

# New Evidence on the Puzzles. Results from Agnostic Identification on Monetary Policy and Exchange Rates.\*

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COMMENTS WELCOME

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## Abstract

Past empirical research on monetary policy in open economies has found evidence of the 'delayed overshooting', the 'forward discount' and the 'exchange rate' puzzles. We revisit the effects of monetary policy on exchange rates by applying Uhlig's (2001) identification procedure that involves sign restrictions on the impulse responses of selected variables. We impose no restrictions on the exchange rate to leave the key question as open as possible. The sign restriction methodology avoids the "price puzzles" of the identification strategies used by Eichenbaum-Evans and by Grilli-Roubini (which are particularly prominent, when using an updated data set). Nonetheless, the puzzles regarding the exchange rates are still there. However, the quantitative features are different. The peak appreciation happens during the first year after the shock for the US-German and the US-UK pair, and during the first two years for the US-Japan pair. This is considerably faster than the three-year horizon found by Eichenbaum-Evans. The exchange rate puzzle becomes more robust than with the Grilli-Roubini recursive identification strategy. Quantitatively, monetary policy shocks seem to have a minor impact on exchange rate fluctuations.

**Keywords:** vector autoregressions, agnostic identification, forward discount bias puzzle, exchange rate puzzle, exchange rates, monetary policy

**JEL codes:**

# 1 Introduction

What are the effects of monetary policy on exchange rates? According to conventional wisdom, there is a sharp conflict between baseline theory and baseline evidence. Dornbusch's (1976) well known overshooting hypothesis predicts that an increase in domestic interest rates relative to foreign interest rates leads to an impact appreciation followed by a persistent depreciation of the domestic currency.

Empirical studies have found different results, however. The most successful studies have followed the lead of Sims (1980), employing vector autoregressions to study these issues in order to sort out the issues of causality of monetary policy shocks, see Sims (1972). Among the advantages of this methodology is the possibility to cleanly formulate and understand the impact of policy changes without violations of the Lucas' critique, see Sims (1982, 1986).

For the issue of the effects of monetary policy on exchange rates, e.g. Eichenbaum and Evans (1995) and Grilli and Roubini (1995, 1996) have found a persistent appreciation for periods up to three years, in contrast to theory. Also Leeper-Sims-Zha (1996) find this result for their larger specifications, see e.g. the impulse response of DOL in the figures 11 and 13, although it is not the main focus there. This finding is known as the 'delayed overshooting puzzle', see figure 1. In particular, this implies a violation of the uncovered interest parity (UIP) condition, and is therefore also often called the 'forward discount puzzle', see figure 2. Note that there may be a forward discount puzzle even without delayed overshooting. Moreover, monetary contractions in the several G-7 countries lead to an impact depreciation or, at least, no significant appreciation of their currencies relative to the US dollar (Sims, 1992, Grilli and Roubini 1995, 1996). This is known as the 'exchange rate puzzle', see figure 3. In addition, the quantitative effect of monetary

policy on exchange rates is far from being clear. The estimated percentages of exchange rate fluctuations due to monetary policy shocks range between 5 and 60 percent (Clarida and Gali (1994), Eichenbaum and Evans (1995), Rogers (1999), Faust and Rogers (2000), Kim and Roubini (2000)).

Recently, this conventional view has come under attack. Most notably, Faust and Rogers (2003) have argued, that one needs to relax dubious identifying assumptions and impose at most rather mild sign restrictions or shape restrictions a priori in order to draw robust conclusions about the impact of monetary policy shocks on exchange rates. They find that no robust conclusions can be drawn regarding the timing of the peak response of the exchange rate, that there is robust evidence in favor of large deviations from UIP due to monetary policy shocks (see figure 2) and that monetary policy shocks may or may not be a cause of exchange rate volatility.

Indeed, we shall see that the identification strategies proposed by Eichenbaum and Evans or by Grilli and Roubini lead to significant "price puzzles" when applied to an updated data set, furthermore calling their original results into questions.

This paper reexamines these issues. We identify monetary policy shocks, using the sign restrictions approach of Uhlig (2005). In particular, we assume that domestic contractionary monetary policy shocks do not lead to decreases in domestic short-term interest rates, increases in domestic prices and increases in domestic monetary aggregates. Hence, we match the conventional wisdom and avoid the price as well as the liquidity puzzle by construction. Crucially, we do not impose any restrictions on the exchange rate to leave the central question as open as possible. We argue that these sign restrictions are plausible because they most directly reflect what economists have in mind (or how economists informally evaluate empirical results) when thinking about monetary policy shocks.

We view this as the continued pursuit of the agenda of Sims (1980).

There, he has argued against the large-scale models popular at that time by stating that “*the connection between ... models and reality - the style in which 'identification' is achieved for these models - is inappropriate, to the point at which claims for identification in these models cannot be taken seriously.*” The advantage of reduced-form vector autoregressions is that it does not need these incredible identification restrictions at all, but for structural vector autoregressions, some identification for the structural shocks is obviously needed. But in line with the quote by Sims, we believe that some of the work using structural VARs similarly are in the danger of employing identification restrictions which are not a priori plausible, while the sign restriction approach allows a much more direct connection with believes derived from theory and the empirical application.

We thus follow the lead of Faust and Rogers (2003), who likewise use sign restrictions to narrow down the set of reasonable identifications for monetary policy shocks. In contrast to these authors, we view the sign restrictions as a means to identifying monetary policy shocks rather than an aid in robustness analysis. Most of the Faust-Rogers restrictions are on-impact only, complemented with very few and hand-selected extra sign restrictions at later horizons. By contrast, we impose sign restrictions on the impulse responses for key monetary policy variables for several periods after the shock. We view this as plausible. This can narrow down the range of possible monetary policy shocks considerably, as has already been argued in Uhlig (1998) discussion of Faust (1998), and indeed this will turn out to be true here as well. On the other hand, we do not impose Faust-Rogers-type shape restrictions that “*the exchange rate response falls between lags 1-2,2-3,3-4,4-6,6-12,12-18,18-36,18-80*”. Rather we wish to leave the response of the exchange rate as agnostically open as possible, since that is the variable of focus.

Following the arguments of Sims and Uhlig (1991), we use a thoroughly Bayesian procedure. Thus, we provide posterior distributions regarding the

parameters of interests - like the time and the size of the peak response - rather than robust 90 percent coverage bands. The sign restrictions imposed take center stage in this paper, as they are key to identification and should be subjected to debate and scrutiny. We also exploit this Bayesian perspective to ask questions concerning the risk a Bayesian investor faces when “betting on the UIP violations in 2. We calculate an implied Sharpe ratio and compare it to Sharpe ratios conventionally observed on e.g. equity markets.

As a benchmark and similar to Faust and Rogers (2003), we apply our identification method to the VAR specifications used by Eichenbaum and Evans (1995) and Grilli and Roubini (1995, 1996). When using conventional identification methods increasing the number of variables in a VAR implies a rising number of assumptions which become increasingly difficult to justify. Adding possibly important variables, while using identification via sign restrictions, is typically rather straightforward, however: thus we do.

Analyzing the US-German, the US-UK and the US-Japanese bilateral exchange rates, we now avoid the price puzzles of Eichenbaum-Evans or Grilli-Roubini by construction. Nonetheless, the delayed overshooting puzzle and the forward discount as well as the exchange rate puzzles are still there and they are sizeable. However, the quantitative features are different. The peak appreciation happens during the first year after the shock for the US-German and the US-UK pair, and during the first two years for the US-Japan pair. This is consirably faster than the three-year horizon found by Eichenbaum-Evans. The forward discount bias puzzle comes with risk. The implied Sharpe ratio for a Bayesian investor can be as high as 2.5, which is five times as high as the annual Sharpe ratio for US stock markets. The exchange rate puzzle becomes more robust than with the Grilli-Roubini recursive identification strategy. Quantitatively, monetary policy shocks seem to have a minor impact on exchange rate fluctuations, which is in contrast to some findings of the previous literature.

We view these results as reconfirming a modified version of the findings of Eichenbaum-Evans and Grilli-Roubini. In essence, their recursive identification strategy was close to correctly identifying monetary policy shocks with the data set they used then, even though that identification strategy no longer seems sensible with an updated data set. Thus, the puzzling behaviour of exchange rate is a feature of monetary policy shocks, once correctly identified. The results of Faust and Rogers (2003) are not in contrast with our results: however, their restrictions are too weak to narrow down the range of reasonable monetary policy shocks sufficiently, thus underscoring the point raised in Uhlig's (1998) discussion of Faust (1998).

The identification method used in this paper, introduced by Uhlig (2005), is related to work by Canova and Pina (1999) and Canova and de Nicolò (2000) who put sign restrictions on impulse response correlations. Like in Bernanke and Mihov (1998) the method concentrates on identifying only the shock of interest rather than aiming at fully identifying the system. Other papers that impose restrictions on impulse responses are Dwyer (1997), Faust (1998) and Gambetti (1999).

A number of papers in the empirical literature concerning monetary policy in open economies employ structural VAR approaches to identify monetary policy shocks instead of using a recursive ordering as Eichenbaum and Evans (1995), Grilli and Roubini (1995, 1996) and Sims (1992) do. E.g. Kim and Roubini (2000) model reaction functions of the monetary authorities and the structure of the economy while Clarida and Gali (1994), Lee and Chinn (1997) and Rogers (1999) use the Blanchard-Quah decomposition, i.e. they distinguish between permanent and transitory effects to identify shocks. Kim (2001) experiments with recursive as well as non-recursive identification schemes to analyze the international transmission of US monetary policy shocks.

In the theoretical literature concerning monetary policy in open economy

settings there are attempts to rationalize the forward discount puzzle. E.g. Gourinchas and Tornell (1996, 2002) explain the delayed overshooting as the interaction of learning about the current state and the intrinsic dynamic response of interest rates to monetary shocks. Studies like e.g. Alvarez, Atkeson and Kehoe (2002, 2003), Chari, Kehoe and McGrattan (2000), or Kollmann (1999) use dynamic general equilibrium models with monetary shocks to explain the behavior of the exchange rate.

The structure of the paper is as follows. After providing a theoretical framework in section 2, section 3 describes our empirical approach. Section 4 reports the results and, finally, section 5 concludes.

## **2 Exchange Rates and Monetary Policy: Some Theory and an Asset Pricing Perspective**

Uncovered Interest Parity (UIP) and the Dornbusch overshooting model are explained well in e.g. Obstfeld-Rogoff (1996), sections 8.2.7 and 9.2. Here, we just provide a brief summary in order to fix notation and to provide a framework for the empirical analysis to follow, complementing both with an asset pricing perspective.

Importantly and as has also been emphasized by Faust and Rogers (2003), we do not consider UIP and the overshooting hypothesis in general, but only conditional on a monetary policy shock. A key question is: how much of a change or deviation from UIP should one expect following a monetary policy shock?

Thus, let  $s_k$  be the impulse response of the log of the exchange rate, understood throughout the paper to be Dollars (“home”) per unit of non-US (“foreign”) currency. Let  $r_k$  and  $r_k^*$  be the impulse response for the US and the foreign short term rates, respectively. This allows the calculation of the



compounded return from investing (or borrowing) at this rate from 0 to  $k$ ,

$$r_{0 \rightarrow k} = \sum_{j=0}^{k-1} r_j, \quad r_{0 \rightarrow k}^* = \sum_{j=0}^{k-1} r_j^*$$

Define the forward discount premium

$$\rho_k = s_0 - s_k + r_{0 \rightarrow k} - r_{0 \rightarrow k}^*$$

which is the gain due to the monetary policy shock (compared to the baseline scenario without that shock) from borrowing foreign currency for the  $k$  periods following the monetary policy shock at the foreign short rate, exchanging it for Dollars, investing it at the US rate, and exchanging it back again in period  $k$ . Note that one can write

$$\rho_k = \sum_{j=1}^k \xi_j$$

where

$$\xi_j = s_{j-1} - s_j + r_{j-1} - r_{j-1}^*$$

is the same gain when executing this hedging strategy only from periods  $j - 1$  to  $j$  following the shock. In particular,  $\rho_j$  stays flat for  $j \geq j^*$ , if  $\xi_j = 0$  for  $j \geq j^*$ . Note furthermore, that the reaction in the impact period is not part of the foreign discount premium. I.e., as long as an investor starts this investment strategy within the “impact month” of the monetary policy shock, he will receive  $\rho_k$ .

(Conditional) uncovered interest parity in the context of our analysis says that one should not be able to make (or loose) money via these hedging strategies, i.e.  $\rho_k \equiv 0$  for all  $k$  or  $\xi_j \equiv 0$  for all  $j$ . This appears to be in conflict with the evidence found in the literature, see figure 2.

The forward discount puzzle is rarely stated in these terms, though (with Faust-Rogers, 2003, being one notable exception). Rather, the forward discount puzzle is an implication of the observations on delayed overshooting.

