a supply, demand and nominal shock together with the 16th and 84th percentiles error bands calculated using Monte Carlo integration. Consistent with the predictions of the model, a supply shock has a persistent positive effect in relative output and a persistent negative effect on relative prices. However, after a supply shock the real exchange rate exhibits a temporary appreciation, which is in contrast with the predictions of the theoretical model. In line with the theoretical model, demand shocks lead to a temporary increase in relative output. This result is obtained by construction, because demand shocks are restricted not to affect relative output in the long run. After a demand shock the real exchange rate exhibits a persistent appreciation. Finally, a monetary policy shock has a temporary effect on relative output and a persistent effect on relative prices. Interestingly, after a monetary shock the real exchange rate overshoots its long-run value almost immediately (in one quarter). This is in contrast to the popular one to three years delayed overshooting found in the earlier literature (see Eichenbaum and Evans, 1995).

The variance decomposition can be used to assess how much of the variance of the real exchange rate is given by supply, demand and monetary shocks over different forecast horizons. The results, given in Table 2, are in line with the ones presented in Clarida and Gali (1994). Nominal shocks are unimportant to explain real exchange rates fluctuations, as demand shocks are the main driver of the real exchange rate. At a horizon of 4 and 20 quarters, 86% and 93% of the variation of the real exchange rate is explained by demand shocks, respectively. By contrast, nominal shocks explain between 1% and 7% of the variance of the real exchange rate at 4 and 20 quarters, respectively. The results also show that supply shocks play a minimal role in explaining real exchange rate fluctuations.

Table 2. Variance Decomposition of the Real Effective Exchange Rate
(Long-run restrictions, Sign restrictions and short-run restrictions)

		Shocks		
Steps		Supply	Demand	Nominal
4 quarters	CG	0.06 [0.01 ; 0.17]	0.86 [0.71 ; 0.94]	0.07 [0.02 ; 0.15]
	SR1	0.03 [0.00 ; 0.11]	0.50 [0.24 ; 0.71]	0.47 [0.23 ; 0.71]
	SR2	0.02 [0.00 ; 0.11]	0.49 [0.24 ; 0.71]	0.48 [0.23 ; 0.71]
8 quarters	CG	0.05 [0.01 ; 0.16]	0.90 [0.78 ; 0.96]	0.03 [0.01 ; 0.08]
	SR1	0.05 [0.01 ; 0.14]	0.58 [0.31 ; 0.80]	0.37 [0.14 ; 0.61]
	SR2	0.04 [0.01 ; 0.14]	0.57 [0.31 ; 0.80]	0.38 [0.14 ; 0.61]
12 quarters	CG	0.05 [0.01 ; 0.16]	0.92 [0.81 ; 0.97]	0.02 [0.01 ; 0.05]
	SR1	0.07 [0.01 ; 0.17]	0.65 [0.35 ; 0.84]	0.28 [0.09 ; 0.54]
	SR2	0.06 [0.01 ; 0.17]	0.66 [0.35 ; 0.84]	0.27 [0.09 ; 0.54]
16 quarters	CG	0.05 [0.01 ; 0.16]	0.92 [0.82 ; 0.98]	0.02 [0.00 ; 0.04]
	SR1	0.08 [0.01 ; 0.19]	0.68 [0.39 ; 0.86]	0.24 [0.07 ; 0.48]
	SR2	0.07 [0.01 ; 0.19]	0.68 [0.39 ; 0.86]	0.24 [0.07 ; 0.48]
20 quarters	CG	0.05 [0.01 ; 0.16]	0.93 [0.83 ; 0.98]	0.01 [0.00 ; 0.03]
	SR1	0.10 [0.01 ; 0.20]	0.70 [0.42 ; 0.87]	0.20 [0.06 ; 0.44]
	SR2	0.08 [0.01 ; 0.20]	0.70 [0.42 ; 0.87]	0.22 [0.06 ; 0.44]

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 quarter horizons. Lag length is 4. In brackets are the 16th and 84th percentiles error bands. CG denotes that identification is based on the Clarida-Gali approach. SR1 refers to the baseline sign restrictions described in Table 1. SR2 denotes the alternative sign restrictions presented in section 7.

5 Identification using short-run sign restrictions

The method based on long-run restrictions described in the previous section is very appealing because it is often justified by economic theory. However, the technique has a series of shortcomings. One is that there are cases in which the impulse responses contradict predictions of the theoretical model. Another is related to the finding of Faust and Leeper (1997) who show that substantial distortions can arise due to small sample biases and measurement errors when using long-run restrictions. In a related study, Chari, Kehoe and McGrattan (2007) show that sVARs with long-run restrictions suffer from a lag-truncation bias. This happens because the available data requires a VAR with a small number of lags, which is a poor approximation of the infinite-order VAR of the observables from the model.

Conventional methods involving zero-short run restrictions, such as the Choleski decomposition, have also been questioned on various grounds. Firstly, they are usually derived from some arbitrary assumptions which 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 mod-

els (see Canova and Pina, 1999). Secondly, they sometimes yield counter-intuitive impulse response functions of key endogenous variables which are not easy to rationalize 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, Eichenbaum and Evans, 1999; Kim and Roubini, 2000). Thirdly, as noted by Sarno and Thornton (2004), the results are often sensitive to the ordering of the variables.

In order to overcome the potential problems of the previous methods, I use an alternative identification procedure based on sign restrictions. Faust (1998), Canova and de Nicoló (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 Fry and Pagan (2007), who highlight some conceptual problems arising from the multiplicity of impulse vectors. Their approach builds on Mountford and Uhlig (2005) and Peersman (2005).

I would like to emphasize that the decision to use the approach of sign restrictions proposed by Fry and Pagan (2007) does not represent a general criticism of other sign restrictions methods. Indeed, some authors have proposed the use of identification schemes that pin down unique impulse responses. Uhlig (2005), for example, proposes the use of a penalty function.

I am interested in analyzing the impact of demand, supply and nominal shocks on the real exchange rate avoiding the problems arising from imposing arbitrary assumptions. Thus, the identification of the shocks is based on the theoretical short-run predictions of the Clarida-Gali model outlined in section 2. In contrast to what I did in the previous section, I will now impose restrictions in the short-run and let the data speak in the long-run.

5.1 VAR model with sign restrictions

Consider the reduced form VAR

$$Y_t = c + B(L)Y_{t-1} + A\epsilon_t \tag{16}$$

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. The endogenous variables Y_t that we include in the VAR are the same as the ones in the previous section.

