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

Almuth Scholl
Humboldt Universität zu Berlin
and Harald Uhlig
Humboldt Universität zu Berlin
Deutsche Bundesbank, CentER and CEPR

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*Address: Humboldt Universität zu Berlin, Wirtschaftswissenschaftliche Fakultät, Spandauer Str. 1, 10178 Berlin, GERMANY. e-mail: ascholl@wiwi.hu-berlin.de, uhlig@wiwi.hu-berlin.de, fax: +49-30-2093 5934, home page http://www.wiwi.hu-berlin.de/wpol/. This research was supported by the Deutsche Forschungsgemeinschaft through the SFB 649 "Economic Risk" and by the RTN network MAPMU. We are grateful to comments received at MIT and at the "Macroeconomics and Reality: 25 years later" conference in Barcelona, April 2005.

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 (2005) 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 (1995) and by Grilli-Roubini (1995, 1996), which are particularly pronounced, when using an updated data set. We find that the puzzles regarding the exchange rates are still there, but that the quantitative features are different. In response to US monetary policy shocks, 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 quicker than the three-year horizon found by Eichenbaum-Evans. There is a robust forward discount puzzle implying a large risk premium. We study this issue, introducing and calculating conditional Sharpe ratios for a Bayesian investor investing in a hedged position following a US monetary policy shock. For foreign monetary policy shocks, we find more robust results than with the Grilli-Roubini recursive identification strategy: the posterior distribution regarding the exchange reaction looks rather similar across countries and VAR specifications. In particular, we find that there seems to be considerable uncertainty regarding the initial reaction of the exchange rate. 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: C32, E58, F31, F42

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 (1972, 1980), employing vector autoregressions to study these issues in order to sort out the issues of causality of monetary policy shocks. 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) employ recursive identification and find a persistent appreciation for periods up to three years, in contrast to theory. Also Leeper, Sims and Zha (1996) find this result for their larger specifications, 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, see Clarida and Gali (1994), Eichenbaum and Evans (1995), Rogers (1999), Faust and Rogers (2000) and Kim and Roubini (2000).

Recently, this conventional view has come under attack. Most notably, Faust and Rogers (2003) argue, that one needs to "relax dubious identifying assumptions stemming from e.g. recursive identification 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 approach of Uhlig (2005) of imposing sign restrictions on impulse response functions. 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. 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. For example, to implement the benchmark view that inflation slows down after a monetary tightening, an on-impact drop in prices followed by a much larger subsequent rise ought to be ruled out as a possible response to a contractionary monetary policy shock. This is exactly what we do here. Thereby, we can narrow down the range of possible monetary policy shocks considerably, as has already been argued in Uhligs (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 identifiying 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 rates 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 Uhligs (1998) discussion of Faust (1998).

The identification method used in this paper, introduced by Uhlig (2005), builds on the classic paper by Leamer (1981) and its macroeconomic implementation by Blanchard (1989) and is related to work by Canova and Pina (1999) and Canova and de Nicolo (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. We complement this literature by introducing and empirically calculating conditional Sharpe ratios for a Bayesian investor investing in a hedged position following a US monetary policy shock.

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 and 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 i_k and i_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,

$$i_{0\to k} = \sum_{j=0}^{k-1} i_j,\ i^*_{0\to k} = \sum_{j=0}^{k-1} i^*_j$$

Define the forward discount premium

$$\rho_k \quad = \quad s_0 - s_k + i_{0\to k} - i^*_{0\to k}$$

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

$$\xi_j = s_{j-1} - s_j + i_{j-1} - i^*_{j-1}$$

is the same gain when executing this hedging strategy only from periods $j-1$ to j following the shock. In particular, ρ_j stays flat for $j \geq j^*$, if $\xi_j = 0$ for $j \geq j^*$. Note furthermore, that the reaction s_0 in the impact period is not part of the foreign discount premium, i.e., we assume that the investor starts the investment strategy after observing the monetary policy shock and the concurrent on-impact movement of the exchange rate. The investment strategy examined here relies on predictable movements, and not on reacting more quickly than the foreign exchange market to news about monetary policy. Put differently, as long as an investor starts this investment strategy within the "impact month" of the monetary policy shock, he will receive ρ_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 and 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

$$q_k = s_k + (p_k^* - p_k) \tag{1}$$

be the impulse response of the real exchange rate. The Dornbusch (1976) overshooting model results from adding to (conditional) UIP the assumptions, that the impulse response for the long run real exchange rate q_k will converge to zero, $q_k \to 0$ due to long run purchasing power parity, that prices are sticky and that the differences in the short rates $i_j - i_j^*$ slowly reverts to zero following the initial monetary policy shock characterized by a liquidity effect $i_0 - i_0^* > 0$. I.e., under these assumptions, q_0 should be large and negative and slowly revert back to zero and 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 implicitely) 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

$$X_{t+k} = (1 - e^{\rho_k})e^{(i_{0\to k}+p_0-p_k)} \tag{2}$$

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 $i_{0\to k}$. Let the random return of that asset between period t and $t+k$ be given by $\tilde{R}_{t+k}$, and let $\tilde{\rho}_{t+k} = i_{0\to k} - \log \tilde{R}_{t+k}$

be the log excess return. As above, the payoff to this strategy is

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

and the asset pricing equation reads

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

Let

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

be the logarithm of the expected excess return. Let

$$r_{0\to k} = i_{0\to k} + p_0 - p_k$$

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

$$\mathrm{SR}_{\tilde{\rho},k} \equiv \frac{\mu_{\tilde{\rho}}}{\sigma_{\tilde{\rho}}} = -\mathrm{corr}_t(m_{t,t+k}, \tilde{\rho}_{t+k})\sigma_m - \mathrm{corr}_t(r_{0\to k}, \tilde{\rho}_{t+k})\sigma_{r_{0\to k}} \tag{4}$$

where $\mathrm{corr}_t(\cdot,\cdot)$ denotes conditional correlation, where σ_m, $\sigma_{\tilde{\rho}}$ and $\sigma_{r_{0\to k}}$ are the conditional standard deviations of $m_{t,t+k}$, $\tilde{\rho}_{t+k}$ and $r_{0\to k}$. This equation defines the Sharpe ratio $\mathrm{SR}_{\tilde{\rho},k}$, 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 $\mathrm{SR}_{\rho,k}$ for the hedging strategy ρ_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,

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

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

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

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

$$a = \tilde{A}\alpha, \tag{5}$$

where $\tilde{A}\tilde{A}' = \Sigma$ is some decomposition of Σ and α 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 Σ. Then, the impulse response $r_a(k)$ for a is given by

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

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 $(r_a(k))_j \leq 0, j \in \mathcal{J}_-$ for some subsets of variables $\mathcal{J}_+$ and $\mathcal{J}_-$ and some horizon $k = 0, \ldots, K$.

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

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[1]. The prior and therefore the posterior belong to the Normal-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 specifica-

[1] In fact, much of the evidence becomes considerably weaker, when using 12 lags instead, which we did in a previous version of this paper.

tions 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 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.

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

US monetary policy shock:

Specification	**EE**	**BIG**
Variables	y, y^*, p, $nbrx$, i, i^*, s	y, y^*, p, p^*, m, m^*, i, i^*, s, r, r^*, $nbrx$
Restrictions:	$p \leq 0$, $nbrx \leq 0$, $i \geq 0$	$p \leq 0$, $m \leq 0$, $nbrx \leq 0$, $i \geq 0$

foreign monetary policy shock:

Specification:	**GR**	**BIG**
Variables:	y, y^*, p, p^*, i, i^*, s	y, y^*, p, p^*, m, m^*, i, i^*, s, r, r^*, $nbrx$
Restrictions:	$p^* \leq 0$, $i^* - i \geq 0$	$p^* \leq 0$, $m^* \leq 0$ $i^* - i \geq 0$

Table 1: Identification of Monetary Policy shocks.