Let  $p_k^*$  and  $p_k$  be the impulse responses of the log price levels in the foreign country resp. in the US, and let

$$q_k = s_k + (p_k^* - p_k) \quad (1)$$

be the impulse response of the real exchange rate. If one adds to (conditional) UIP the assumptions, that the impulse response for the long run real exchange rate  $q_k$  will converge to zero,  $q_k \rightarrow 0$  due to long run purchasing power parity, that prices are sticky and that the differences in the short rates  $r_j - r_j^*$  slowly reverts to zero following the initial monetary policy shock characterized by a liquidity effect  $r_0 - r_0^* > 0$ , one obtains the Dornbusch (1976) overshooting model, that  $q_0$  should be large and negative and slowly revert back to zero and that the domestic currency should appreciate on impact. However, the empirical literature has found delayed overshooting in response to US monetary policy contraction, see figure 1 and no significant foreign-currency appreciation or even foreign-currency depreciation in response to a foreign monetary policy contraction, see 3.

There are a couple of things to note at this point. First, the Dornbusch overshooting hypothesis requires a number of auxiliary assumptions beyond conditional UIP. Second, even if there is overshooting, conditional UIP might be violated if the quantitative magnitudes do not satisfy 1.

Second, the hedging strategies described above are conditional on a single monetary policy shock only. To literally execute such a strategy in practice, where one wishes to only exploit possible gains from a single monetary policy shock, one would need to "insure" away all other influences such as other contemporaneous and all future shocks until maturity  $k$ .

Third and perhaps most importantly: while many papers in the literature - including Faust and Rogers (2003) - have documented (explicitely or implicitly) significant violations of conditional UIP, this may not suffice for an investor contemplating exploiting this deviation at some date  $t$ . The hedging

position executed for a single dollar at stake is a random variable with payoff in terms of US goods given by

$$X_{t+k} = (1 - e^{\rho_k})e^{(r_{0 \rightarrow k} + p_0 - p_k)} \quad (2)$$

if executed in the “hypothetical” manner of insuring against all other shocks. Think of  $X_{t+k}$  as a component of a portfolio bearing exchange rate risk due to a monetary policy shock. Our aim is to study the price for the risk of this component in isolation.

Let  $e^{m_{t,t+k}}$  be the stochastic discount factor of this investor between  $t$  and  $t+k$ . Standard asset pricing theory implies that

$$0 = E_t[e^{m_{t,t+k}} X_{t+k}]$$

Compare this with the general problem of short-selling any Dollar-denominated asset upon the occurrence of a monetary policy shock at date  $t$ , and investing the proceeds at the short rate  $r_{0 \rightarrow k}$ . Let the random return of that asset between period  $t$  and  $t+k$  be given by  $\tilde{R}_{t+k}$ , and let  $\tilde{\rho}_{t+k} = r_{0 \rightarrow k} - \log \tilde{R}_{t+k}$  be the log excess return. As above, the payoff to this strategy is

$$\tilde{X}_{t+k} = (1 - e^{\tilde{\rho}_{t+k}})e^{(r_{0 \rightarrow k} + p_0 - p_k)}$$

and the asset pricing equation reads

$$0 = E_t[e^{m_{t,t+k}} \tilde{X}_{t+k}] \quad (3)$$

Let

$$\mu_{\tilde{\rho}} = \log E_t[e^{\tilde{\rho}_{t+k}}]$$

be the logarithm of the expected excess return. Let

$$r_{0 \rightarrow k}^{\text{real}} = r_{0 \rightarrow k} + p_0 - p_k$$

be the real short rate. The asset pricing equation (3) can be rewritten as

$$\text{SR}_{\tilde{\rho},k} \equiv \frac{\mu_{\tilde{\rho}}}{\sigma_{\tilde{\rho}}} = -\text{corr}_t(m_{t,t+k}, \tilde{\rho}_{t+k})\sigma_m - \text{corr}_t(r_{0 \rightarrow k}^{\text{real}}, \tilde{\rho}_{t+k})\sigma_{r_{0 \rightarrow k}^{\text{real}}} \quad (4)$$

where  $\text{corr}_t(\cdot, \cdot)$  denotes conditional correlation, where  $\sigma_m$ ,  $\sigma_{\tilde{\rho}}$  and  $\sigma_{r_{0 \rightarrow k}^{\text{real}}}$  are the conditional standard deviations of  $m_{t,t+k}$ ,  $\tilde{\rho}_{t+k}$  and  $r_{0 \rightarrow k}^{\text{real}}$ . This equation defines the Sharpe ratio  $\text{SR}_{\tilde{\rho},k}$ , expressed in terms of log returns. For equity (and the reverse of the strategy described here, i.e. for going long on equity and borrowing at the short rate) and an investment horizon of one year, Sharpe ratios of 0.3 to 0.5 are common, as is well-known from the literature. For a discussion in the context of DSGE models, see e.g. Uhlig (2004).

We thus evaluate the Sharpe ratio  $\text{SR}_{\rho,k}$  for the hedging strategy  $\rho_k$  from the perspective of a Bayesian investor, who is able to “insure” against all other current and future shocks, but remains uncertain about the precise impact of monetary policy shocks on the forward discount premium due to uncertainty regarding the reduced-form dynamics of the economy as well as uncertainty regarding the precise nature of monetary policy shocks.

## 3 Our approach

### 3.1 Methodology

Consider a vector autoregression in reduced form,

$$Y_t = B(L)Y_{t-1} + u_t, \quad E[u_t u_t'] = \Sigma$$

for some vector of variables  $Y_t$ , coefficient matrices  $B(L)$  and a variance-covariance matrix for the one-step ahead prediction error  $\Sigma$ . The key to identification is to represent the one-step ahead prediction error  $u_t$  as a linear combination of orthogonalized “structural” shocks,

$$u_t = Av_t, \quad E[u_t u_t'] = I$$

Traditional identification strategies impose a recursive ordering or structural restrictions on  $A$  or  $A^{-1}$ . Here, we use the methodology of sign restrictions as in Uhlig (2005).

As a consequence, it is not necessary to identify all structural shocks. Identifying a single shock is equivalent to identifying an impulse vector:

**Definition 1** *The vector  $a \in \mathbb{R}^m$  is called an impulse vector, iff there is some matrix  $A$ , so that  $AA' = \Sigma$  and so that  $a$  is a column vector of  $A$ .*

Simple matrix algebra shows that any impulse vector  $a$  can be characterized by

$$a = \tilde{A}\alpha, \quad (5)$$

where  $\tilde{A}\tilde{A}' = \Sigma$  is some decomposition of  $\Sigma$  and  $\alpha$  is an  $m$ -dimensional vector of unit length. Let  $r_i(k) \in \mathbb{R}^m$  be the vector response at horizon  $k$  to the  $i$ -th shock in a Cholesky-decomposition of  $\Sigma$ . Then, the impulse response  $r_a(k)$  for  $a$  is given by

$$r_a(k) = \sum_{i=1}^{\infty} \alpha_i r_i(k). \quad (6)$$

The identifying restrictions we shall impose to identify an impulse vector characterizing monetary policy shocks are that  $(r_a(k))_j \geq 0, j \in \mathcal{J}_+$  and  $(r_a(k))_{j \in \mathcal{J}_-} \leq 0, j \in \mathcal{J}_-$  for some subsets of variables  $\mathcal{J}_+$  and  $\mathcal{J}_-$  and some horizon  $k = 0, \dots, K$ .

We use a Bayesian prior for the reduced form VAR parameters  $(B, \Sigma)$  and an independent uniform prior for  $\alpha$ . The uniform prior for  $\alpha$  assures that the implied prior for  $a$  is independent of the specific decomposition  $\tilde{A}\tilde{A}' = \Sigma$  of  $\Sigma$  and can even be random, as long as the choice of the decomposition is independent of  $\alpha$ .

A Bayesian VAR with 6 lags in levels of the logs of the series has been fitted to the data except for using interest rates directly. No constant or time trend are included. The choice of 6 lags follows the choices made in the literature<sup>1</sup> The prior and therefore the posterior belong to the Normal-

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<sup>1</sup>In fact, much of the evidence becomes considerably weaker, when using 12 lags instead,

Wishart family (see Uhlig (1994) for a detailed discussion of the properties). Results are obtained by taking draws from the posterior for the VAR coefficients  $B$  and draws from the space of possible impulse vectors. Inference statements are based on those joint draws that satisfy the sign restrictions for the impulse responses. We use 500 draws satisfying the restrictions for drawing posterior inferences. We typically show the median as well as the 16% and 84% quantiles of the distribution for the points on the impulse response functions.

For further methodological details, see Uhlig (2005).

### 3.2 Identification of Monetary Policy Shocks in Open Economies

The first choice to be made is the selection of variables. To assure comparability and similar to Faust and Rogers (2003), we shall use the specifications used by Eichenbaum and Evans (1995) as well as Grilli and Roubini (1995, 1996). Both specifications investigate countries pairwise, e.g. the US and Germany or the US and the UK. Typical monetary policy variables like short term interest rates and price levels are included, as are data on industrial production and the exchange rate.

The advantage of the Eichenbaum-Evans or **EE** specification is the inclusion of the ratio of nonborrowed reserves to total reserves as a monetary aggregate, which a number of researchers have argued to be closely related to monetary policy choices, see e.g. Christiano and Eichenbaum (1992), Strongin (1995). The advantage of the Grilli-Roubini or **GR** specification is that it treats the US and the “foreign” country in a more symmetric manner, since typically data on reserves are not available to the same extent for the “foreign” countries.

When using conventional identification methods including more variables

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which we did in a previous version of this paper.

implies a rising number of assumptions which become increasingly difficult to justify. The advantage of identification via sign restrictions is that we do not need many more assumptions. Hence, in a second step, we study the robustness of our results and extend both benchmark VAR specifications by adding possibly important variables, and analyze the resulting **BIG** VAR.

We study three country pairs: the US and Germany, the US and the UK, and the US and Japan. We employ monthly data from 1975.07 to 2002.07. For a detailed description of the variables and the specifications, see appendix 6.1.

### 3.3 Identifying Monetary Policy Shocks

To identify monetary policy shocks we shall impose that domestic price variables as well as monetary aggregates like the the ratio of nonborrowed to total reserves or the money stock do not rise. Thus, “price puzzles” are avoided by construction. We do not impose a sign restriction on the reaction of industrial production, since the results in Uhlig (2005) suggest that there is little evidence that GDP will fall in reaction to a contractionary monetary policy shock.

We furthermore either impose that interest rates do not fall or, alternatively, impose that they do not fall more than the interest rate in the other country. These two choices reflect two prototypes of the game played between the two monetary authorities. In the first, one can think of the central bank as a Stackelberg leader: if both interest rates go up, then it and not necessarily the other central bank was the ultimate cause. In the second, the central bank is viewed as a follower: deviations from the interest rate set by the other central bank are viewed as surprises.

In principle, the game played between two monetary authorities could be rather complicated. There is no really good reason to a priori rule out e.g. a game whereby which central banks alternate in who is using the higher

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US monetary policy shock:		
Specification	<b>EE</b>	<b>BIG</b>
Variables	$Y, Y^*, P, NBRX, i, i^*, S$	$Y, Y^*, P, P^*, M, M^*, i, i^*, S, r, r^*, NBRX$
Restrictions:	$P \leq 0,$ $NBRX \leq 0, i \geq 0$	$P \leq 0, M \leq 0,$ $NBRX \leq 0, i \geq 0$

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foreign monetary policy shock:		
Specification:	<b>GR</b>	<b>BIG</b>
Variables:	$Y, Y^*, P, P^*, i, i^*, S$	$Y, Y^*, P, P^*, M, M^*, i, i^*, S, r, r^*, NBRX$
Restrictions:	$P^* \leq 0,$ $i^* - i \geq 0$	$P^* \leq 0, M^* \leq 0$ $i^* - i \geq 0$

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Table 1: Identification of Monetary Policy shocks.



interest rate or choose other complicated reaction functions. We view the two possibilities investigated here as two particularly plausible benchmarks. To impose the commonly held view that US monetary policy is leading and other countries are following, we shall impose the restriction  $i \geq 0$  when identifying US monetary policy shocks, but  $i^* - i \geq 0$ , when identifying foreign monetary policy shocks. For the **BIG** VAR, one can also think of these choices as reflecting a causal ordering of the monetary policy choices, with the US ordered first.

The identification restrictions imposed are summarized in table 1.

For the restriction horizon, we have used  $k = 0, \dots, K = 5$ , i.e. half a year, throughout. Choosing shorter restriction leaves too much room for spurious effects, while imposing a longer horizon imposes an implausibly long duration for the liquidity effect. Uhlig (2005) contains some discussions how results vary when applying the methodology to the analysis of monetary policy shocks in US data. There, output seems to rise rather than fall, if the restriction horizon is extended to one or even two years. This turns out to be even more forcefully true here. It might be interesting to understand more deeply, why this might be the case, but that discussion would be beyond the scope of this paper. We therefore have chosen to omit results for other choices for  $K$ .

## 4 Results

### 4.1 US monetary policy shocks

For a US monetary policy shock, figure 4 contains the key results, while figures 5 to 15 contain further details.

We have plotted all the impulse response functions for our benchmark identification in figure 5 (two pages). The vertical lines in some of the impulse response diagrams denote the sign restrictions we have imposed. Note that

- by construction - there is no liquidity puzzle and no price puzzle. In line with Uhlig (2005), we find no significant effect on real output. We believe that these are reasonable looking results, and shall thus turn to a discussion of the results regarding the exchange rate response.