My aim is to identify three types of innovations: a supply, a demand and a mon-

etary shock, $\epsilon_t' = [\epsilon^s \ \epsilon^d \ \epsilon^m]$. I identify these shocks using a sign restriction approach. Since the shocks are assumed to be orthogonal, so that $E[\epsilon_t \epsilon_t'] = I$, the variance-covariance matrix of equation (16) is equal to: $\Sigma = AA'$. For any orthogonal decomposition of A , we can find an infinite number of possible orthogonal decomposition 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 lower triangular. 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 Nicoló (2002), I consider $P = \prod_{m=1}^{N-1} \prod_{n=m+1}^N Q_{m,n}(\theta)$, where $Q_{m,n}(\theta)$ is an orthonormal rotational matrix of the form:

$$Q_{m,n} = \begin{bmatrix} 1 & 0 & \dots & 0 & 0 \\ 0 & \cos(\theta) & \dots & -\sin(\theta) & 0 \\ \dots & \dots & 1 & \dots & \dots \\ 0 & \sin(\theta) & \dots & \cos(\theta) & 0 \\ 0 & 0 & \dots & 0 & 1 \end{bmatrix} \quad (17)$$

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

In a 3 variable model we have a 3x3 rotational matrix Q and 3 bivariate rotations.⁷ The angles $\theta = \theta_1, \dots, \theta_3$, and the rows m and n are rotated such that

$$P = \begin{bmatrix} \cos(\theta_1) & -\sin(\theta_1) & 0 \\ \sin(\theta_1) & \cos(\theta_1) & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} 1 & 0 & 0 \\ \dots & \cos(\theta_2) & -\sin(\theta_2) \\ 0 & \sin(\theta_2) & \cos(\theta_2) \end{bmatrix} \begin{bmatrix} \cos(\theta_3) & 0 & -\sin(\theta_3) \\ 0 & 1 & 0 \\ \sin(\theta_3) & 0 & \cos(\theta_3) \end{bmatrix}$$

My estimation strategy follows Fry and Pagan (2007) and Peersman (2005) and is carried out as follows. Firstly, all possible rotations are produced by varying the rotation angles $\theta_1, \theta_2, \theta_3$ in the range $[0, \pi]$. For practical purposes, I grid the interval $[0, \pi]$ into M points. 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, $A_j (j = 1, \dots, M^3)$ as follows:

$$R_{j,t+k} = [I - B(L)]^{-1} A_j \epsilon_t \quad (18)$$

where $R_{j,t+k}$ is the matrix of impulse responses at horizon k . In order to identify

⁷In general terms, we have a total of $N(N-1)/2$, where N is the number of variables.

the shock v of interest, sign restrictions can be imposed on $p \leq n$ variables over the horizon $0, \dots, K$ in the form:

$$R_{j,t+k}^{p, v} \leq 0 \quad (19)$$

The sign restrictions are imposed based on the Clarida-Gali open economy model as summarized in Table 1 over the time horizon $k = 0, \dots, K$. Details about the number of periods for which the restrictions hold are given below.

To identify a supply shock, s , I impose that relative output does not decrease, relative prices do not increase for four quarters ($K = 4$) and that the real exchange rate does not appreciate for one quarter ($K = 1$):⁸

$$\begin{aligned} R_{j,t+k}^{y-y^*, s} &\geq 0, \quad k = 0, \dots, 4 \\ R_{j,t+k}^{p-p^*, s} &\leq 0, \quad k = 0, \dots, 4 \\ R_{j,t+k}^{q, s} &\leq 0, \quad k = 0, 1 \end{aligned}$$

The restrictions to identify a demand shock, d , are that relative output and relative prices do not decrease for four quarters and that the real exchange rate does not depreciate for one quarter:

$$\begin{aligned} R_{j,t+k}^{y-y^*, d} &\geq 0, \quad k = 0, \dots, 4 \\ R_{j,t+k}^{p-p^*, d} &\geq 0, \quad k = 0, \dots, 4 \\ R_{j,t+k}^{q, d} &\geq 0, \quad k = 0, 1 \end{aligned}$$

A monetary policy shock, m , is identified by restricting relative output and relative prices not to decrease for four quarters and the real exchange rate not to appreciate for one quarter:

$$\begin{aligned} R_{j,t+k}^{y-y^*, m} &\geq 0, \quad k = 0, \dots, 4 \\ R_{j,t+k}^{p-p^*, m} &\geq 0, \quad k = 0, \dots, 4 \\ R_{j,t+k}^{q, m} &\leq 0, \quad k = 0, 1 \end{aligned}$$

Out of all possible rotations I select those that satisfy the sign restrictions of the impulse responses of the three shocks. Impulse responses are constructed using a

⁸Changing the values of the number of quarters for which the restrictions are binding has no effect on the conclusions of the results.

Monte Carlo experiment. For each Monte Carlo draw, I draw one rotation out of all possible rotations and check if the imposed restrictions are satisfied for all shocks. Solutions that satisfy all the restrictions are kept and the others are discarded. In practice I repeat this procedure until 1000 draws satisfying the restrictions are found.

5.1.1 Summarizing the information

A common way to present the information is by reporting the median of the impulse responses based on a certain number of solutions. As noted by Fry and Pagan (2007), this may not provide a useful measure. Let us consider as an example the responses of two variables to one shock. Say that $\tau_1(i)$ represents the impulse responses of variable 1 and $\tau_2(i)$ represents the impulse responses of variable 2, where i indexes the values of θ . What is usually presented is $med((\tau_1(i)))$ and $med((\tau_2(i)))$. If $\tau_1(i)$ was monotonic in θ , this would be $\tau_1(med(\theta(i)))$. Thus, the median of the impulse responses would be generated by the same model, represented by $med(\theta(i))$. Given that there is no guarantee of monotonicity, the median of the impulse responses will generally be associated with different values of θ . Thus, the median of the impulse responses will not come from a single model. This will happen for the impulse responses of all variables and for all shocks. As a consequence, the shocks identified are no longer orthogonal to each other. Apart from the fact that the assumption of uncorrelated shocks is essential to estimate a VAR model, if correlations are non-zero, the variance decomposition does not exist.⁹

Fry and Pagan (2007) suggest locating a unique vector of θ 's such that the impulses are closest to its median while maintaining the orthogonality condition. This method works as follows. Impulse responses are calculated based on those that satisfy the sign restrictions as described above. Given that the impulse responses are not unit free, they are standardized by subtracting their median and dividing by their standard deviation. These standardized impulse responses are included in a vector ϕ_i for each value of θ_i (in the 2 variable case below ϕ is a 4×1 vector as there are 4 impulse responses). The choice of θ is the one that minimizes the following quantity.

$$\Psi(\theta_i) = \phi_i' \phi_i \tag{20}$$

Substituting θ_{\min} into $Q(\theta)$ will produce a set of orthogonal shocks and the descriptive statistics are computed based on this rotation.