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. Consider a surprise rise in the interest rate of some country A of interest, which may be contemporaneous with a surprise rise of interest rates in some other country B. With the first interpretation, one can think of the central bank of country A as a Stackelberg leader: if both interest rates go up, then it and not necessarily the central bank of country B was the ultimate cause. In the second, the central bank of country A is viewed as a follower: deviations from the interest rate set by the other central bank of country B are viewed as the key surprise, to which the central bank in country A reacts quickly.

In principle, the game played between two monetary authorities could be rather complicated. There is really no good reason to a priori rule out e.g. a game in which central banks alternate as to who is using the higher interest rate. And it is easy to think of considerably more complicated games. 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 $k = 0, \ldots, K = 11$, i.e. one 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 13 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 e.g. the peak of nominal exchange rate appreciation is very similar to the posterior for the peak of the real exchange rate appreciation: we therefore concentrate on the real exchange rate only.

The results show that the US-German and US-UK bilateral real exchange rates appreciate until approximately the forth to twelfth month and then depreciates in our benchmark identification. The delay is slightly shorter for Germany than for the UK, and considerably longer in the US-Japan case, with most of the distribution centered around months 8 to 24. 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 6. The joint distribution is shown both for the first five years as well as for the first 24 months. Also, the distribution of the beyond-impact change of the real exchange rate $q(k)-q(0)$ is shown. 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 7 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. 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 8 (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. This price puzzle is considerably more pronounced than in the original source, because the recursive identification procedures have been applied here to our updated data set, using data from 1975 to 2002. 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 13 (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 9 and 10. 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 finds 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 the recursive identification as plausible for this updated data set, however, as we have argued before[2]. 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 ξ_k for one-month holding periods in figure 11 and the corresponding Sharpe ratios in figure 12. 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 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 .

[2]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.

4.2 Foreign Monetary Policy Shocks

For a foreign monetary policy shock, figure 14 contains the key results, while figures 15 to 20 contain further details.

The first line of figure 14 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. The results for nominal and real exchange rates are very similar throughout: we thus only show the latter.

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 15 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.

Figures 16 and 17 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. For our specification and on impact, the posterior probability for a positive exchange rate response is somewhere between 30% and 60%. This probability then rises and plateaus somewhere near 70% to 90% within three years. 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 shows the advantage of using the sign restriction methodology as opposed to their original identification strategy, when applying it to this updated data set.

Figure 18 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 15. 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 19 as well as 20 (two pages) for completeness and to emphasize, that our identification strategy is reasonable.

The last line of figure 14 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 21 contains the variance decomposition for the exchange rate movements explained by a US monetary policy shock, while figure 22 contains the fraction of exchange rate variation explained by foreign monetary policy shocks.

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 22 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-Roubinin specifications as well as the sharp peaks in the Eichenbaum-Evans specifications may be viewed as a further odity, when applying these conventional identification strategies, which are avoided with the sign restriction approach.

For the sign restriction approach, the 86% quantile rarely moves beyond 30% for any country or specification. When considering the joint contribution of both monetary policy shocks the median estimate ranges between 10 and 20 percent, and the 86% quantile reaches up to 45 percent of the exchange rate movements.

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 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 inconclusive 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) identifiying restrictions compared to 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
ρ	forward discount premium (accumulated)	derived
ξ	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^*, y

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: Eichenbaum and Evans assume the recursive ordering $[y, p, y^*, i^*, nbrx, i, s]$. A US monetary policy contraction is identified with a fall in $nbrx$.

Grilli-Roubini: Grilli and Roubini assume a recursive ordering $[y^*, p^*, y, p, i, i^*, s]$. A foreign monetary policy contraction is identified with a rise in the foreign short term interest rate.

Faust-Rogers: Faust and Rogers consider the EE specification and analyze the impact of a monetary innovation. Two sets of restrictions are assumed. The first set consists of money restrictions: on impact p, y, y^*, $nbrx$, s are larger than or equal to zero and i and i^* are less than or equal to zero; p at horizon 80 is not larger than at horizon 36; y^* is no more than one-half of that of y on impact and the decline in i^* is not larger than one-half of the decline in i on impact. The second set consists of shape restrictions on the exchange rate: the exchange rate responses falls between lags 1-2, 2-3, 3-4, 4-6, 6-12, 12-18, 18-36, 18-80.

Faust and Rogers report results under three combinations of these restrictions: money restrictions only, money and shape restrictions, and neither money or shape restrictions.