The first line of figure 4 shows the impulse response of the real exchange rate to a US monetary policy contraction and should be compared to figure 1. The second line shows the posterior distribution of the peak appreciation, i.e. the distribution for the month containing the lowest point of an impulse response drawn from the posterior and shown in the first line. The posterior distribution for the peak of real exchange rate appreciation is compared to the posterior for the peak of the nominal exchange rate appreciation in figure 6: the differences are small.

The results show that the US-German and US-UK bilateral real exchange rates appreciate until approximately the seventh to twelfth month and then depreciates in our benchmark identification. Note that there is a longer delay in the US-Japan case, with most of the distribution centered around months 8 to 24. I.e., our results suggest that there is evidence of a delayed overshooting, in contrast to Kim and Roubini (2000) and Faust and Rogers (2003), but that the delay is considerably shorter than the three-year horizon found by Eichenbaum and Evans (1995). The delay is a matter of months, not a matter of years.

It could be that the posterior distribution for the peak months gives misleading results, if e.g. impulse responses showing mild appreciations peak early and those with strong appreciations peak late. That this is not the case can be seen from the posterior joint distributions of the peak and its altitude both for the real and the nominal exchange rate in figure 7.

Figure 8 concentrates on the first 24 months after the shock and shows the distribution for the real exchange rate  $q(k)$  as well as the distribution for the beyond-impact change of the real exchange rate  $q(k) - q(0)$ . Apparently,

the posterior is rather sharply peaked for the US-Germany and the US-UK case. It is rather remarkable that there is a sharp peak in the posterior distribution for the *change* of the exchange rate compared to the on-impact response, i.e. for  $q(k) - q(0)$  for the US-Germany and the US-UK pair, with greater diffusion in the US-Japan case. After the on-impact change, the exchange rate drops by a further one percent for US-Germany within 7 months and for the US-UK within 2 to 5 months. There is also considerable mass on this event in the US-Japan case, but there is additional mass on early and mild as well as late and somewhat stronger additional appreciations.

Figure 9 compares the impulse response functions for the three identifications we investigated, namely the two sign restriction identifications in the **EE** and the **BIG** specification as well as the original recursive Eichenbaum-Evans identification in the **EE** specification. Note that there is only a minor difference between the result of the **EE** and **BIG** specification, showing that our results are fairly robust and the identification method indeed easily generalizable, as claimed above. Also, the evidence looks rather similar across the three country pairs. The difference to the results from the Eichenbaum-Evans identification is considerable, though.

The reason for this difference can be understood from figure 10 (two pages), where we have compared all the impulse response functions for the variables of the **EE** specification in the US-UK case. Note in particular the huge price puzzle emerging for the Eichenbaum-Evans identification. We believe that this strong and long positive reaction of the price level casts considerable doubt on this identification strategy and therefore on the results for the exchange rate response (and we doubt that Eichenbaum and Evans would have stuck to this strategy in light of the new data and the new results). Obviously, there are additional impulse response functions for the **BIG** specification and similar comparisons could be made for the US-Germany and the US-Japan case. We show the impulse responses for the

**BIG** specification in figure 15 (two pages). These do not change the key insights, however.

The third and forth line of figure 4 shows the impulse response for the forward discount premium and the resulting Sharpe ratio for a Bayesian investor. A comparison for our three identification procedures can be found in figures 11 and 12. The impulse response for the forward premium should be compared to the theory figure 2. The results can be described as follows. There is a forward discount premium, but there is also considerable uncertainty regarding its size or whether it is even positive. When taking this uncertainty into account and calculating the Sharpe ratio, one find values between 1 and all the way up to 2.5 for the **EE** specification and investment horizons for up to two years and values around 0.5 to 1.7 for the **BIG** specification, due to the additional posterior coefficient uncertainty. The results are quite a bit higher for the Eichenbaum-Evans recursive identification: there, the Sharpe ratio reaches values near 4 before slowly moving back to zero. We do not view their recursive identification as plausible, however, as we have argued before<sup>2</sup>. We therefore conclude, that there is indeed a sizeable forward discount bias puzzle, offering rewards to risk which exceed the corresponding annual US stock market Sharpe ratio by a factor of up to 5, but that the reward for risk is not quite as extreme as suggested by the recursive identification by Eichenbaum and Evans.

We similarly investigate the forward premium  $\xi_k$  for one-month holding periods in figure 13 and the corresponding Sharpe ratios in figure 14. The numbers are very similar, and thus do not change the insights.

We conclude from this that the reward for the risk of betting on violations

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<sup>2</sup>Nonetheless one may wonder, which types of shock are the cause of these reactions. There must be *some* shocks which generate these kinds of high Sharpe ratios. Here we only argue, that they are not shocks to monetary policy. If they are due to e.g. restrictions of capital movements or changes in taxation, these high Sharpe ratios may not be “exploitable” for smart investors. Investigating this issue further is surely interesting but beyond the scope of this paper.

of the uncovered interest parity is higher by a factor but not by an order of magnitude, compared to Sharpe ratios typically calculated for asset markets. It may be puzzling why financial markets offer such a high reward for risk in general. While the market for foreign exchange offers even higher rewards, it is not drastically different from other asset markets in that respect.

The last line of figure 4 concerns the variance decomposition of the movements of exchange rates: we shall discuss this together with the variance contributed by foreign monetary policy shocks in subsection 4.3 .

## 4.2 Foreign monetary policy shocks

For a foreign monetary policy shock, figure 16 contains the key results, while figures 17 to 23 contain further details.

The first line of figure 16 shows the impulse response of the real exchange rate and should be compared to the theory picture 3. The second line shows the posterior probability for that impulse response function to be positive in any given month. A comparison of that posterior distribution for real versus nominal exchange rates is given in figure 17: the differences are minor.

The results show that, again, there is a gradual response only, i.e. a delayed overshooting, and that there is considerable probability for the exchange rate to appreciate rather than depreciate on impact. The results here are rather robust across countries. This is in contrast to Grilli and Roubini (1995, 1996) who find the exchange rate puzzle for Germany but not for the UK.

Figure 18 compares the response of the real exchange rates for the sign restriction approach in the **GR** as well as the **BIG** specification to the results from the recursive Grilli-Roubini identification. The Grilli-Roubini identification leads to very different results, depending on the country pair, whereas the results for the sign restriction approach look considerably more alike. In particular, the behaviour of the US-Germany pair looks puzzling in the

Grilli-Roubini identification.

Figures 19 and 20 provide further comparisons between **GR** and the original Grilli-Roubini identification on the distribution of the sign of the response for the real exchange rate as well as the price level. As one can see, there is more variation in the results regarding the sign distribution for the original Grilli-Roubini identification. There also is a considerably large price puzzle. We believe that this casts doubts on the results obtained with their original identification strategy.

Figure 21 compares the impulse responses of all the variables (and variables derived from it) for the three specification and the US-UK pair: we have actually chosen the pair which should be most favorable to the Grilli-Roubini identification, based on the rather “orthodox looking results of the third line in figure 18. Note again, however, that there is a huge price puzzle for the Grilli-Roubini identification: foreign prices keep on rising for a long time before returning to zero, following a contractionary foreign monetary policy shock. For our sign restriction approach, this price puzzle is avoided by construction.

There are more impulse response functions available for the sign restriction approach both for the **GR** and the **BIG** specification. They are shown in figures 22 as well as 23 (two pages) for completeness and to emphasize, that our identification strategy is reasonable.

The last line of figure 16 concerns the variance decomposition of the movements of exchange rates, which we discuss next in subsection 4.3 .

### 4.3 Monetary Policy Shocks and Exchange Rate Volatility

Figure 24 contains the variance decomposition for the exchange rate movements explained by a US monetary policy shock, while figure 25 contains the fraction of exchange rate variation explained by foreign monetary policy

shocks. In figure ?? we use the **BIG** specification to plot the joint contribution of US and foreign monetary policy shocks to the volatility of exchange rates.

US monetary policy shocks account for somewhere between 2 and 10 percent of the exchange rate fluctuations at the median estimate, independent of the horizon and country, for both sign restriction specifications. The number is smaller at most horizon for the Eichenbaum-Evans identification except for a rather sharp peak 6 to 8 months after the shock.

For foreign monetary policy shocks, figure 25 delivers a similar result with a somewhat broader range for the numbers. The Grilli-Roubini specification delivers very different results, depending on the country pair chosen: less than 5 percent (at the median) is explained in the US-German case, nearly 30 percent is explained at a horizon one to three years out in the US-UK case and there is a sharp peak one year out in the US-Japan case.

These rather different behaviours for the Grilli-Roubini specifications as well as the sharp peaks in the Eichenbaum-Evans specifications may be viewed as a further oddity, when applying these conventional identification strategies, which are avoided with the sign restriction approach.

For the sign restriction approach, the 86

All in all, our results are in contrast to Eichenbaum and Evans (1995) who estimated a percentage of 42, 26 and 23 for Germany, UK and Japan at lags 31 – 36. Other examples for studies which find that monetary shocks have a substantial contribution in explaining real exchange rate fluctuations are Clarida and Gali (1994) and Rogers (1999). Our result that monetary policy shocks do not seem to be important for exchange rate fluctuations is compatible with the weaker result of Faust and Rogers (2003) who state that the percentage might be anything between 8 and 56. However, since we are imposing more identifying assumptions, we find a narrower range than they do. Our results are also compatible with Kim and Roubini (2000) who

estimate a percentage of 5, 16 and 17 at long horizons.

## 5 Conclusions

This paper has estimated the effects of monetary policy shocks on the US-German, the US-UK and the US-Japanese bilateral exchange rates by applying an agnostic identification method recently proposed by Uhlig (2005). A priori theorizing has been made explicit by imposing sign restrictions on the impulse responses of selected variables for a certain period following the shock. We view this identification strategy as a consequent pursuit of the agenda put forth by Sims (1980) of avoiding incredible identifying restrictions. In particular, we impose no or only weak restrictions on the exchange rate to leave the key questions as open as possible.

We have followed conventional wisdom and have assumed that domestic contractionary monetary policy shocks do not lead to decreases in domestic short-term interest rates, increases in domestic prices and increases in domestic monetary measures.

It has turned out that the evidence on the delayed overshooting, the forward discount and the exchange rate puzzles remain, but that their quantitative properties change. Applying our identification scheme to a benchmark VAR we have found that with 2/3 probability the US-German and the US-UK exchange rates appreciate for about 9 months and then depreciate. Only the US-Japanese exchange rate shows a delayed overshooting up to 24 months.

Regarding the forward discount puzzle, we evaluate the risk inherent in exploiting the forward discount premium by calculating the Sharpe ratio for a Bayesian investor, and find values in line considerably higher but not dramatically different from those found on asset markets. We do find a robust exchange rate puzzle, in the sense that there is a rather a stable and incon-



clusive pattern across several countries and specifications, in contrast to the results from standard identification procedures. Quantitatively, monetary policy shocks seem to have a minor impact on exchange rate fluctuations, again in contrast to some of the literature.

Our results are sharper than the results of Faust and Rogers (2003) who find that the behavior of the exchange rate is very sensitive to different identification schemes. The sharpening is due to imposing more (and we argue, sensible) identifying restrictions rather than the rather loose on-impact restrictions in Faust and Rogers. In contrast to these authors, we do not restrict the reaction of the exchange rate, though, and view our methodology as providing posterior distributions, rather than a sensitivity analysis.