⁹This problem remains even if only one single shock is identified, as in Uhlig (2005), since there is nothing that ensures it is uncorrelated with the remaining shocks in the model when considering the median of impulse responses.

6 Empirical results

6.1 Estimates of the baseline model

I now turn to the empirical findings, by presenting the benchmark results from implementing the VAR described in Section 5.

Figure 3 shows the impulse responses of relative output, relative inflation and the exchange rate with respect to the three shocks of interest. The impulse responses suggest that a supply shock leads to a persistent depreciation of the real exchange rate. More precisely, the US dollar depreciates 0.5% on impact and it continues to decline until it reaches a depreciation of 1.3% after 9 quarters. In addition, a supply shock generates a persistent increase in relative output and a persistent decrease in relative prices.

After a demand shock, there is a persistent rise of around 0.8% in relative output. By contrast, there is a temporary effect on relative prices, which increase 0.07% on impact. The real exchange rate appreciates 2% on impact and it continues to rise up to 3.7% after 9 quarters.

Finally, a monetary policy (or nominal) shock leads to a temporary depreciation of the US dollar. After an initial depreciation of 2%, the real exchange rate reaches its minimum value after 3 quarters and then reverts to equilibrium (consistent with PPP), showing no statistically significant reaction after 10 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 (it increases 0.5% on impact and shows no significant reaction after 6 quarters) and a persistent rise in relative prices, which increase by 1% after 9 quarters.

The long-run predictions of the Clarida-Gali model are satisfied empirically except for the impact of a demand shock on relative output. According to the theoretical model, only supply shocks influence the long run level of output. However, the empirical results show that demand shocks also have a persistent effect on output.

Table 2 compares the variance decomposition of the real exchange rate using the sign restrictions (SR1) and the Clarida-Gali (CG) methodologies. From Table 2 it is clear that there is a stark difference in the results from both identification procedures. Using the sign-restrictions method we conclude that a substantial share of the variation in the real exchange rate is explained by monetary shocks over short horizons but at long horizons their relative importance is reduced. Indeed, at a 4 quarters horizon the contribution of monetary policy shocks to real exchange rate fluctuations is 47%. At a horizon of 20 quarters only 20% of the variation of the

real exchange rate is explained by monetary policy shocks. By contrast, demand shocks explain a substantial proportion of the variance of the real exchange rate both at short and long horizons (50% and 70% at a horizon of 4 quarters and 20 quarters respectively). The results also show that supply shocks play a minimal role in explaining real exchange rate fluctuations. In particular, I find that supply shocks explain 3%, 7% and 10% of the movement in the real exchange rate at a 4 quarters, 12 quarters and 20 quarters respectively.

In summary, using sign-restrictions the findings indicate that demand shocks have been the main determinant of real exchange rate fluctuations both at short and long horizons. Although monetary policy shocks are not the main drivers of the real exchange rate, they are nevertheless important over short horizons. This result is consistent with the theoretical model in that monetary policy shocks have transitory effects on the exchange rate. Thus, most of its impact should be materialized in the short run. By contrast, using long-run restrictions in the same fashion as Clarida and Gali (1994) leads to the conclusion that nominal shocks are unimportant to explain real exchange rates fluctuations. In the Clarida-Gali identification scheme, demand shocks account for most of the variance of the real exchange rate.

One source of difference between the two methods relies on the shocks identified. The time series of the shocks using both methodologies is presented in Figure 4 and Table 3 reports the correlations of the shocks across both methodologies.

Table 3. Correlation of shocks across methodologies

Clarida-Gali	Sign restrictions		
	Supply	Demand	Nominal
Supply	0.64	0.69	-0.01
Demand	-0.43	0.65	-0.46
Nominal	-0.23	0.28	0.89

The Table shows that the correlation is high for nominal shocks (0.89). For supply and demand shocks the correlations are low, 0.64 and 0.65 respectively. The table also indicates that part of the supply shocks of the long-run restrictions approach are captured by the demand shocks using sign restrictions. In addition, the demand shocks of the Clarida-Gali approach are now picked up by supply and nominal shocks.

6.2 Interpreting Real Exchange Rate Fluctuations

Variance decompositions reveal which shocks are important in explaining the variance of the real exchange rate across different horizons. However, this measure does not provide a complete picture of the nature of real exchange rate fluctuations. In order

to relate the sources of exchange rate movements with economic factors it is useful to analyze the historical decomposition of the real exchange rate, which refers to the contribution of each of the shocks to the path of the real exchange rate.

Figure 5 compares the historical decomposition of the real exchange rate for the period 1976:1 to 2006:4 using the Clarida-Gali and the sign restrictions methodologies. The historical decomposition calculated using the Clarida-Gali approach reveals that demand shocks were the main drivers of the real exchange rate and that there is little role for other shocks. Using the sign restriction approach we get a different perspective on the sources of exchange rate movements. In the rest of this section I will concentrate on the historical decomposition based on sign restrictions.

From the graph of the real effective exchange rate in Figure 1, it is possible to distinguish a first episode of dollar depreciation between 1978 and 1980. The historical decomposition shows that monetary policy shocks were the main drivers of exchange rate fluctuations during this episode.

Between 1980 to 1985 the dollar appreciates significantly. As described in Obstfeld (1995), observers differ as to whether important shifts in fundamental factors such as the Volcker disinflation, the Reagan fiscal expansion and the fiscal contraction outside the US can explain this rise. The historical decomposition allows us to distinguish the contribution of monetary policy and demand shocks to the real exchange rate increase. The Volcker disinflation is clearly linked to a monetary policy shock and the fiscal expansion in the US and fiscal contraction abroad is associated with a demand shock. Figure 5 shows that monetary policy shocks play a dominant role in the dollar appreciation at the beginning of this period and demand shocks become the main contributor to exchange rate movements between 1983 and 1985. The relative importance of monetary policy shocks may be attenuated due to the fact that as a result of the dollar appreciation, other industrialized countries faced depreciating currencies and inflationary pressures. The policy response to this was a contractive monetary policy and consequent interest rate increases, which lowers the magnitude of the interest rate differential.