6.4 Figures

6.4.1 Theory

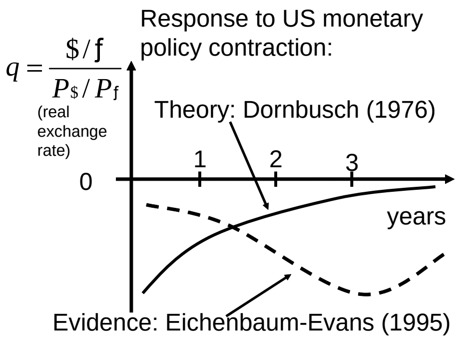

Figure 1: A stylized representation of the forward discount puzzle

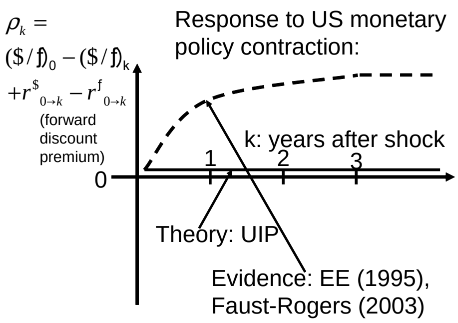

Figure 2: A stylized representation of the forward discount puzzle

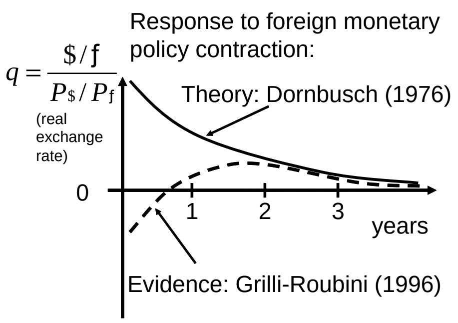

Figure 3: A stylized representation of the exchange rate puzzle

6.4.2 Responses to a US Monetary Policy Contraction.

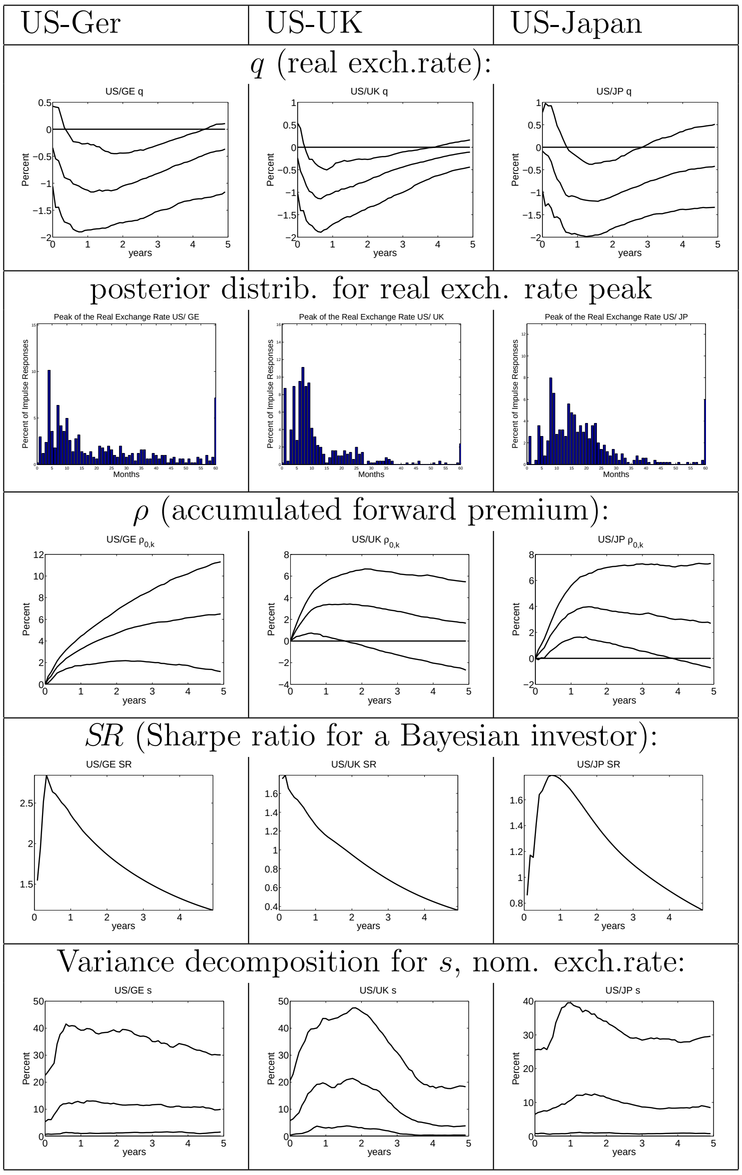

Figure 4: Key results for a **US monetary policy shock**, using the benchmark identification in the **EE** specification. Shown are the results for the response of the real exchange rate q, the posterior distribution for its peak, the response of the (accumulated) forward premium, the Sharpe ratio of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

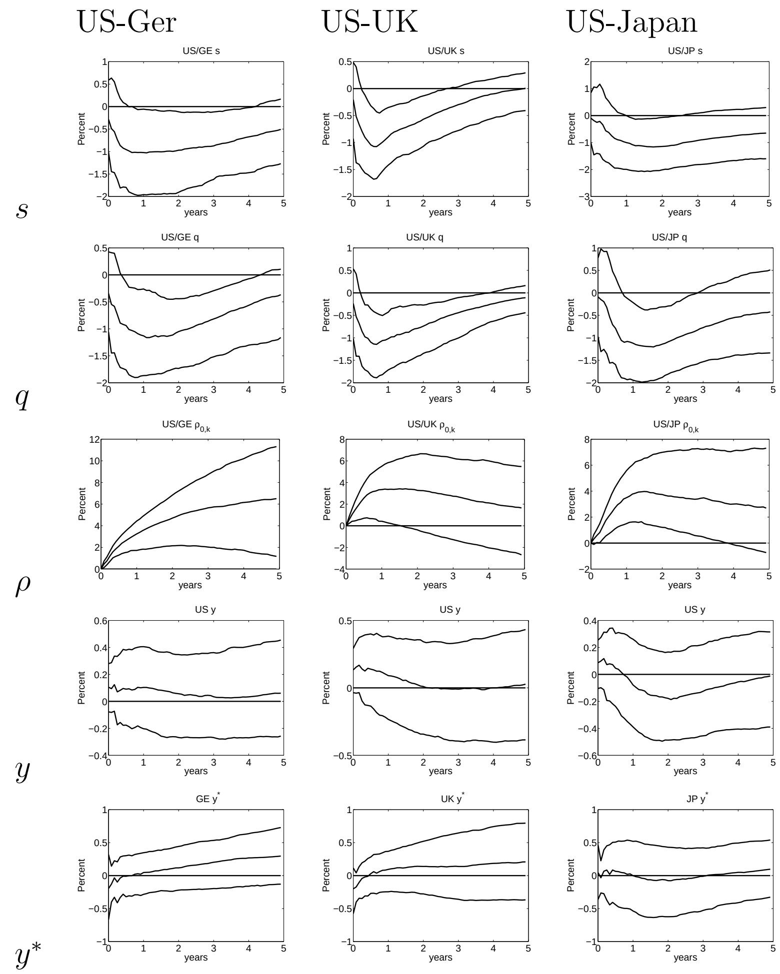

Figure 5: Part 1 of the impulse responses for a US monetary policy contraction, using the benchmark identification for the **EE** specification. Note that there is little difference between the response of the nominal and the real exchange rates.

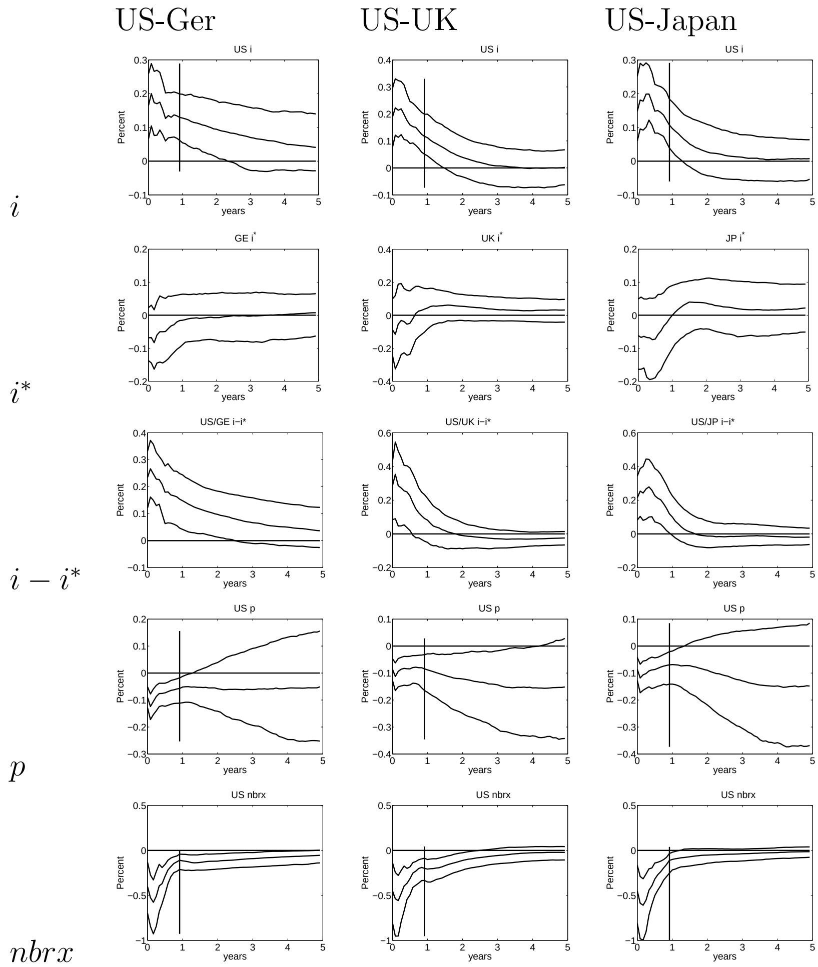

Figure 5 continued. Part 2 of the impulse responses for a US monetary policy contraction, using the benchmark identification in the **EE** specification.