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## 6 Appendix

### 6.1 Data

variable	description	source
$y$	index of US industrial production	IMF Washington, line 66
$p$	US consumer price index	IMF Washington, line 64
$nbrx$	US non-borrowed reserves/total reserves	Fed. Reserve Bank St. Louis
$m$	US money supply M1	Fed. Reserve Bank St. Louis
$i$	US 3-months treasury bill rate	IMF Washington, line 60c
$r$	US 10-year government bond yield	IMF Washington, line 61
$y^*$	foreign industrial production	IMF Washington line 66
$m^*$	foreign money supply M1 (UK: M0)	IMF Washington, line 39, (UK: Bank of England)
$p^*$	foreign consumer price index	IMF Washington, line 64
$i^*$	foreign 3-months treasury bill rate	IMF Washington, line 60c
$r^*$	foreign 10-year government bond yield	IMF Washington, line 61
$s$	nominal exchange rate in dollar per foreign currency	Fed. Reserve Bank St. Louis
$q$	real exchange rate per foreign currency	derived
$\rho$	forward discount premium (accumulated)	derived
$\xi$	forward discount premium (for one period)	derived

## 6.2 Specifications

The following VAR specifications have been used:

**EE** (Eichenbaum-Evans):  $Y, Y^*, P, NBRX, i, i^*, S$

**GR** (Grilli-Roubini):  $Y, Y^*, P, P^*, i, i^*, S$

**BIG** (big VAR specification:)  $Y, Y^*, P, P^*, M, M^*, NBRX, i, i^*, r, r^*, S$

## 6.3 Identifications in the literature

The following is a short description of some of the identifications used in the literature

Eichenbaum-Evans: assume a recursive ordering with the variables as listed in the **EE** specifications above. They identify a US monetary policy contraction with a rise in  $NBRX$ . **CHECK THIS**

(Grilli-Roubini: **TO BE COMPLETED.**

Faust-Rogers: **TO BE COMPLETED.**



## **6.4 Figures**

### **6.4.1 Theory**

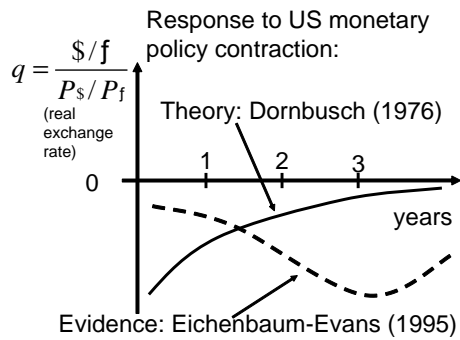


Figure 1: A stylized representation of the forward discount puzzle

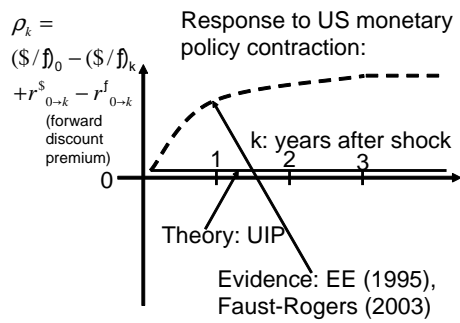


Figure 2: A stylized representation of the forward discount puzzle

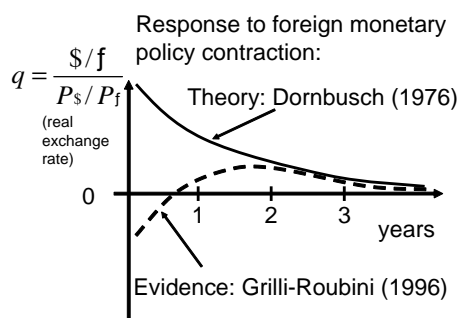


Figure 3: A stylized representation of the exchange rate puzzle

#### 6.4.2 Responses to a US monetary policy contraction.

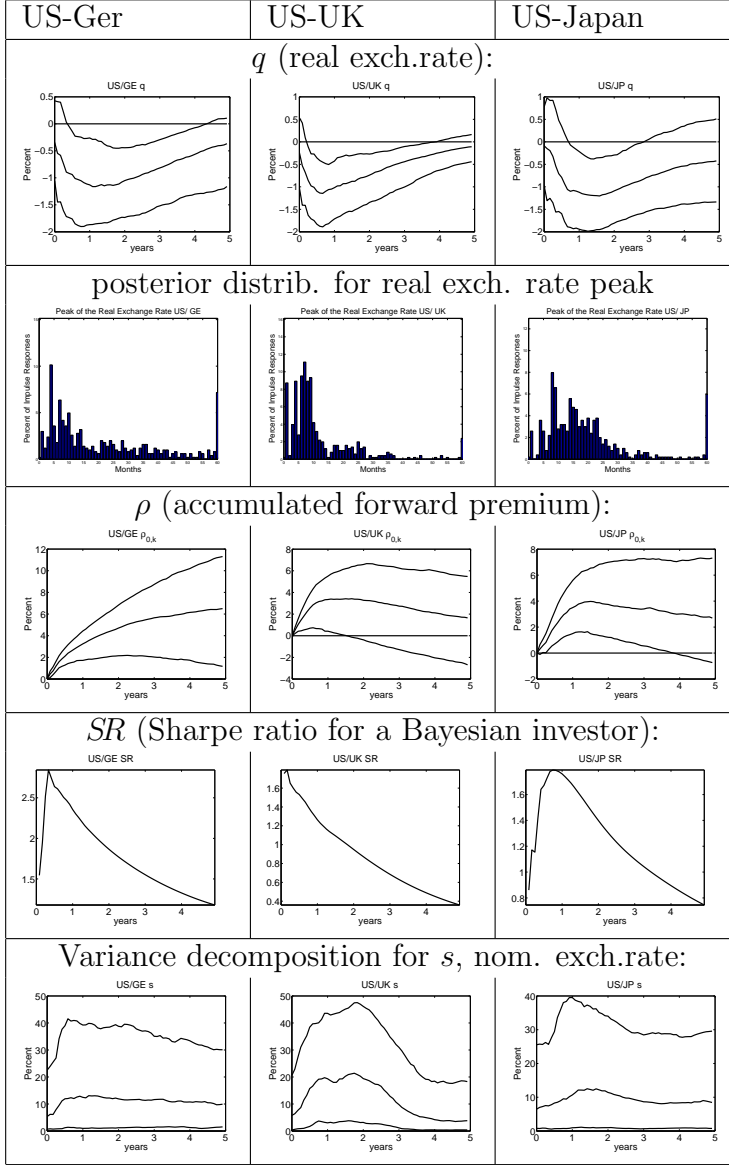


Figure 4: Key results for a **US monetary policy shock**, using the benchmark identification in the **EE** specification. Shown are the results for the response of the real exchange rate  $q$ , the posterior distribution for its peak, the response of the (accumulated) forward premium, the Sharpe ratio of a Bayesian investor and the variance decomposition of the nominal exchange rate.

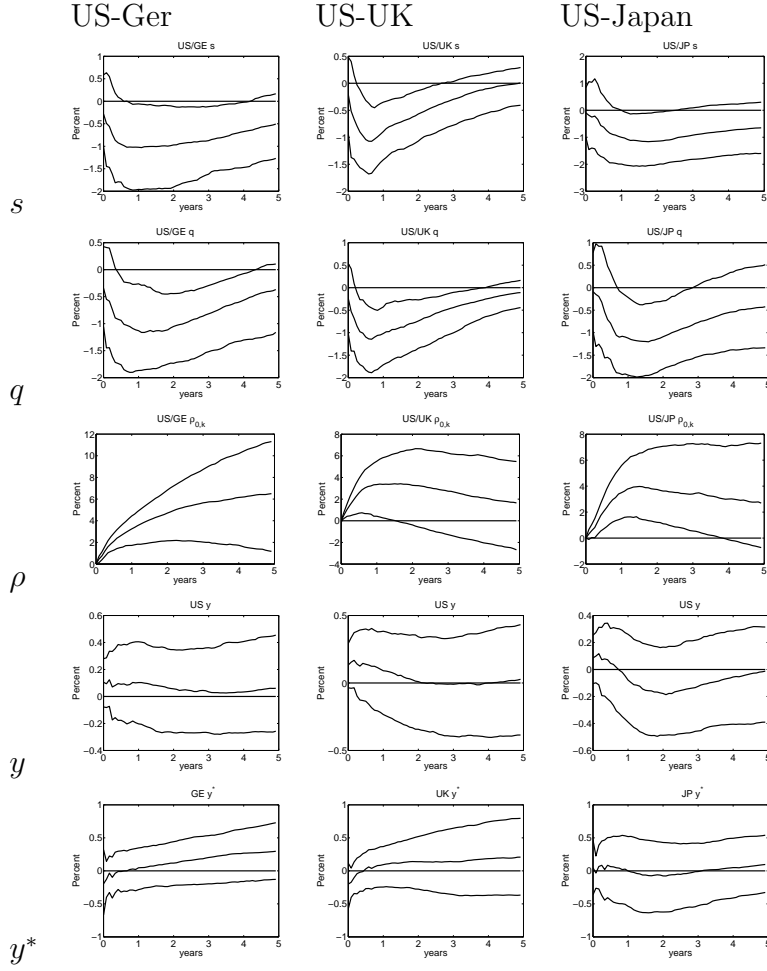


Figure 5: Part 1 of the impulse responses for a US monetary policy contraction, using the benchmark identification for the **EE** specification. Note that there is little difference between the response of the nominal and the real exchange rates.

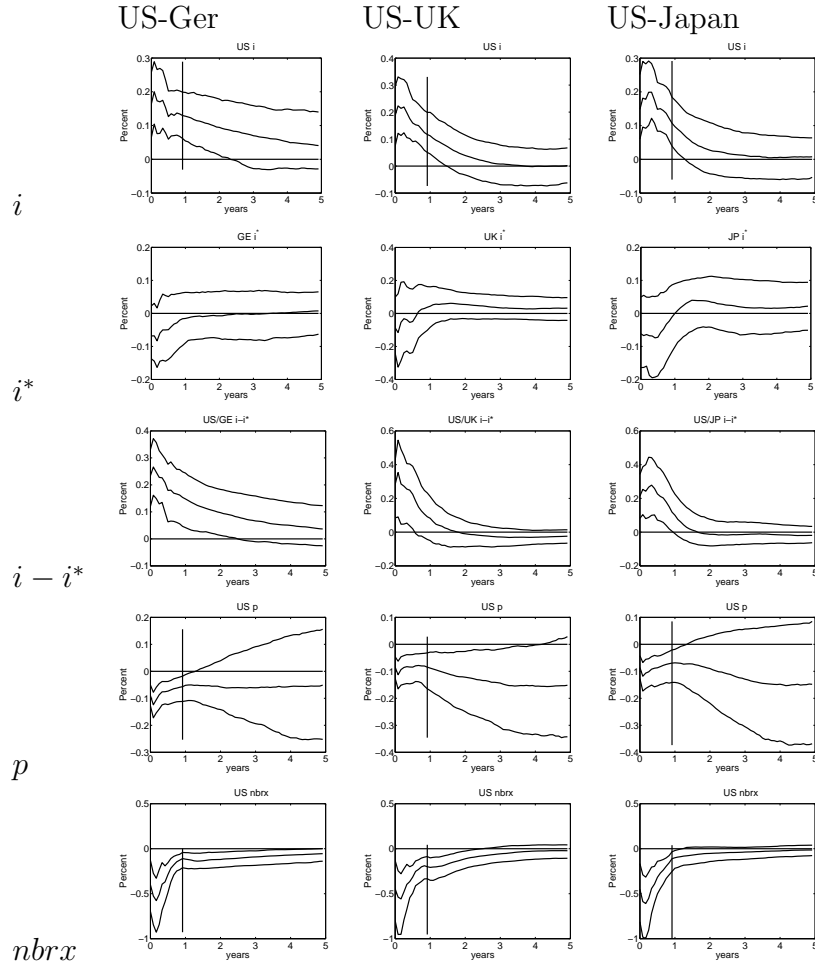


Figure 5 continued. Part 2 of the impulse responses for a US monetary policy contraction, using the benchmark identification in the **EE** specification.

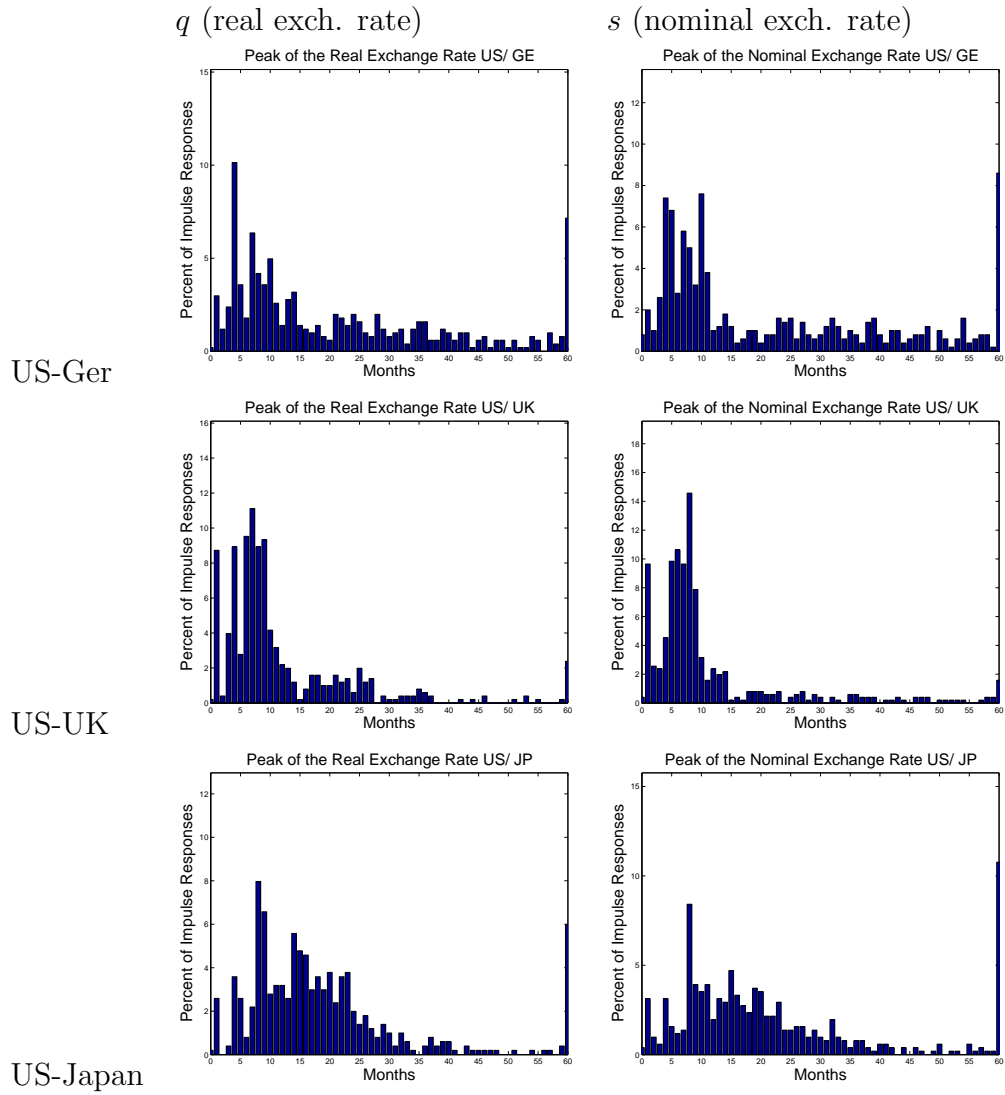


Figure 6: The distribution of the peak response of the real and the nominal exchange rate for the benchmark identification, **EE** VAR.