In the period from 1986 until 1988 the dollar declined significantly. This period is usually not identified as one of dollar weakness if one considers its value with respect to the 1976-2006 average. However, there is a decline with respect to the previous episode of appreciation. Figure 5 shows that both monetary and demand shocks played an important role in explaining the real exchange rate movement during this episode. This pattern squares very well with some events that took place during this period. In October 1987 there was a US and global stock market collapse which should have lead to a decrease in demand. In response to the crash the Fed decreased

interest rates.

Between 1989 and 1995 the dollar was more stable than in the previous ten years. The decline from 1990 to 1991 is mainly due to demand shocks given that the US economy moved into a mild recession during this period in the context of the Gulf War. A further decline of the dollar took place within the ERM crisis of 1992-1993.

Starting in 1995, there is a continuous increase in the value of the dollar until 2002. The historical decomposition reveals that this was mainly due to positive demand shocks and that the contribution of monetary policy shocks was much less important between 1995 and 2000. Many observers in fact highlighted that initially the dollar strength during this period was associated with a healthy US economy and strong demand (see Truman, 2006).

7 Robustness: estimating alternative models

Empirical results often depend on modelling assumptions and variables definitions. Thus, in this section I estimate different VAR specifications and I also use alternative variables definitions to assess the robustness of my results.

7.1 Alternative sign restrictions

A main difference that emerges from the results between both methods is the sign of the response of the real exchange rate to a supply shock. Using long-run restrictions I find that the exchange rate appreciates on impact. This is in contrast to the predictions of the model. According to the theoretical model the real exchange rate should depreciate in the short-run. I use this prediction to estimate the baseline VAR with sign restrictions and consequently the exchange rate depreciation after a supply shock is obtained by construction.

In this subsection I assess the robustness of my results to estimating the VAR with sign restrictions without imposing a sign on the response of the real exchange rate to a supply shock.

Figure 6 compares the impulse responses of the baseline model estimated with sign restrictions shown in Figure 3 (hatched lines) with the ones obtained using the alternative sign restrictions (solid and dashed lines). Overall the impulse responses mirror those obtained for the baseline VAR except that now the exchange rate tends to appreciate after a supply shock but the appreciation is not significant.

Table 2 presents the variance decomposition of the real exchange rate using the alternative sign restrictions (SR2). The results are very similar to the ones using the baseline sign restrictions (SR1).

7.2 Sub-Sample analysis

Financial markets in the G7 countries have witnessed substantial changes over the sample period. For example, capital controls have been 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 in two sub-samples (1976:1-1989:4 and 1990:1-2006:4) and estimate the impulse responses for each of them in order to check whether regime shifts change the results. The advantage of breaking the sample is that I avoid 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 split between the two samples given that it could be defined as the starting point for the recent wave of financial globalization.¹⁰

Impulse responses are shown in figures 7.A. and 7.B. For the first sub-sample, impulse responses mirror those obtained for the full sample. The only difference that emerges is that the short-run contribution of nominal shocks increases with respect to the baseline case (the variance decomposition is not shown here to preserve space but is available upon request).

In the second sub-sample, nominal shocks appear to have a weaker effect on the real exchange rate. In fact, the real exchange rate shows no statistically significant reaction to nominal shocks after 5 quarters. The relative contribution of nominal shocks to real exchange rate fluctuations declines for this period.

7.3 Alternative exchange rate measure

I test for the sensitivity of the results by using the real effective exchange rate taken from the US Federal Reserve Board Statistics instead of the one sourced from the IFS. This index is CPI-based and includes a wider set of countries. Figure 8 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 (dashed lines). The impact of each of the shocks 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 here to preserve space but available upon request) is very similar to that of the baseline specification.

¹⁰Some authors have chosen 1982 as a the split between subsamples (see e.g. Kim, 1999 and Canova and De Nicoló, 2002). I don't analyze the results based on this break because the sample size becomes too small.

7.4 Extended Model

In this section I include the interest rate differential (i_t) into the baseline VAR to check for the robustness of my results. Thus, the vector of endogenous variables is now $Y_t' = [\Delta y_t \quad \Delta p_t \quad i_t \quad \Delta q_t]$. My aim is to identify three types of innovations: a supply, a demand and a monetary shock. The restrictions imposed for relative output, relative prices and real exchange rate are the same as in Table 1. I now include additional restrictions on the interest rate differential which are in line with the aggregate supply-demand diagram and also confirm the conventional undergraduate textbook intuition. Firstly, I impose that the interest rate differential does not increase after a supply and monetary shock. Secondly, I restrict the interest rate differential not to decrease after a demand shock. I assume that the restrictions for relative output and relative prices are binding for four quarters ($k = 4$) and that the restrictions for relative interest rate and real exchange rate are binding for one quarter ($k = 1$).

I estimate the model applying the same methodology as the one described in Section 5 for the case of a 4 variable VAR and 3 shocks. I present the results based on the minimization of the distance with respect to the median and the 16th and 84th percentiles error bands.

Figure 9 shows the impulse responses of relative output, relative inflation, relative interest rates and the exchange rate with respect to the three shocks of interest. The responses to a supply and demand shocks are very similar to the ones of the three variable VAR. The only point to highlight is that supply and demand shocks have a temporary impact on the relative interest rate (a supply shock decreases the relative interest rate and a demand shock increases it). In line with the baseline estimation, after an expansionary monetary policy shock the real exchange rate depreciates, reaching a minimum after 3 quarters and then reverts to equilibrium (consistent with PPP). The effects of a monetary policy shock are insignificant after eight quarters. By and large, the behavior of the real exchange rate is consistent with the delayed overshooting hypothesis. I also find that a monetary policy shock has a temporary effect on relative output and relative interest rate differential. By contrast, relative prices show a persistent response to monetary policy shocks.

The contributions of supply, demand and nominal shocks to real exchange rate fluctuations are very similar to those of the three variable VAR, with monetary shocks exhibiting a slightly higher importance in the four variables model. (not presented here for brevity).

7.5 Other Methods: Zero short-run restrictions

In order to gain a further understanding of the sources of real exchange rates fluctuations, it is informative to identify the shocks using other methods. In particular, I examine the impact of monetary shocks using zero short-run restrictions in the same fashion as Eichenbaum and Evans (1995). This identification strategy is different from Clarida-Gali or even the sign restriction approach, but it is illustrative to analyze it and compare the results with the other techniques. I highlight that this identification strategy yields a significant “puzzle”, thus casting doubt on its validity.