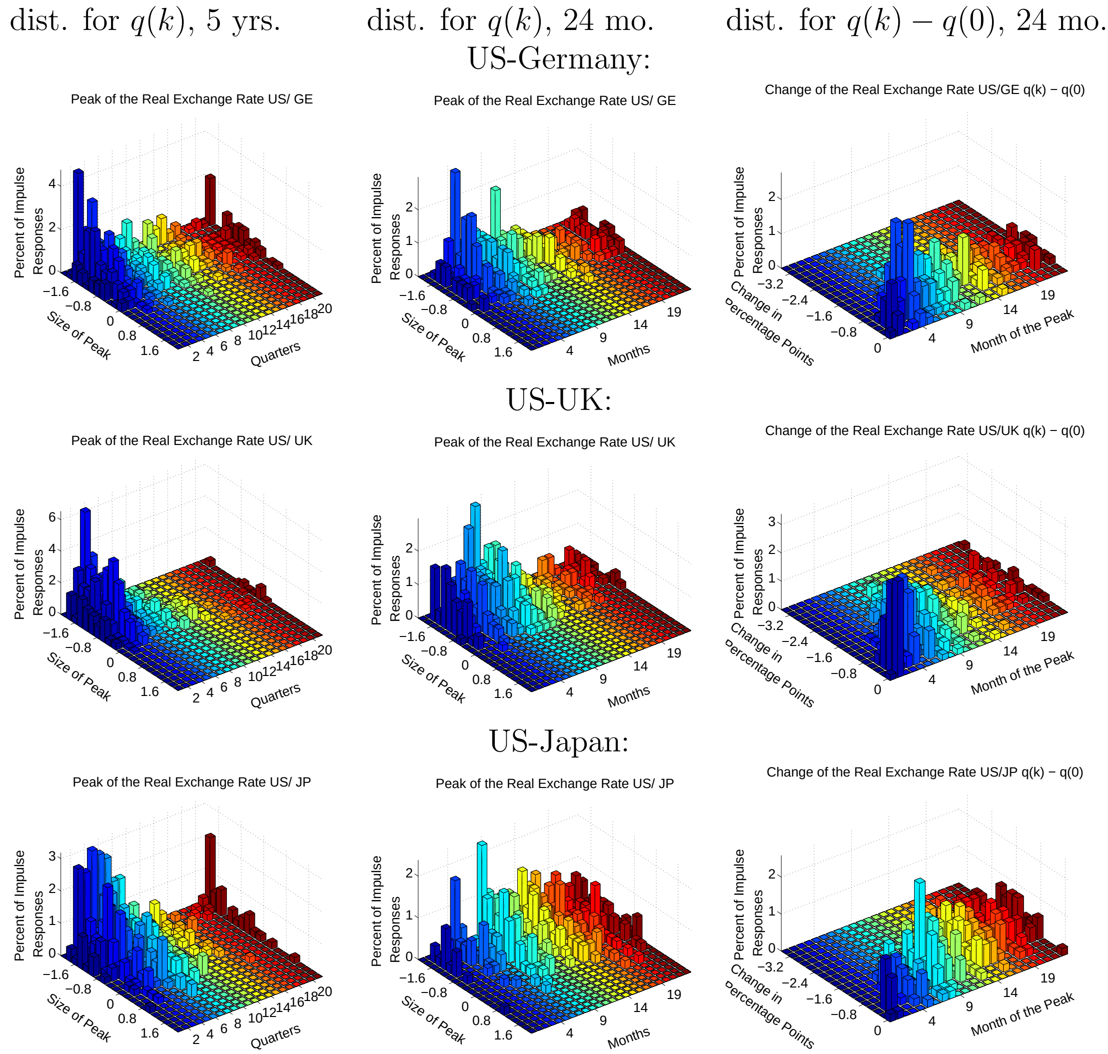

Figure 6: **Real exchange rate** q**.** Posterior distribution for the size and location of the peak appreciation $q(k)$ as well as for the change relative to the impact reaction $q(k) - q(0)$, conditional on a US monetary policy contraction, for the benchmark identification, **EE** VAR. The left column shows the distribution over the first 20 quarters, whereas the other two columns use a monthly scale for a closer examination of the first 24 months.

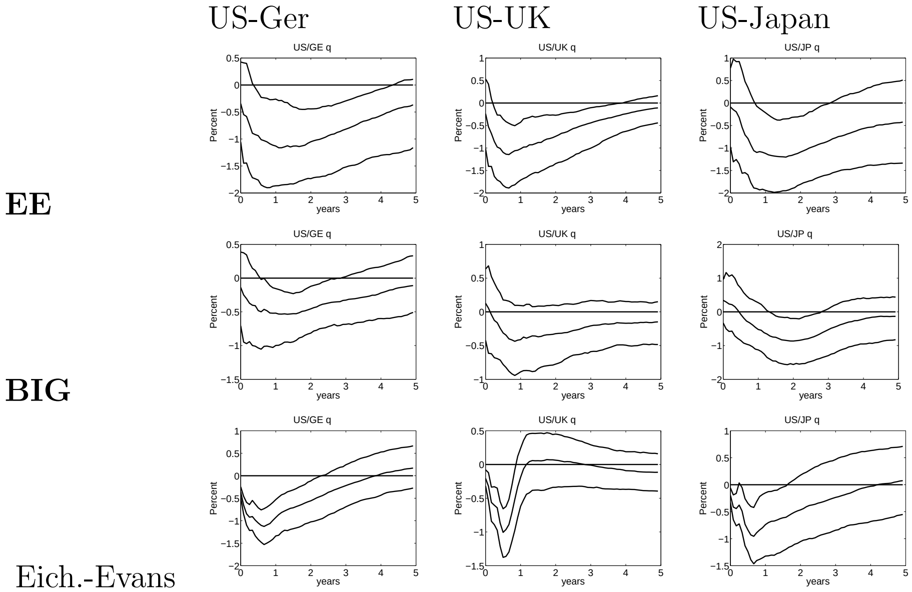

Figure 7: Impulse response function for the real exchange rate, conditional on a US monetary policy contraction. We compare the benchmark specification identification in the **EE** VAR specification to the identification in the **BIG** VAR to the original Eichenbaum-Evans recursive identification.

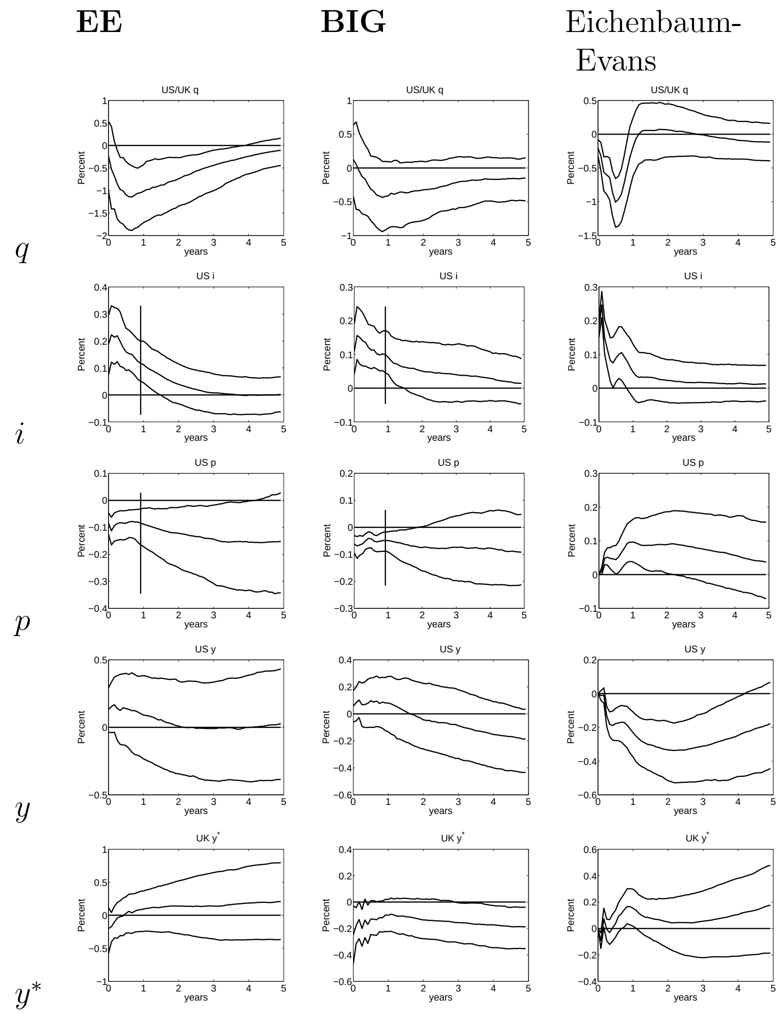

Figure 8: Comparison of results. For a US monetary policy contraction and the US-UK country pair, we compare the benchmark identification in **EE** to the **BIG** specification and the original identification of Eichenbaum-Evans. Note the "price puzzle" in the Eichenbaum-Evans column.