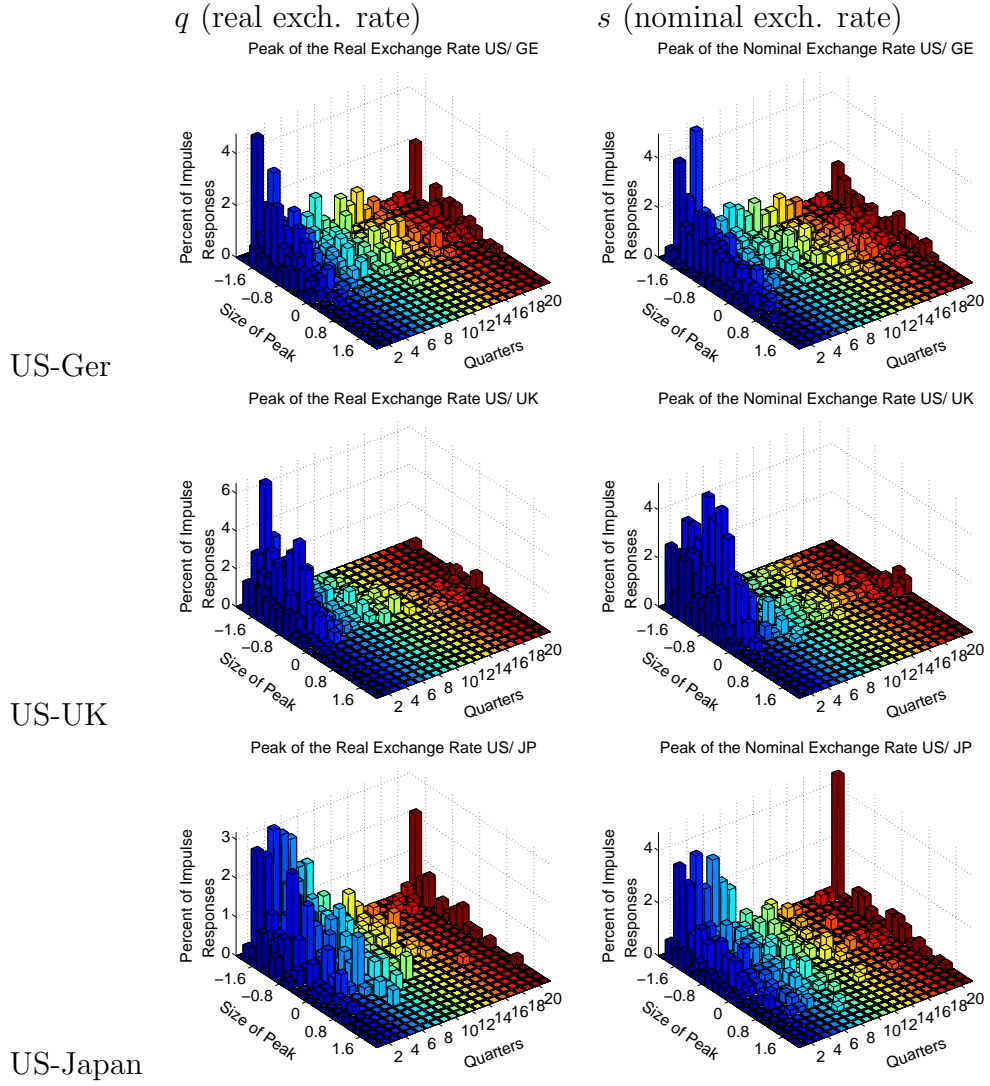


Figure 7: The axis shows quarters. Posterior distribution for the size and location of the peak appreciation of the real and nominal exchange rate, conditional on a US monetary policy contraction, for the benchmark identification, **EE** VAR.



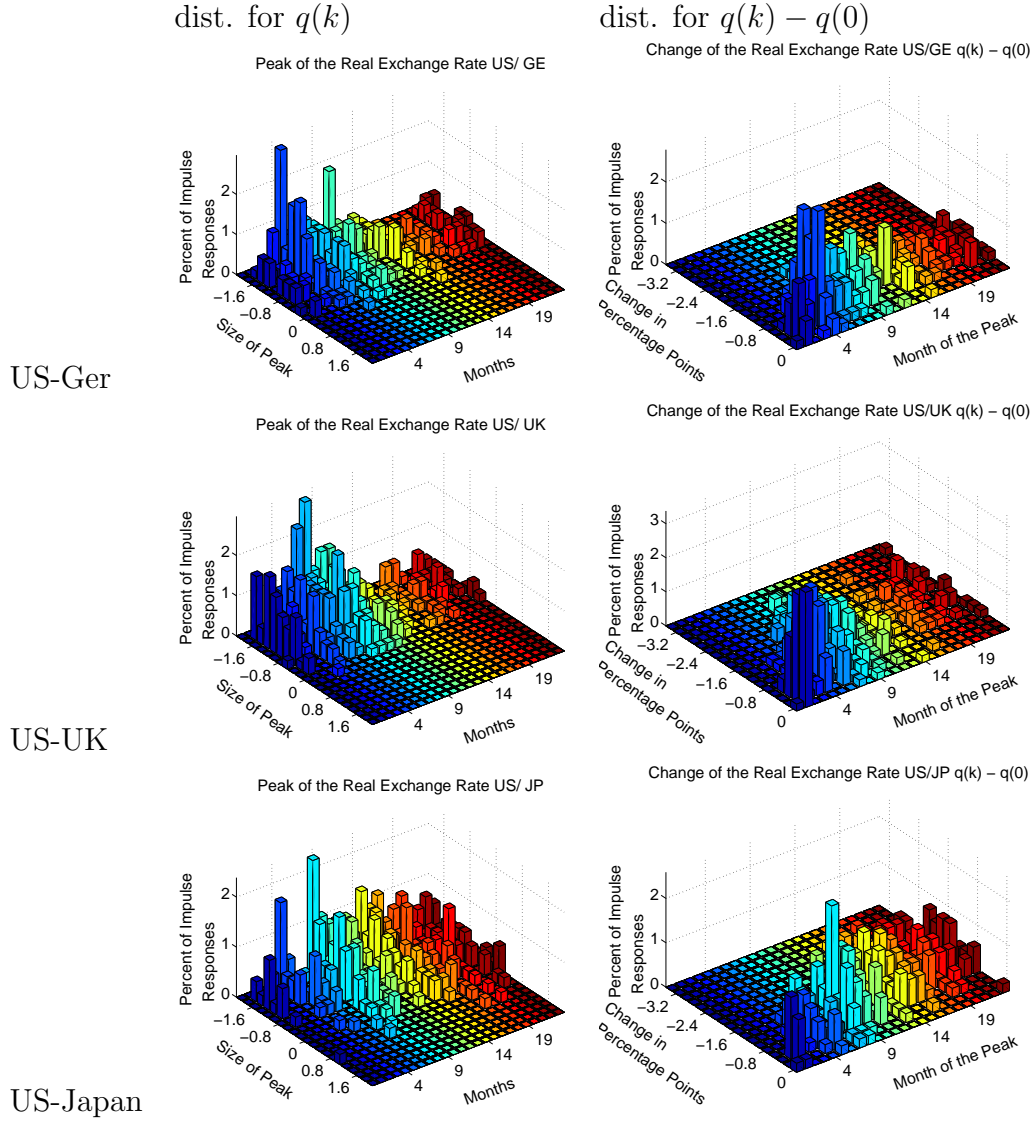


Figure 8: **Real exchange rate  $q$** . Closer examination of the first 24 months following the shock. The axis shows months. Posterior distribution for the size and location of the peak appreciation  $s(k)$  as well as for the change relative to the impact reaction  $s(k) - s(0)$ , conditional on a US monetary policy contraction, for the benchmark identification, **EE** VAR.

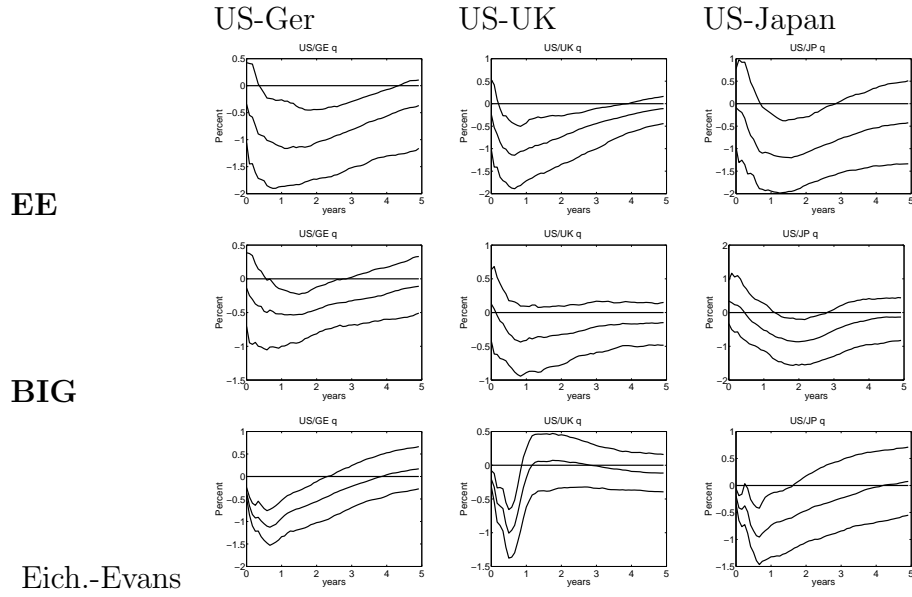


Figure 9: Impulse response function for the real exchange rate, conditional on a US monetary policy contraction. We compare the benchmark specification identification in the **EE** VAR specification to the identification in the **BIG** VAR to the original Eichenbaum-Evans recursive identification.

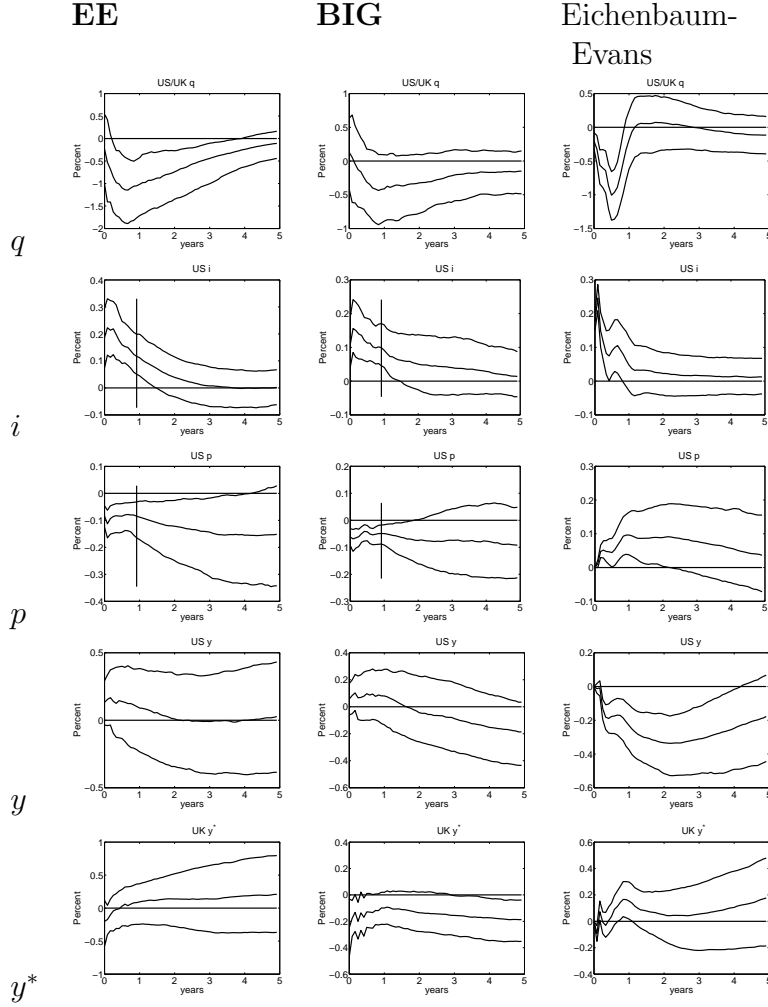


Figure 10: Comparison of results. For a US monetary policy contraction and the US-UK country pair, we compare the benchmark identification in **EE** to the **BIG** specification and the original identification of Eichenbaum-Evans. Note the “price puzzle” in the Eichenbaum-Evans column.