Figure 10 shows the impulse responses using the Choleski decomposition. I identify the monetary policy shocks with innovations in the interest rate differential. The VAR model is estimated with the interest rate differential ordered third and the real exchange rate fourth.

The results show that a monetary contraction leads to a temporary appreciation of the real exchange rate. In contrast to the sign restriction method, overshooting occurs on impact. Interestingly, prices go up for two quarters and decrease afterwards. This response of prices to a monetary tightening resembles the so-called “price puzzle” pointed out by Sims (1992). The response of relative output is insignificant over all horizons.

The only point to highlight about the variance decomposition is that according to the recursive approach monetary policy shocks only explain 10% of the movement of the real exchange rate at all horizons.

8 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. A great deal of theoretical literature has focused on confirming this belief. However, the empirical evidence on the role of monetary policy shocks has not given clear cut answers on the link between monetary policy and exchange rate movements. In addition, this work has often been criticized due to the lack of credible identifying assumptions.

This paper has focused on one specific question: How important are nominal and real shocks as drivers of the US real exchange rate? In order to address this question, this paper starts with the estimation of a VAR model with long-run restrictions as in Clarida and Gali (1994). Using this conventional estimation strategy I find that monetary shocks are unimportant to explain exchange rate fluctuations.

I also estimate a VAR model with short-run sign restrictions to identify sup-

ply, demand and monetary shocks. The sign restrictions are also derived from the Clarida-Gali open economy macro model but focus on the short-run predictions of the model. Using this method I get a different perspective on the sources of exchange rate fluctuations. In particular, I find that even though demand shocks have been the main driver of exchange rate fluctuations, monetary shocks explain 47% of the fluctuations in the real exchange rate at a 4 quarters horizon.

These results have important implications. Firstly, they reveal that, in contrast to previous findings, the contribution of monetary policy shocks to real exchange rate movements is high at short horizons. This reconciles the focus of the theoretical literature on monetary policy shocks with the empirical evidence. Secondly, my paper also shows that conventional estimation techniques find a much less relevant role for monetary policy. These results demonstrate how different models may give rise to different results. I emphasize that conventional estimation strategies have a number of shortcomings. I take these findings as being part of the general criticism on the use of arbitrary assumptions when estimating VAR models.

Figure 1. Data

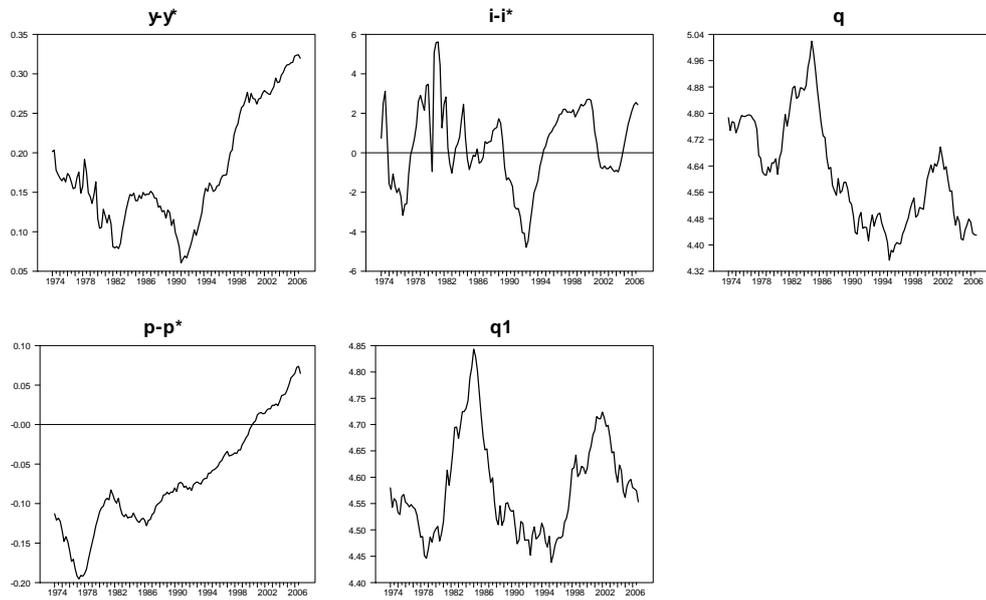
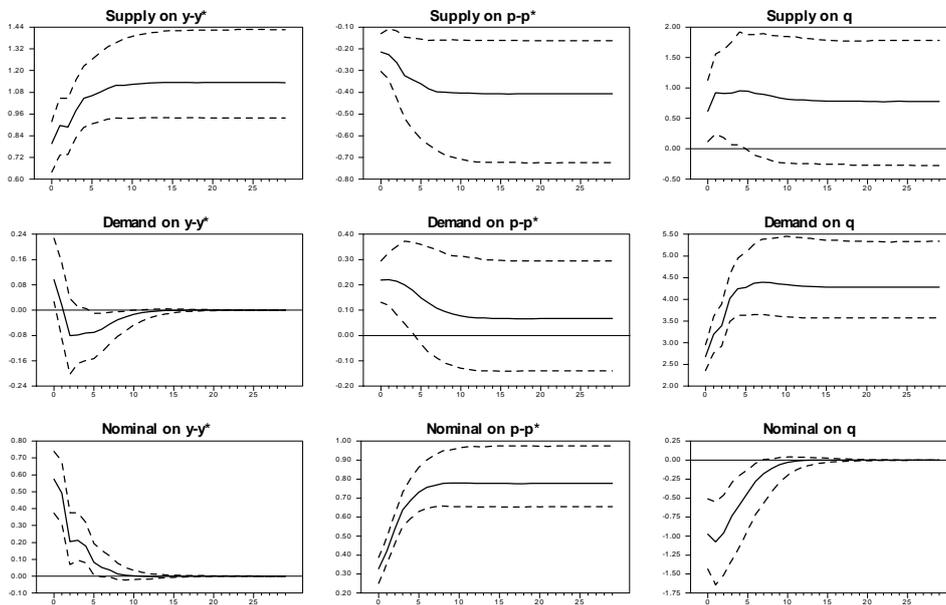
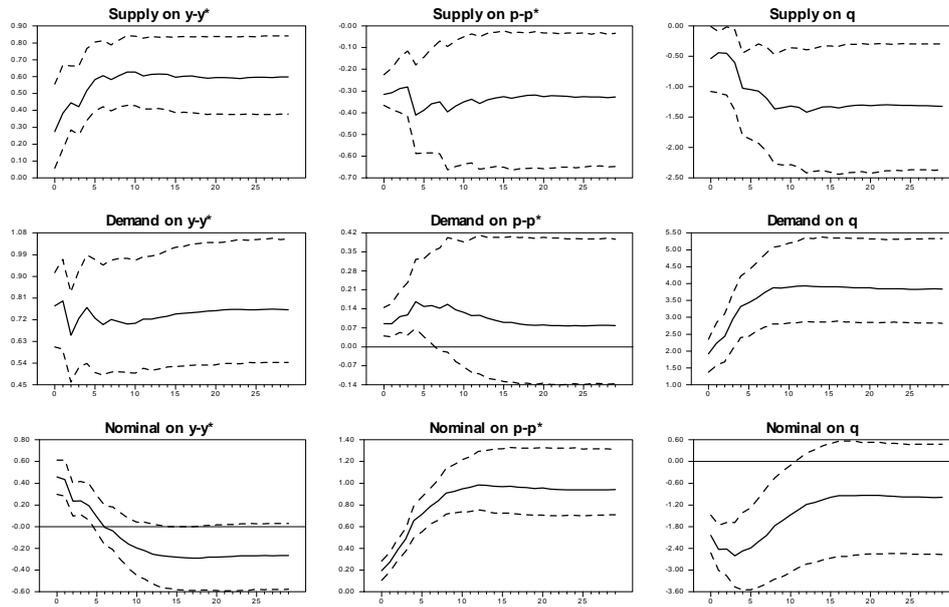