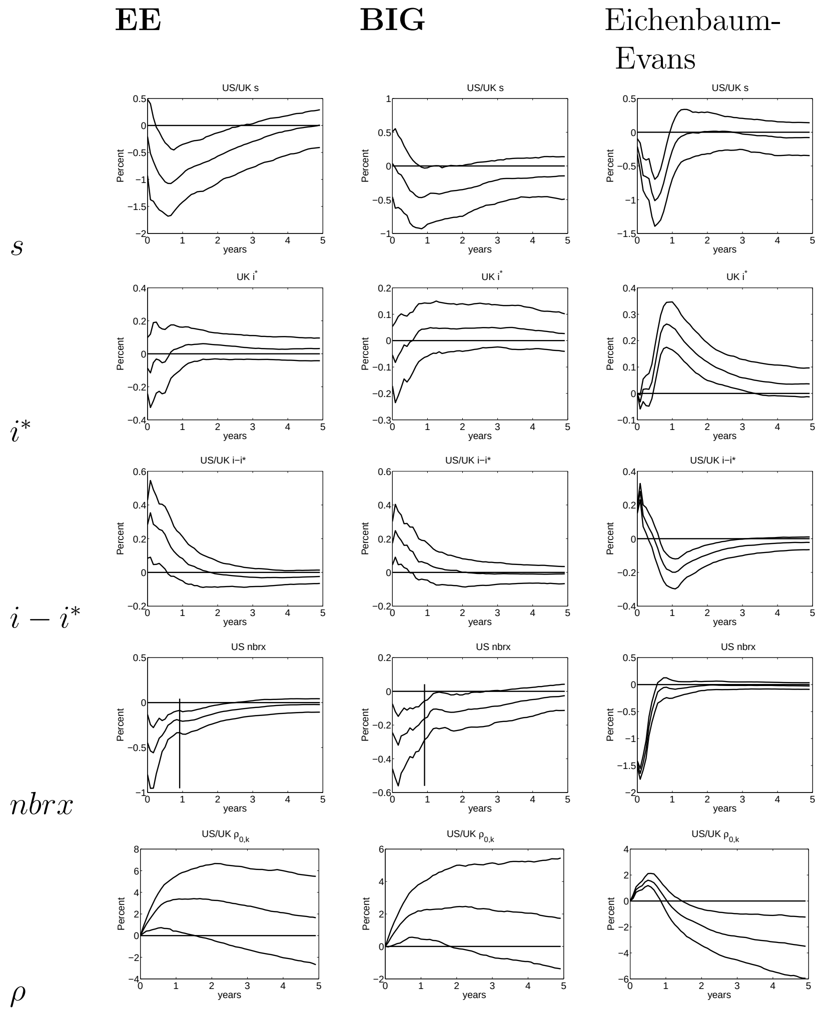

Figure 8 continued. Part 2 of the comparison.

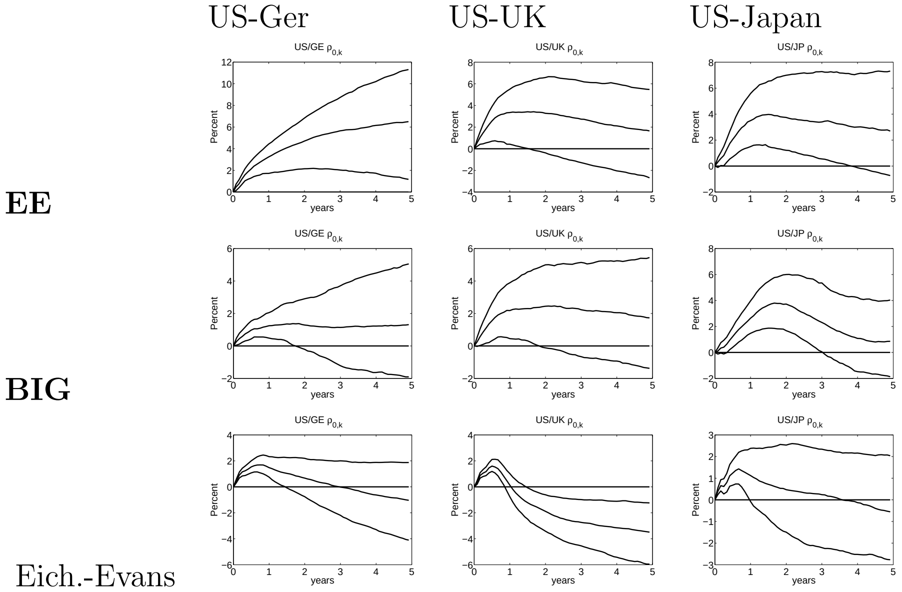

Figure 9: Impulse responses for the forward discount premium ρ_k, conditional on a US monetary policy contraction.

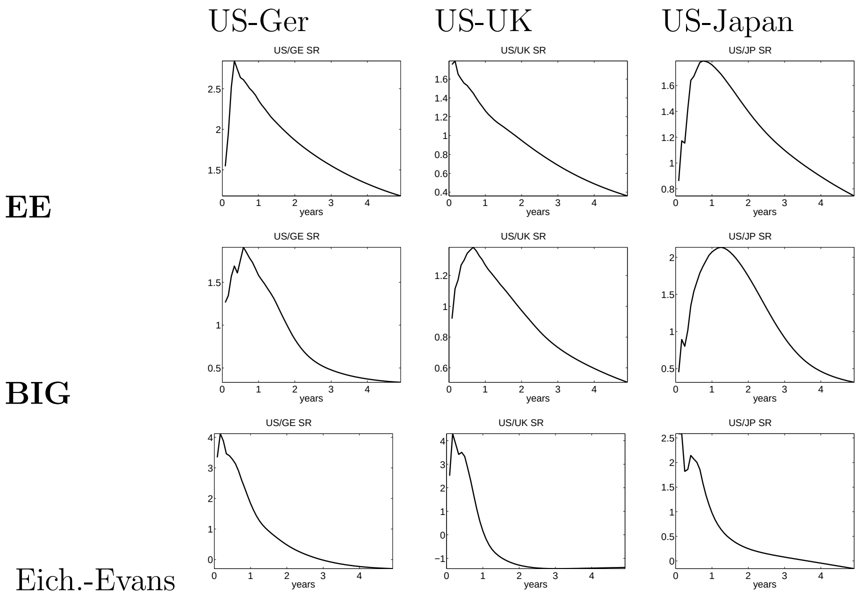

Figure 10: Impulse responses for the Sharpe ratio of a Bayesian investor, conditional on a US monetary policy contraction.