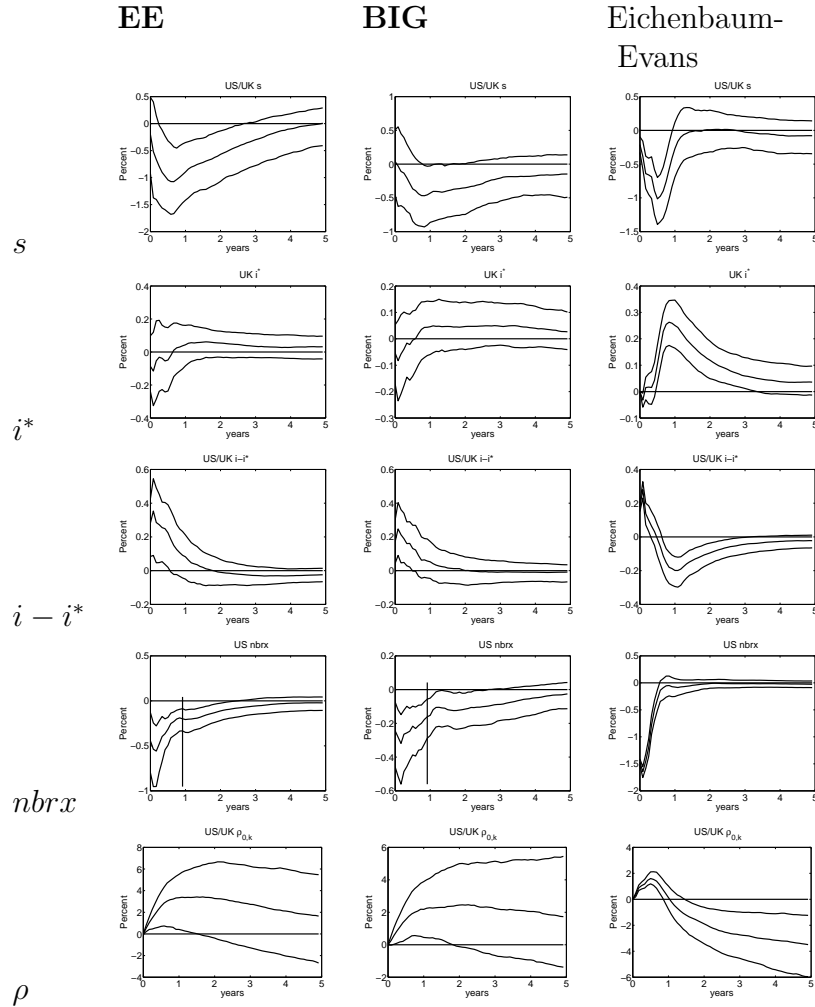


Figure 10 continued. Part 2 of the comparison.

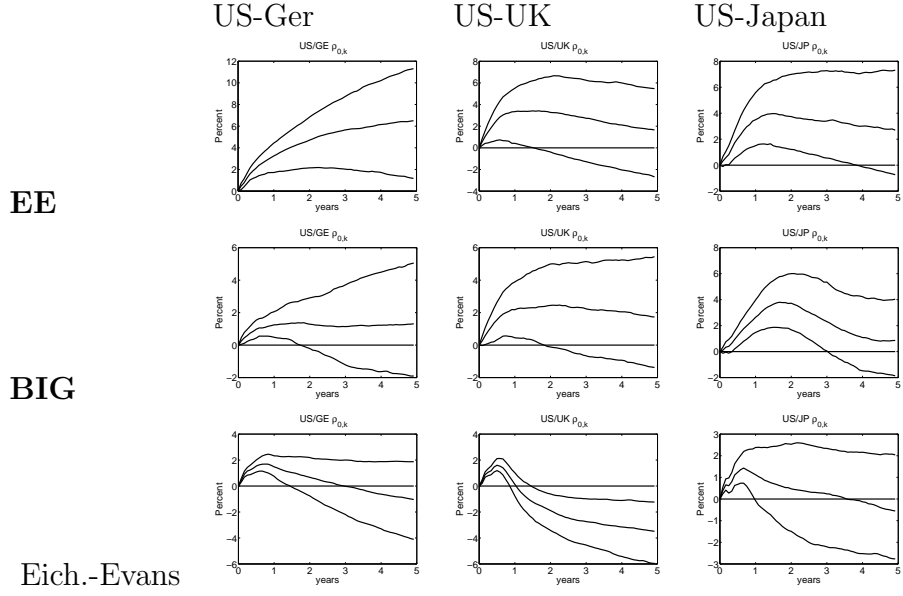


Figure 11: Impulse responses for the forward discount premium  $\rho_k$ , conditional on a US monetary policy contraction.

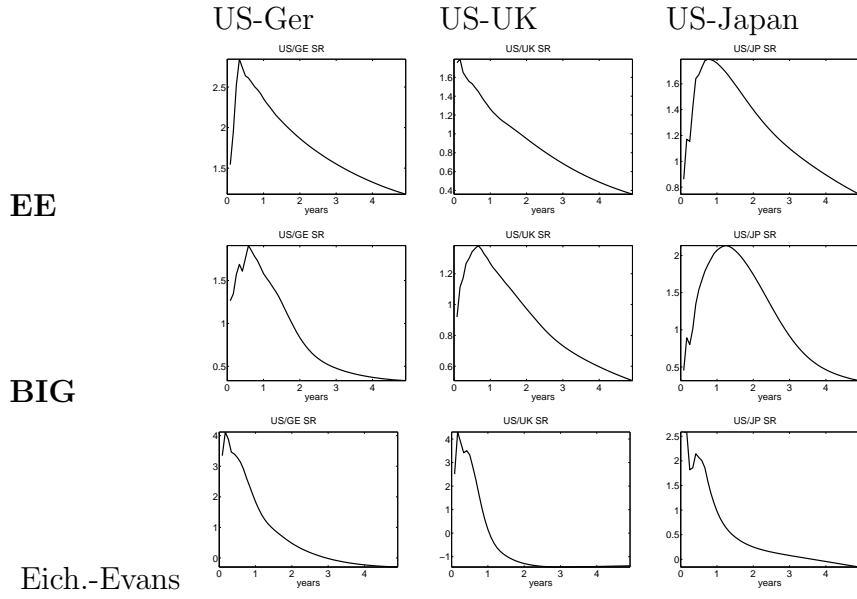


Figure 12: Impulse responses for the Sharpe ratio of a Bayesian investor, conditional on a US monetary policy contraction.

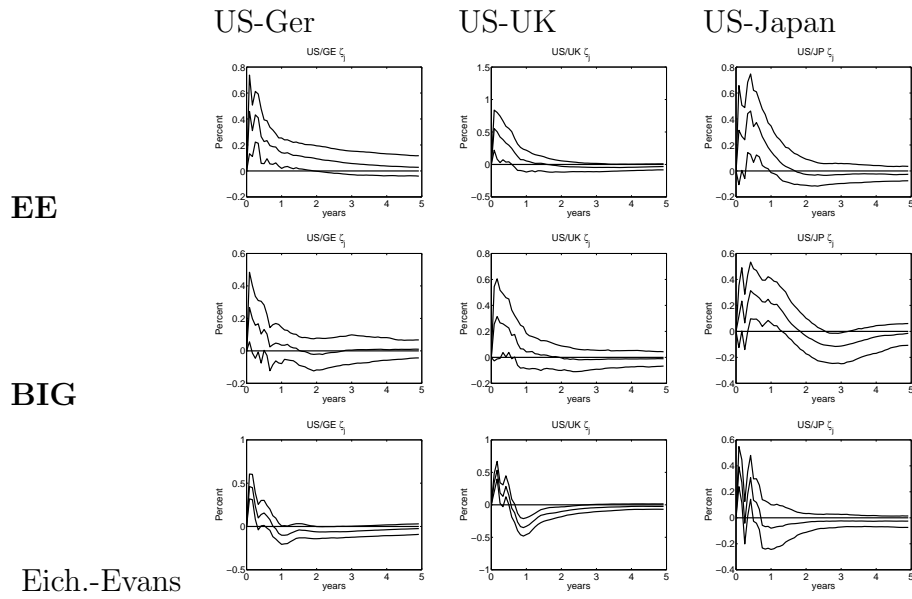


Figure 13: Impulse responses for the forward discount premium  $\xi_k$  for a one-month arbitrage strategy from periods  $k-1$  to  $k$  after a contractionary US monetary policy shock.

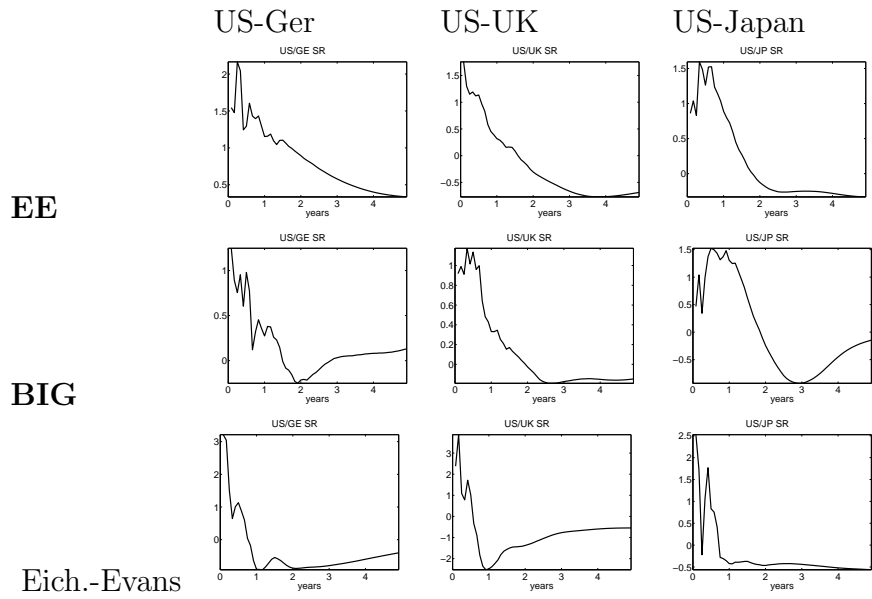


Figure 14: Impulse responses for the Sharpe ratio for one-period arbitrage trades  $\xi_k$  between  $k-1$  and  $k$  of a Bayesian investor, conditional on a US monetary policy contraction.

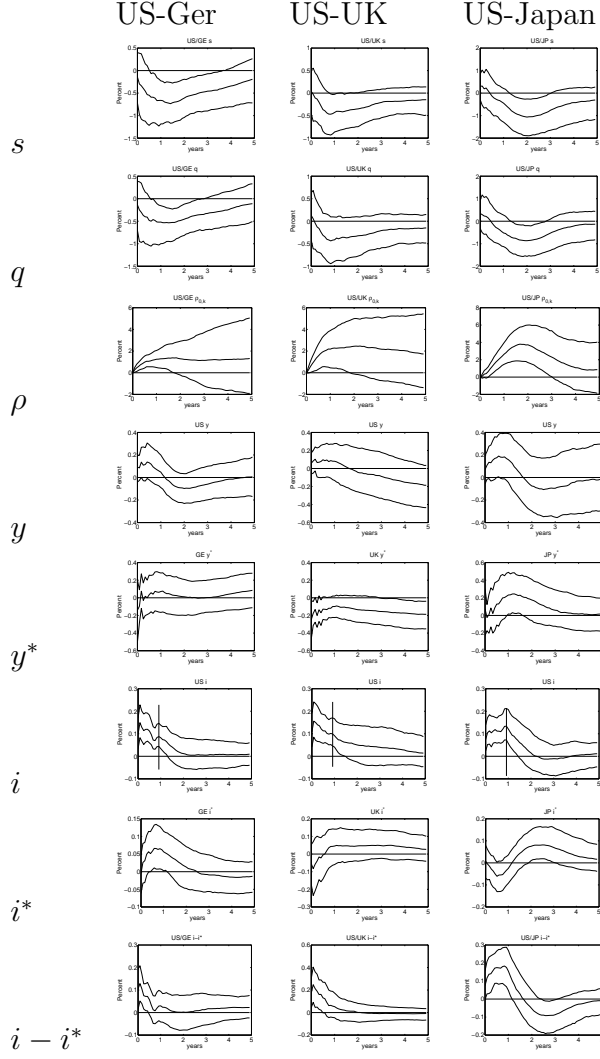


Figure 15: This figure shows impulse responses to a US monetary policy contraction for the **BIG** VAR, using  $K = 5$ .