Figure 2. Impulse Responses based on Clarida-Gali identification



Notes: The figure shows the impulse responses to a demand, supply and nominal shocks using the Clarida-Gali identification approach based on long-run restrictions. The solid lines represent the point estimates and the dashed lines represent the 16th and 84th percentiles error bands.

Figure 3. Impulse Responses based on sign restrictions



Notes: The figure shows the impulse responses to a demand, supply and nominal shocks using sign restrictions. The solid lines are calculated based on the minimization of the distance with respect to the median as explained in Section 5. The dashed lines represent the 16th and 84th percentiles error bands.

Figure 4. Structural shocks

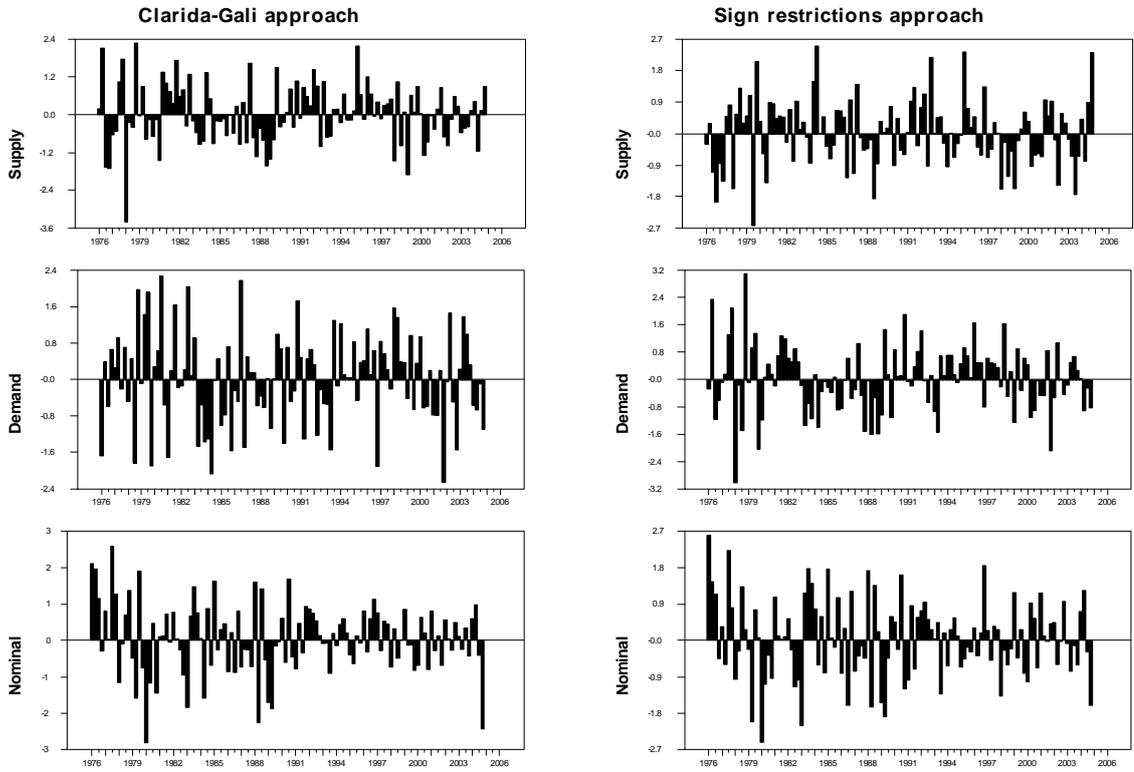


Figure 5. Historical decomposition

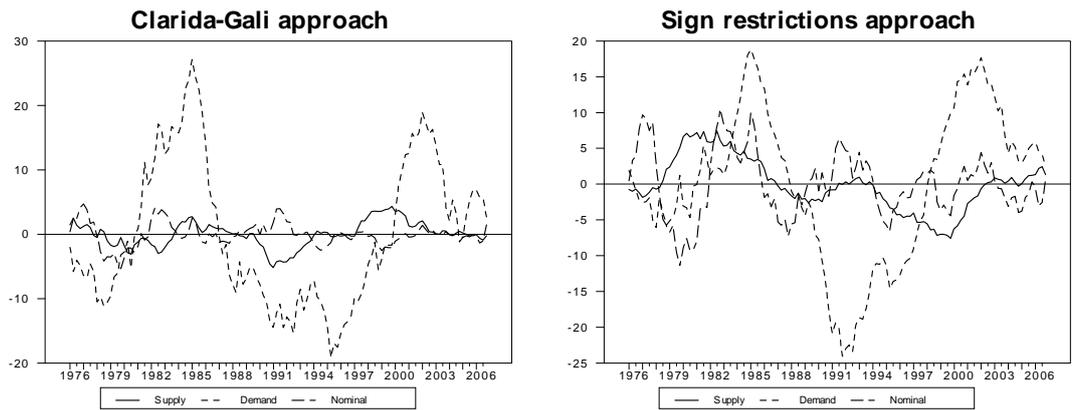
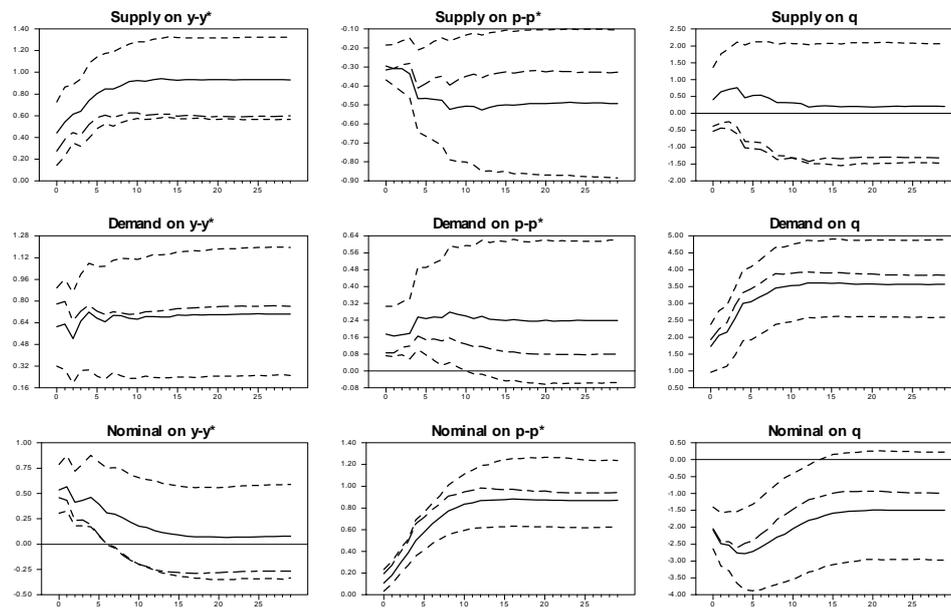


Figure 6. Impulse responses based on alternative sign restrictions



Notes: The figure compares the impulse responses to a demand, supply and nominal shocks using the baseline sign restrictions of Figure 4.3 (hatched lines) with the ones obtained using alternative sign restrictions (solid and dashed lines). The alternative specification relaxes the restriction on the real exchange rate in the case of a supply shock.

Figure 7.A. Impulse responses subsample 1976:01-1989:04

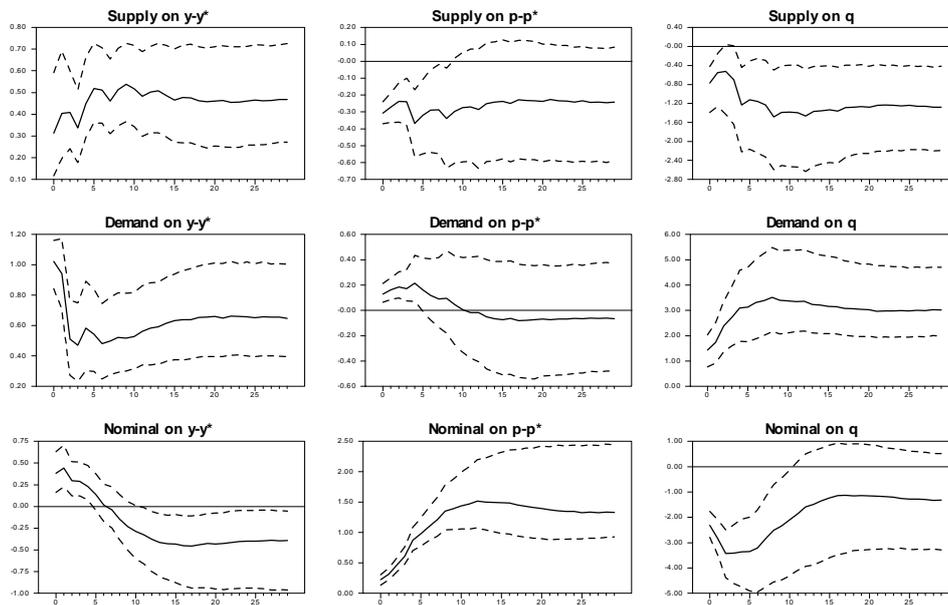
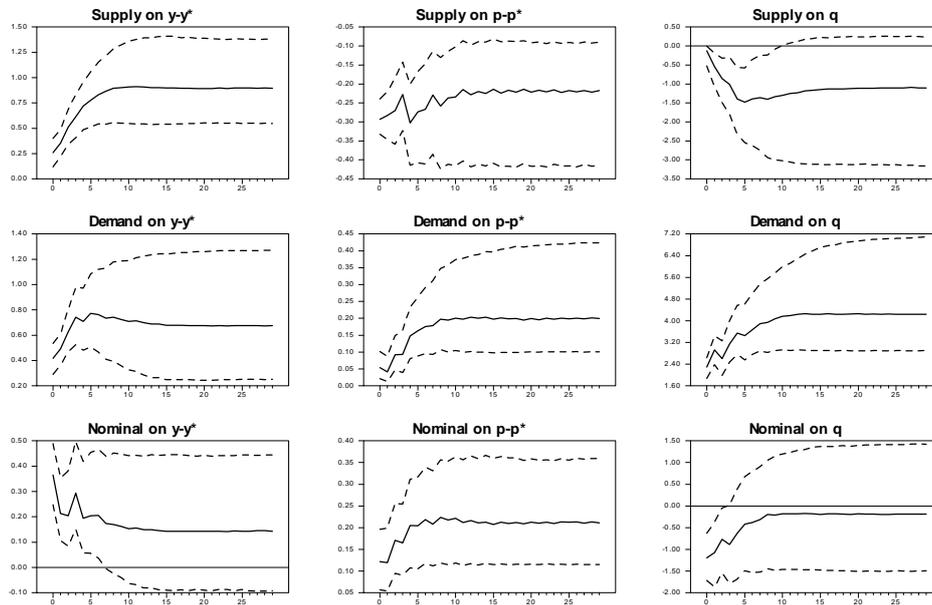
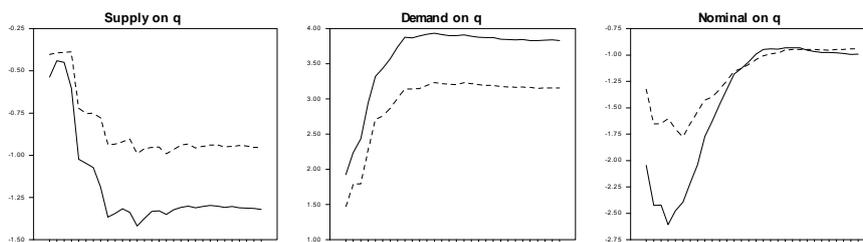