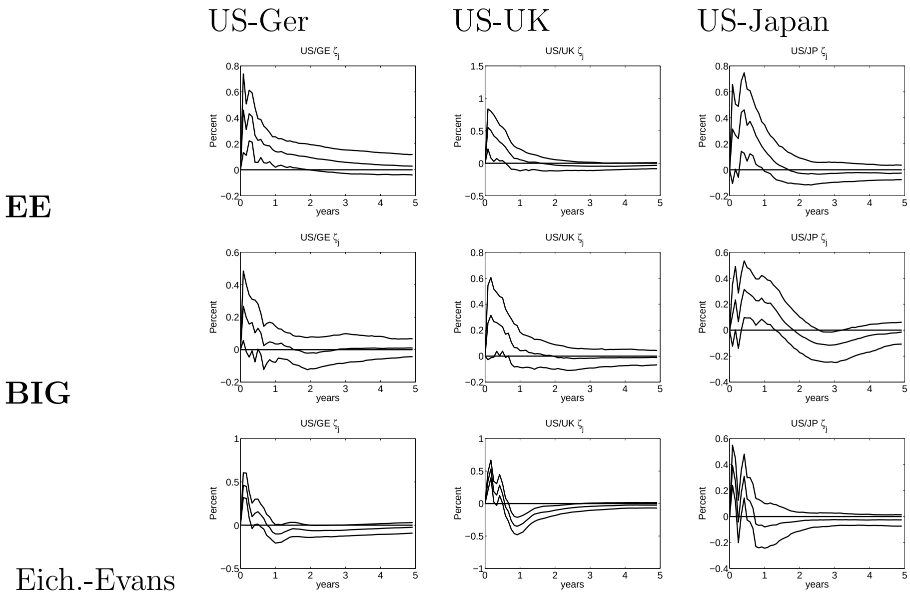

Figure 11: Impulse responses for the forward discount premium ξ_k for a one-month arbitrage strategy from periods $k-1$ to k after a contractionary US monetary policy shock.

US-Ger US-UK US-Japan

EE

BIG

Eich.-Evans

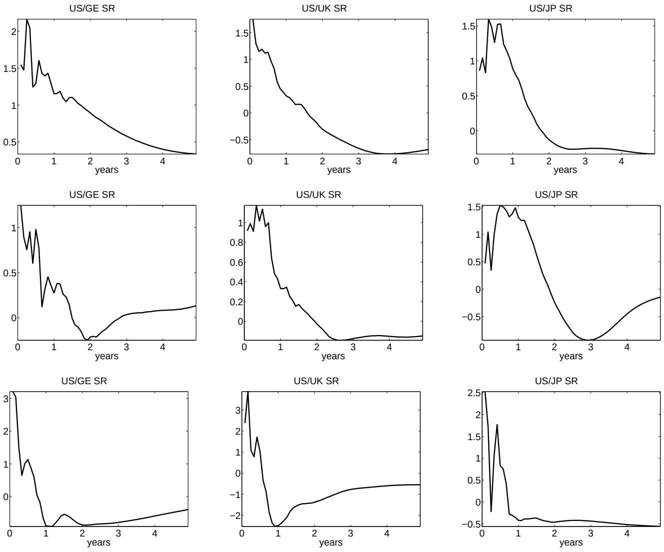

Figure 12: Impulse responses for the Sharpe ratio for one-period arbitrage trades ξ_k between $k-1$ and k of a Bayesian investor, conditional on a US monetary policy contraction.

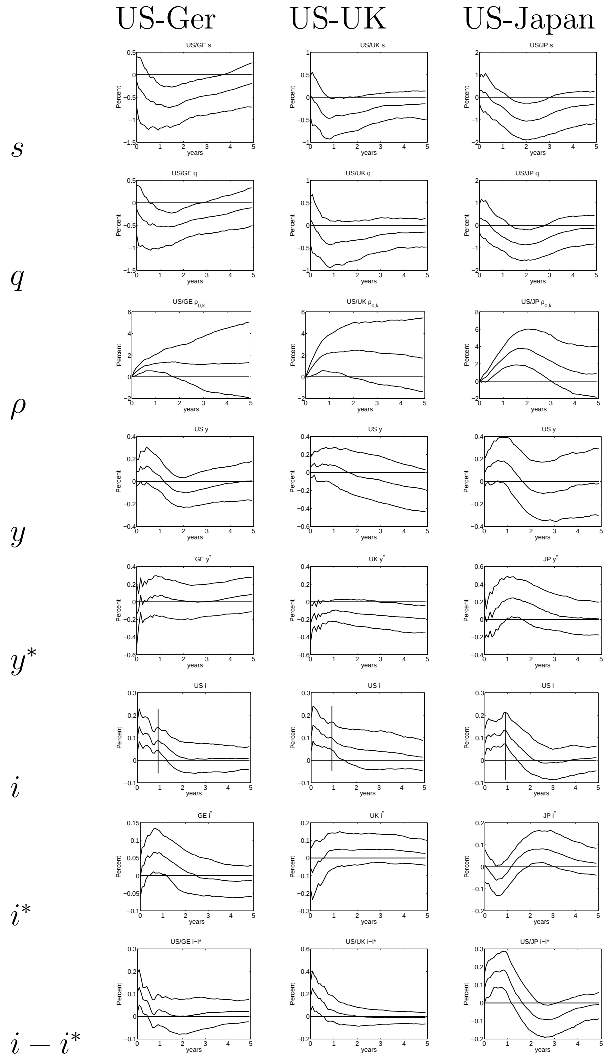

Figure 13: This figure shows impulse responses to a US monetary policy contraction for the **BIG** VAR, using $K = 11$.

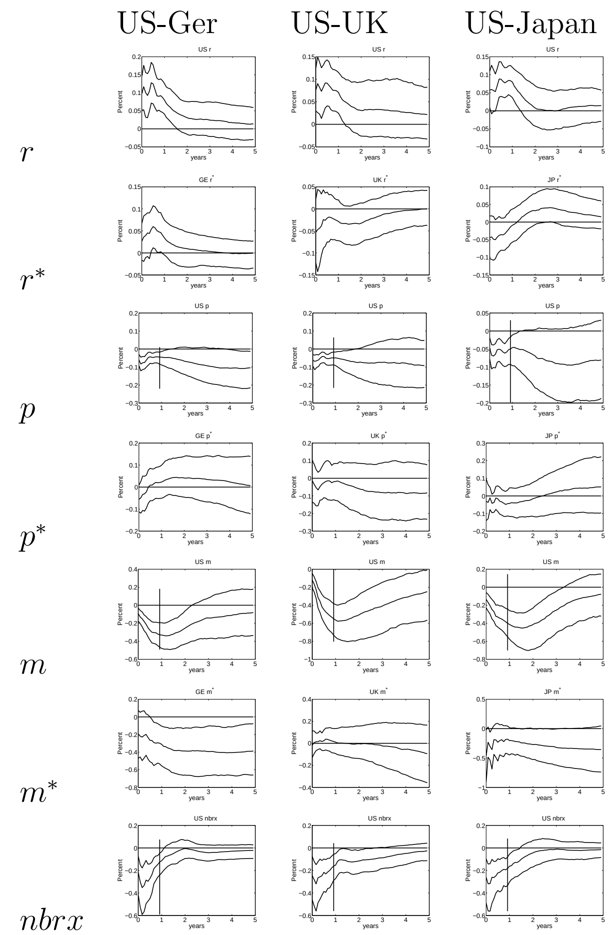

Figure 13, impulse responses to US monetary policy contraction for the **BIG** VAR, continued.