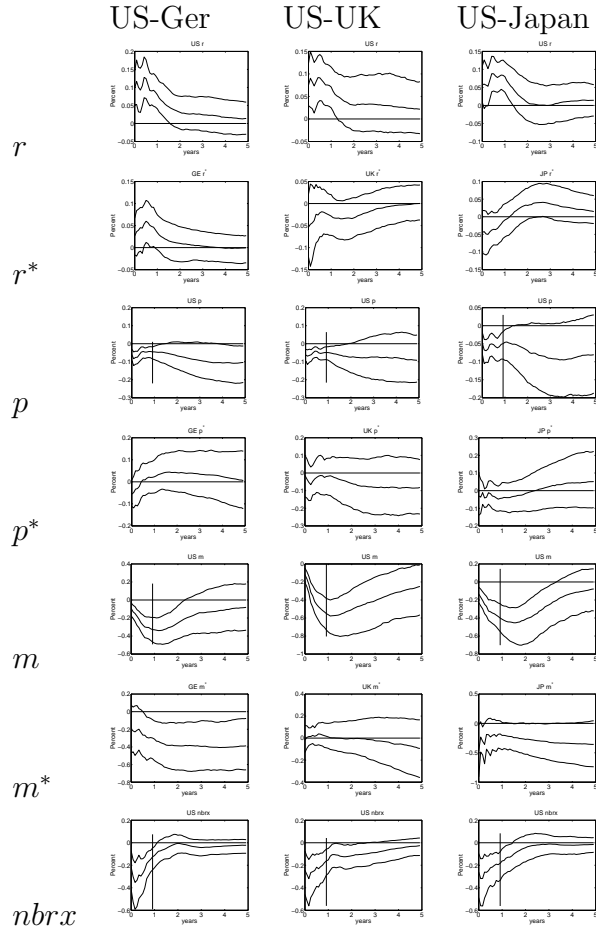


Figure 15, impulse responses to US monetary policy contraction for the **BIG** VAR, continued.

### 6.4.3 Foreign monetary policy shocks

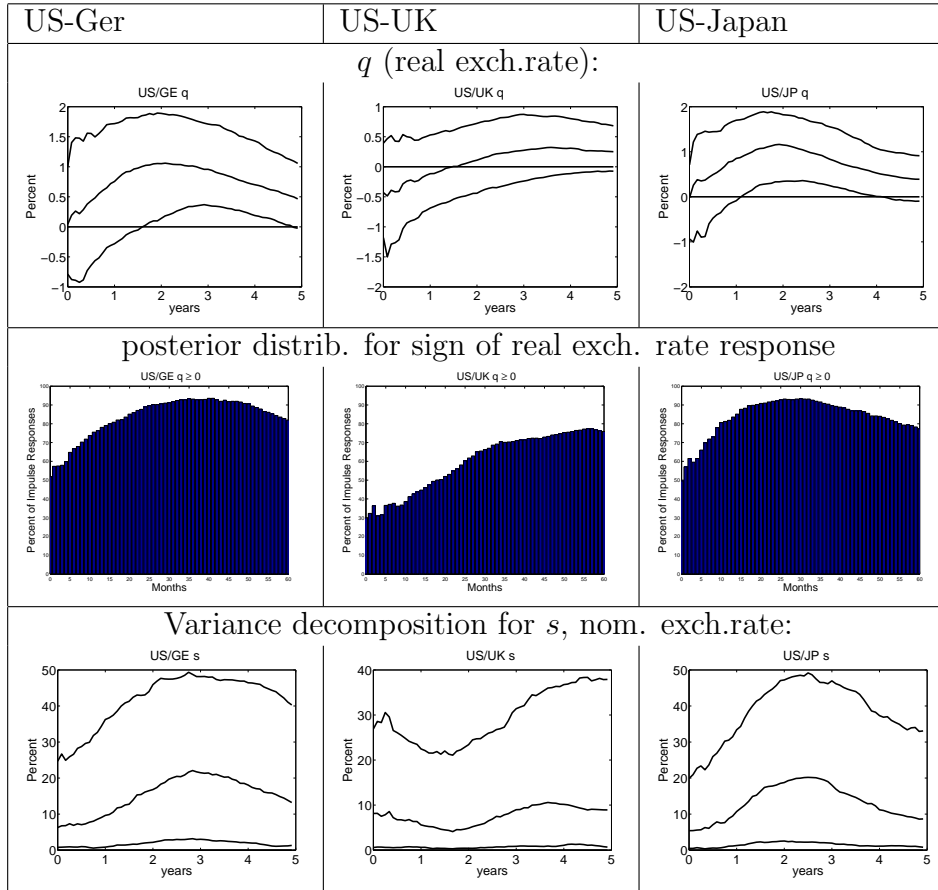


Figure 16: Key results for a **foreign monetary policy contraction**, using the benchmark identification with  $i^* - i \geq 0$  in the **GR** specification. Shown are the results for the response of the real exchange rate  $q$ , the posterior distribution for its sign distribution and the variance decomposition for the nominal exchange rate.

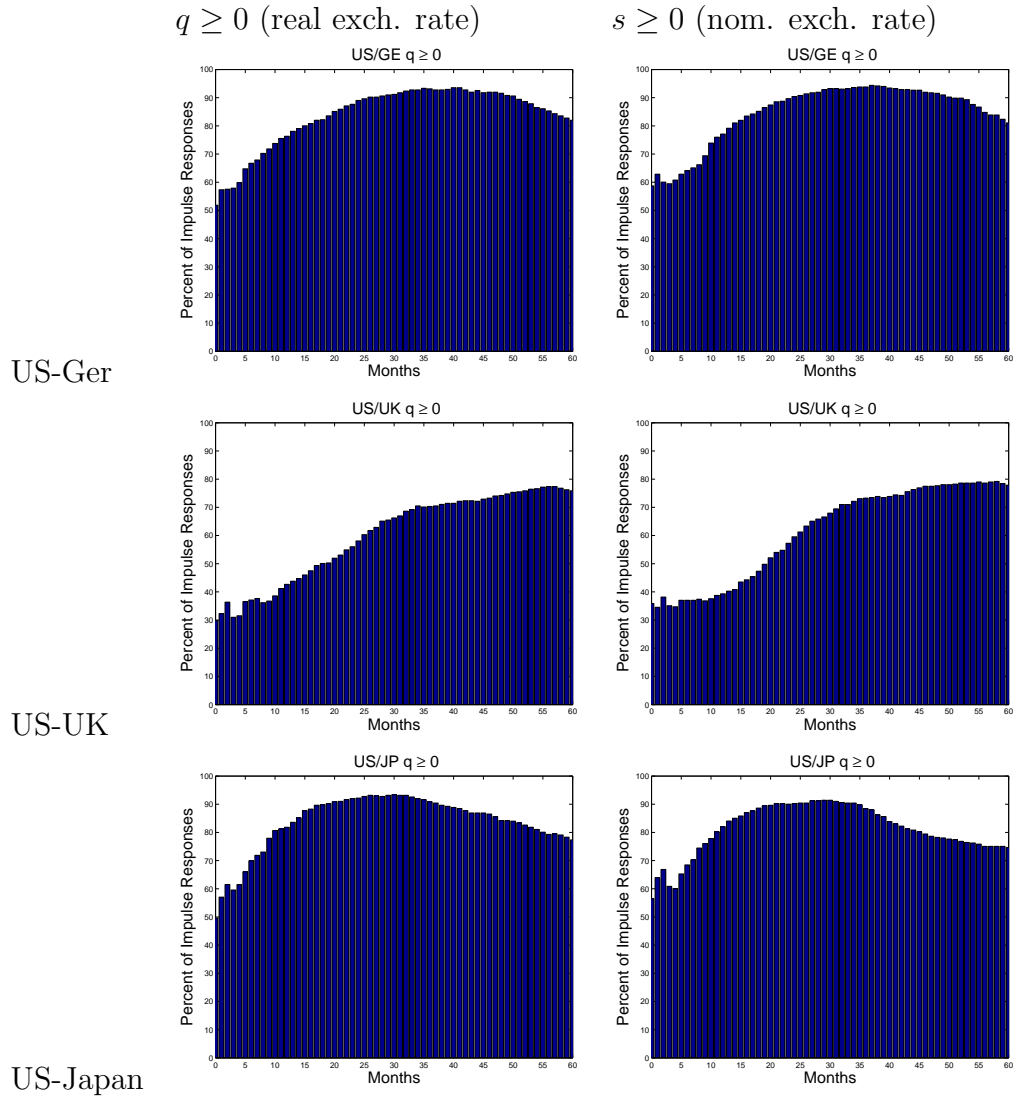


Figure 17: Posterior distribution for a positive real or nominal exchange rate response to a foreign monetary contraction for the benchmark identification, **GR** VAR.

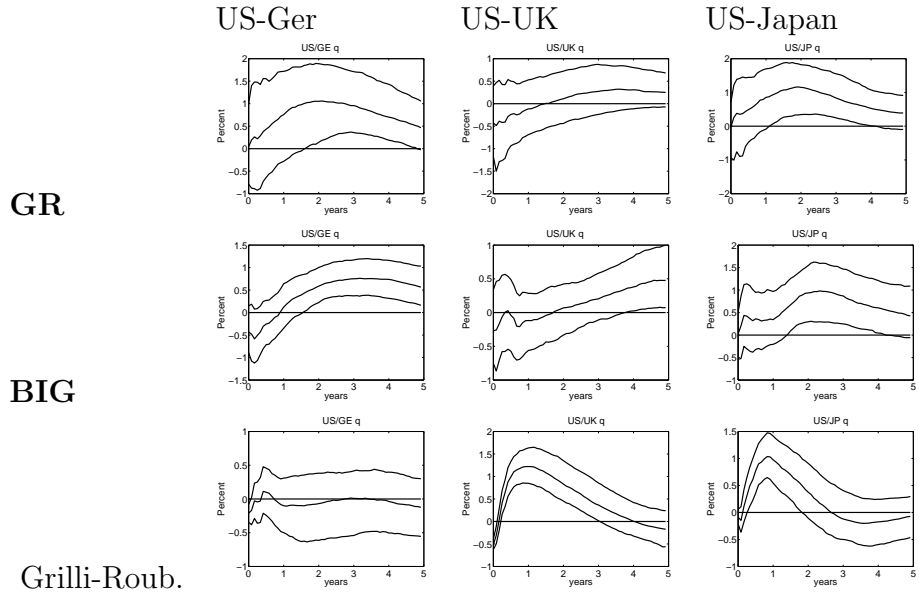


Figure 18: Impulse response function of the real exchange rate, conditional on a foreign monetary policy contraction. We compare the benchmark identification and imposing  $i^* - i \geq 0$  in the **GR** VAR specification to the **BIG** VAR specification as well as the recursive Grilli-Roubini identification.

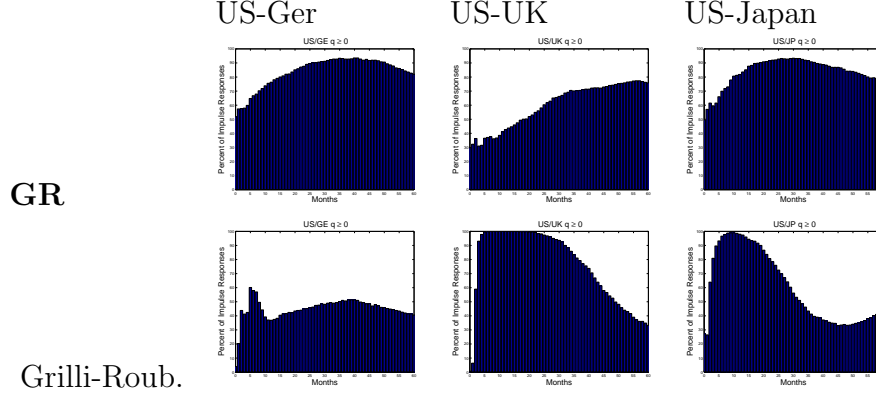


Figure 19: Distribution for the sign  $q \geq 0$  of the real exchange rate, conditional on a foreign monetary policy contraction. We compare the benchmark identification and imposing  $i^* - i \geq 0$  in the **GR** VAR specification to the **BIG** VAR specification as well as the recursive Grilli-Roubini identification.

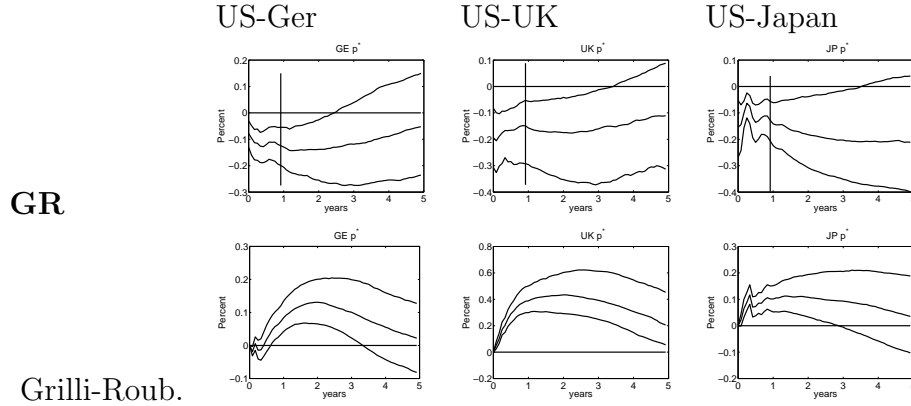


Figure 20: Impulse response function of the price level, conditional on a foreign monetary policy contraction. We compare the benchmark identification and imposing  $i^* - i \geq 0$  in the **GR** VAR specification to the **BIG** VAR specification as well as the recursive Grilli-Roubini identification. Note that there is a considerable price puzzle in the Grilli-Roubini specification, which is avoided with sign restrictions by construction.