Figure 7.B. Impulse responses subsample 1990:01-2006:04



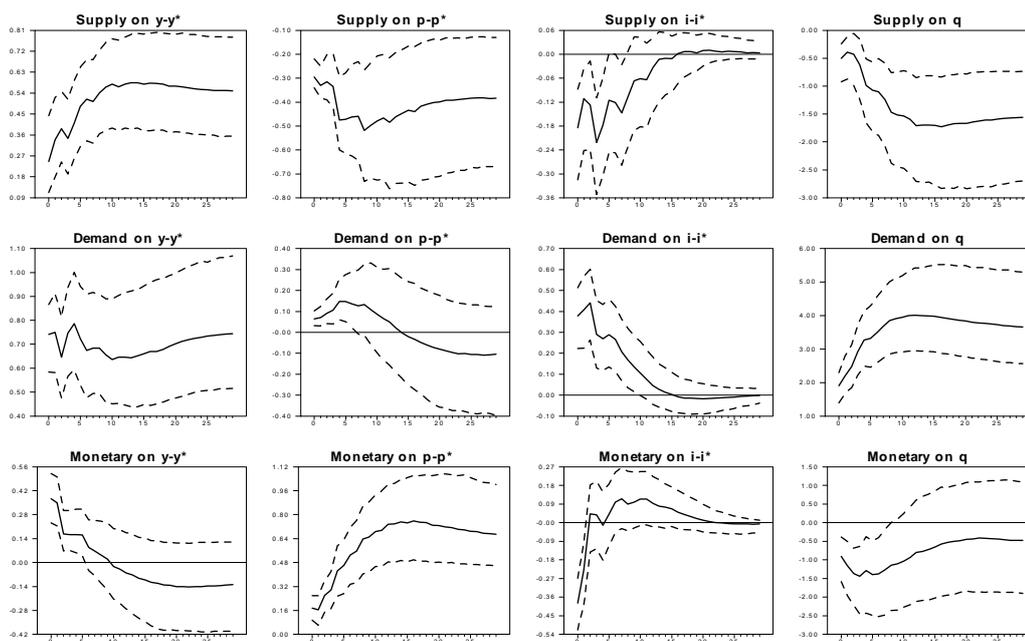
Notes: The figure shows the impulse responses to a demand, supply and nominal shocks using sign restrictions for two subperiods. The solid lines are calculated based on the minimization of the distance with respect to the median as explained in Section 5. The dashed lines represent the 16th and 84th

Figure 8. Impulse responses alternative REER



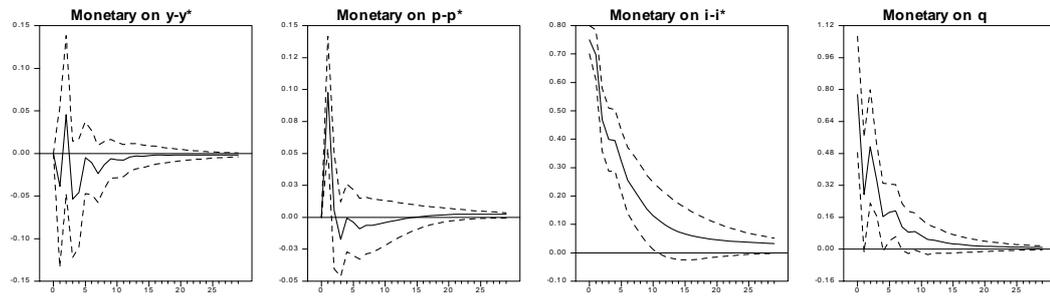
Notes: The figure compares the impulse responses of the real exchange rate to a demand, supply and nominal shocks using the baseline model presented in Figure 3 (solid lines) with the ones obtained when the model is estimated using the real effective exchange rate sourced from the Federal Reserve Board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

Figure 9. Impulse Responses extended model



Notes: The figure shows the impulse responses to a demand, supply and nominal shocks using sign restrictions for the 4 variable VAR. The solid lines are calculated based on the minimization of the distance with respect to the median as explained in section 3. The dashed lines represent the 16th and 84th percentiles error bands.

Figure 10. Impulse responses Choleski decomposition



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

A Appendix

Table A1. Tests for unit roots

	$y - y^*$	$p - p^*$	$i - i^*$	q	$y - y^*$	$p - p^*$	$i - i^*$	q
	Levels				Fist Differences			
ADF ^{AIC}	-1.54 (0.810)	-1.49 (0.827)	-3.22* (0.085)	-2.31 (0.426)	-5.85 (0.000)***	-5.78 (0.000)***	-4.83 (0.000)***	-8.71 (0.000)***
ADF ^{BIC}	-1.49 (0.828)	-0.33 (0.917)	-2.48 (0.335)	-1.81 (0.692)	-10.22 (0.000)***	-4.06 (0.000)***	-8.98 (0.000)***	-8.71 (0.000)***
KPSS	0.29***	1.17***	0.09	0.71**	0.07	0.04	0.05	0.07

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 trend-stationary alternative. The latter tests the null of trend-stationarity. The critical values of the ADF test are -3.15, -3.45 and -4.03 for the 10%, 5% and 1% significance levels respectively. The critical values for the KPSS test are 0.12, 0.15 and 0.22 for the 10%, 5% and 1% significance levels respectively. AIC denotes that the lag length was selected according to the Akaike Criterion and BIC denotes that it was selected based on the Schwartz criterion. In parenthesis are p-values. The sample period is 1976-2006. *, **, *** indicates rejection of the null at 10%, 5% and 1% respectively.

Table A2. Test of cointegrating rank

rank=r	Trace	95% crit. value	p-value	Trace	95% crit. value	p-value
	Baseline Model			Extended Model		
r=0	33.531	42.770	0.315	55.295	63.659	0.214
r=1	18.194	25.731	0.338	31.806	42.770	0.405
r=2	5.886	12.448	0.485	14.984	25.731	0.583
r=3	-	-	-	4.542	12.448	0.667

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 variables of the baseline model are $y - y^*$, $p - p^*$ and q and the variables of the extended model are $y - y^*$, $p - p^*$, $i - i^*$ and q . The sample period is 1976-2006. The VAR model is estimated with 4 lags.

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