6.4.3 Responses to a Foreign Monetary Policy Contraction

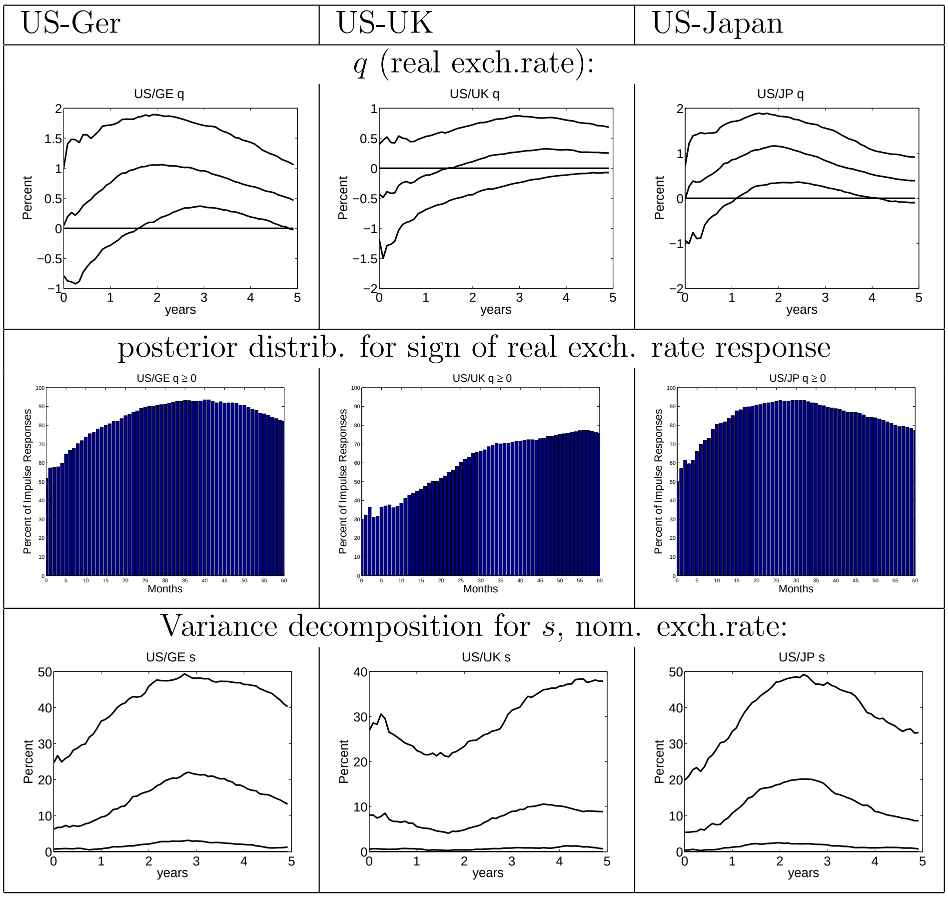

Figure 14: Key results for a **foreign monetary policy contraction**, using the benchmark identification with $i^* - i \geq 0$ in the **GR** specification. Shown are the results for the response of the real exchange rate q, the posterior distribution for its sign distribution and the variance decomposition for the nominal exchange rate.

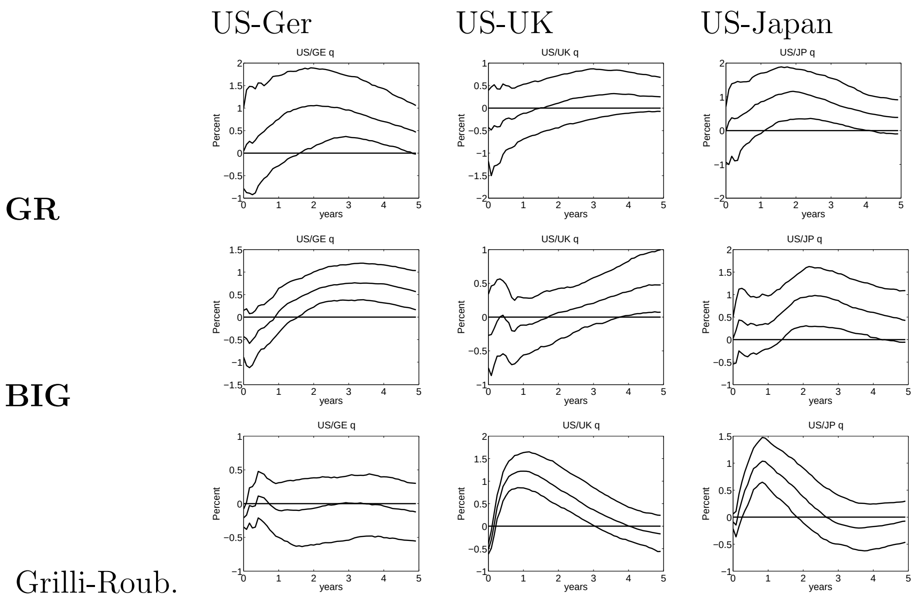

Figure 15: Impulse response function of the real exchange rate, conditional on a foreign monetary policy contraction. We compare the benchmark identification and imposing $i^* - i \geq 0$ in the **GR** VAR specification to the **BIG** VAR specification as well as the recursive Grilli-Roubini identification.

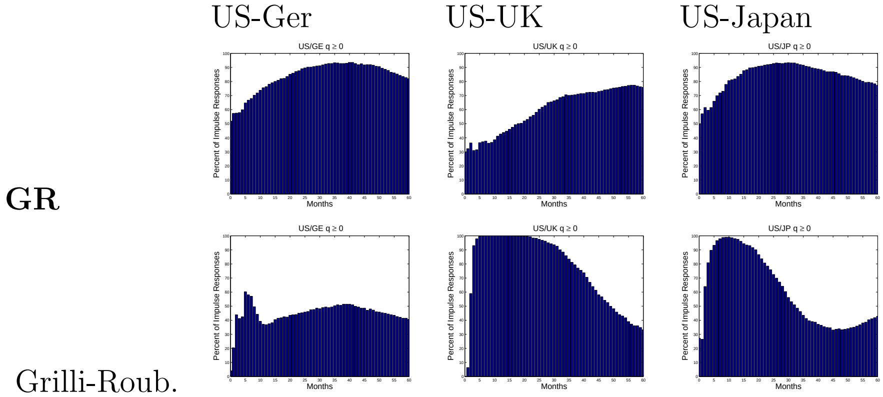

Figure 16: Distribution for the sign $q \geq 0$ of the real exchange rate, conditional on a foreign monetary policy contraction. We compare the benchmark identification and imposing $i^* - i \geq 0$ in the **GR** VAR specification to the **BIG** VAR specification as well as the recursive Grilli-Roubini identification.

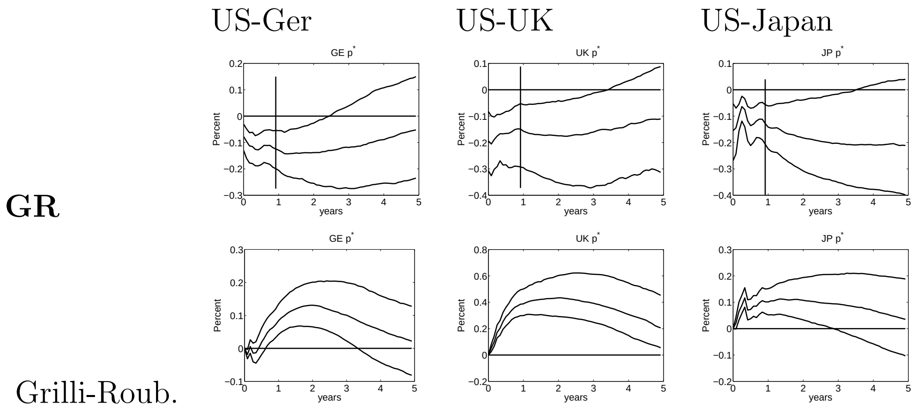

Figure 17: Impulse response function of the price level, conditional on a foreign monetary policy contraction. We compare the benchmark identification and imposing $i^* - i \geq 0$ in the **GR** VAR specification to the recursive Grilli-Roubini identification. Note that there is a considerable price puzzle in the Grilli-Roubini specification, which is avoided with sign restrictions by construction.