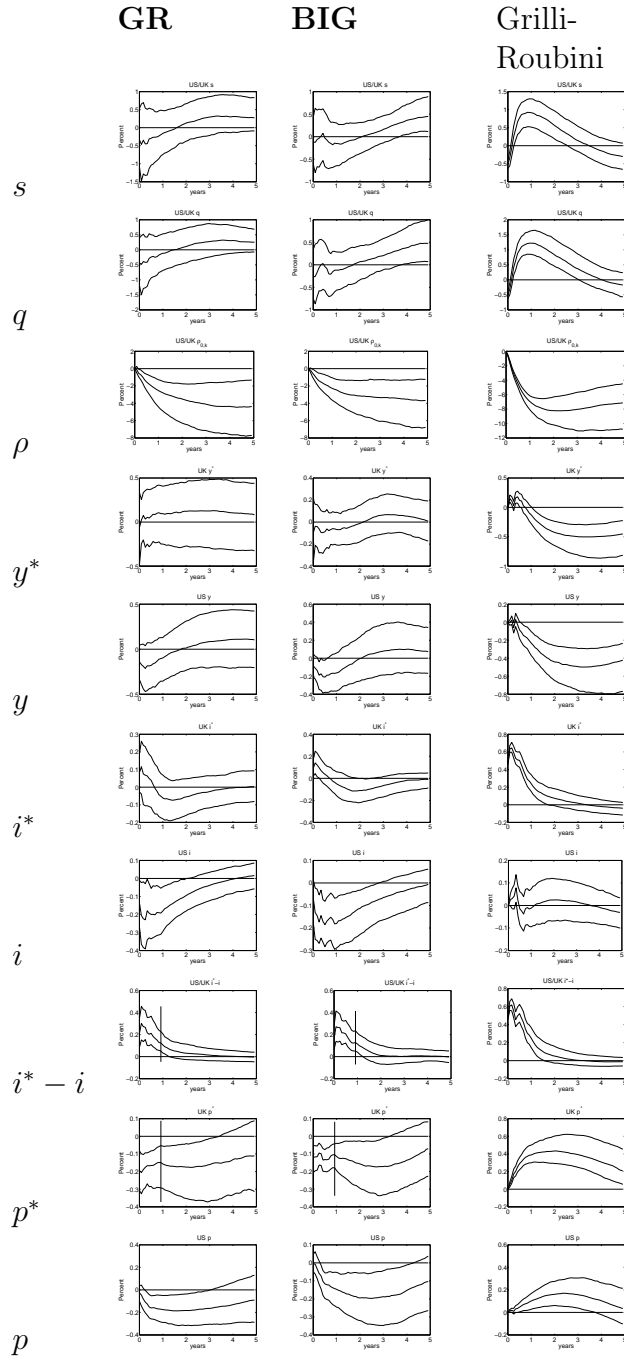


Figure 21: Comparison of results. For a foreign monetary policy contraction and the US-UK country pair, we compare the benchmark identification in the **GR** specification to the **BIG** specification as well as the original identification of Grilli-Roubini.

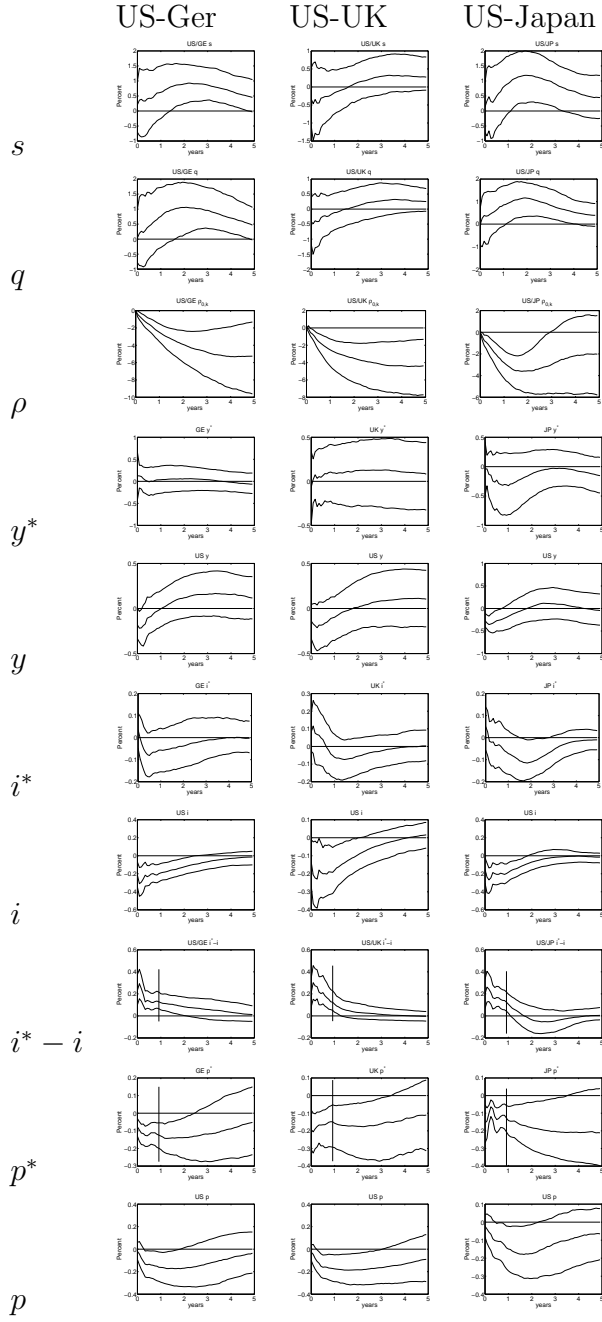


Figure 22: Impulse responses to a foreign monetary policy contraction in the GR VAR, using the restriction  $i^* - i \geq 0$ .



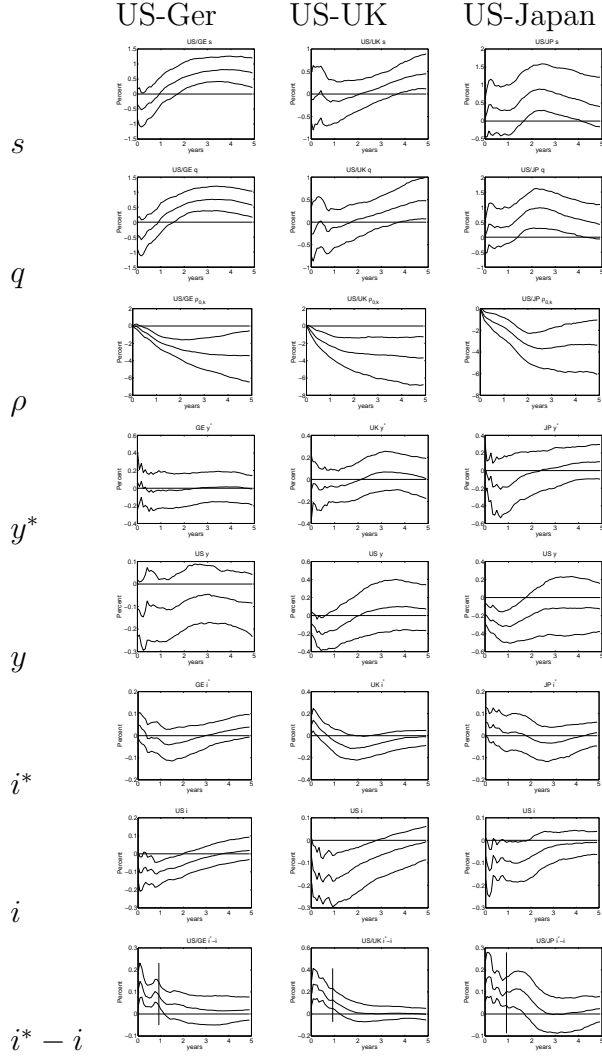


Figure 23: Impulse responses to a foreign monetary contraction. Identification and  $i^* - i \geq 0$  for the **BIG** VAR.

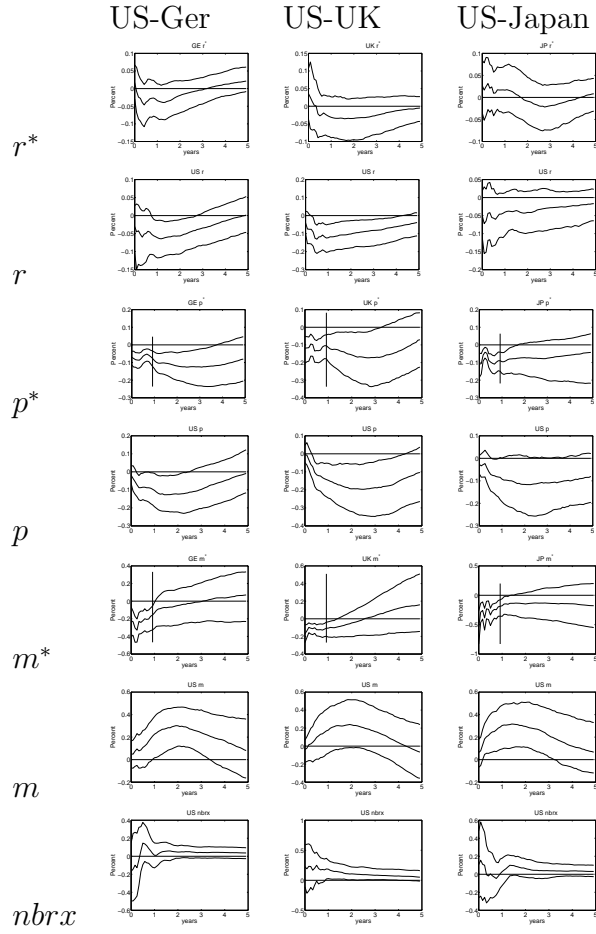


Figure 23, impulse responses to a foreign monetary contraction for the **BIG** VAR, continued.

#### 6.4.4 Variance decompositions

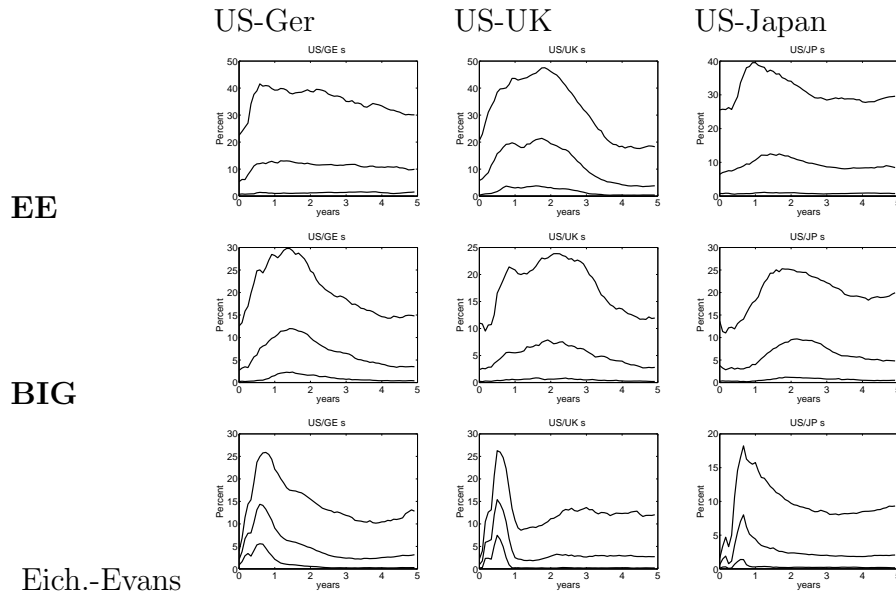


Figure 24: Variance decompositions: the contribution of US monetary policy shocks to the variance of the nominal exchange rate.

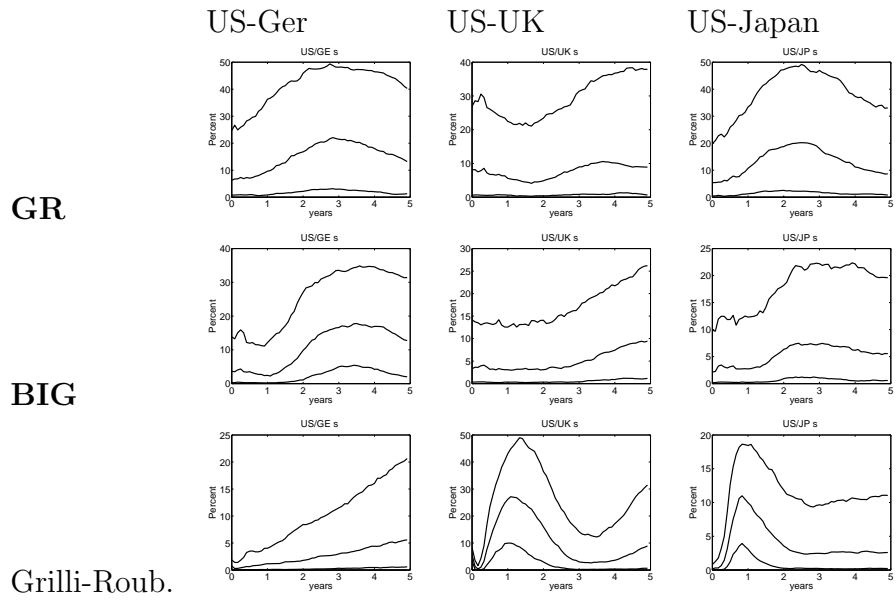


Figure 25: Variance decompositions: the contribution of foreign monetary policy shocks to the variance of the nominal exchange rate.