GR **BIG** Grilli-Roubini

s

q

ρ

y^*

y

i^*

i

$i^* - i$

p^*

p

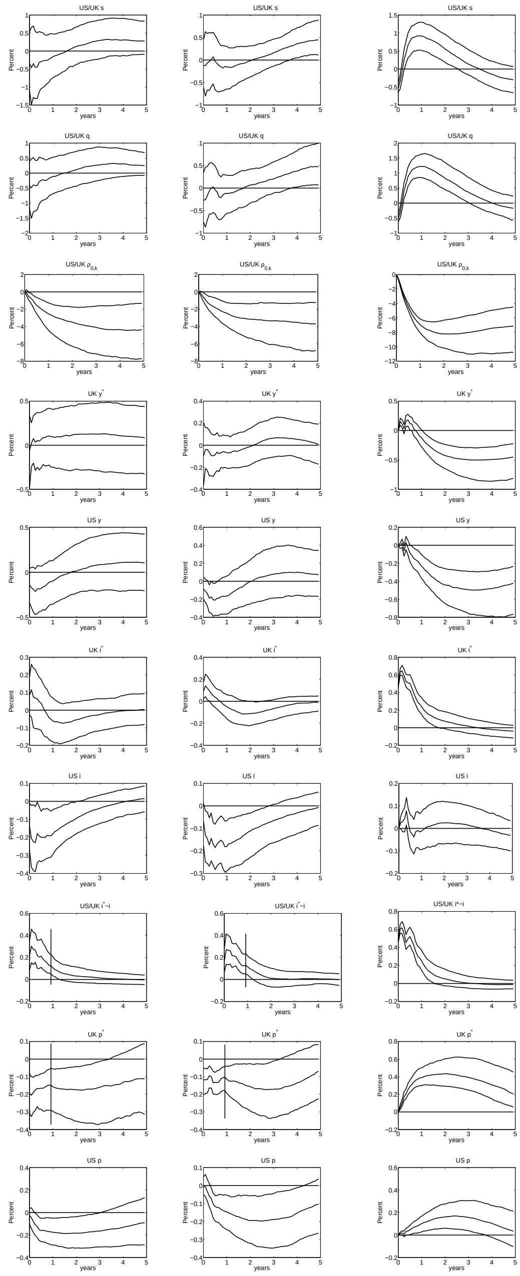

Figure 18: Comparison of results. For a foreign monetary policy contraction and the US-UK country pair, we compare the benchmark identification in the **GR** specification to the original identification of Grilli-Roubini.

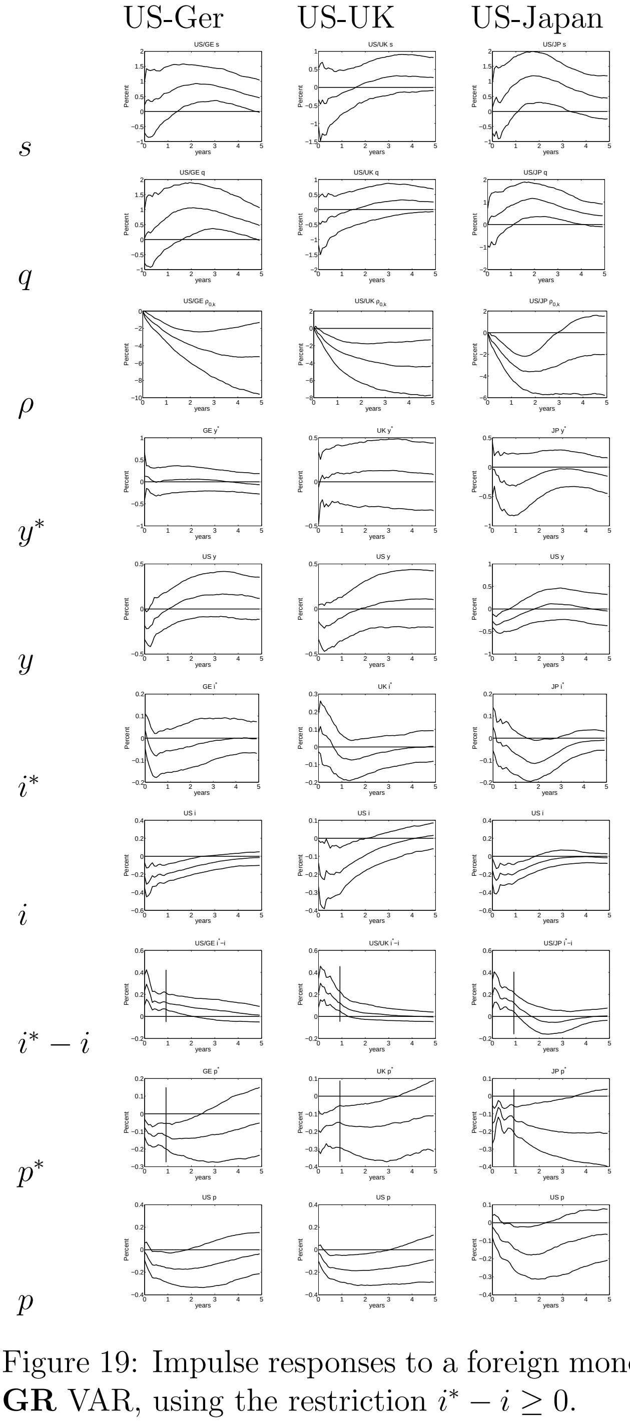

Figure 19: Impulse responses to a foreign monetary policy contraction in the **GR** VAR, using the restriction $i^* - i \geq 0$.

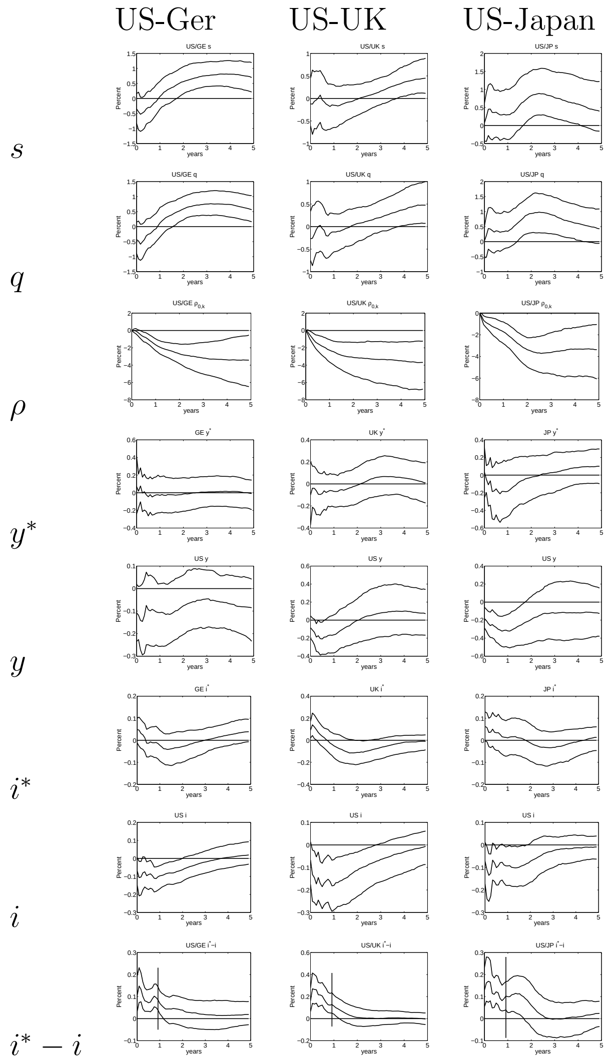

Figure 20: Impulse responses to a foreign monetary contraction. Identification and $i^* - i \geq 0$ for the **BIG** VAR.

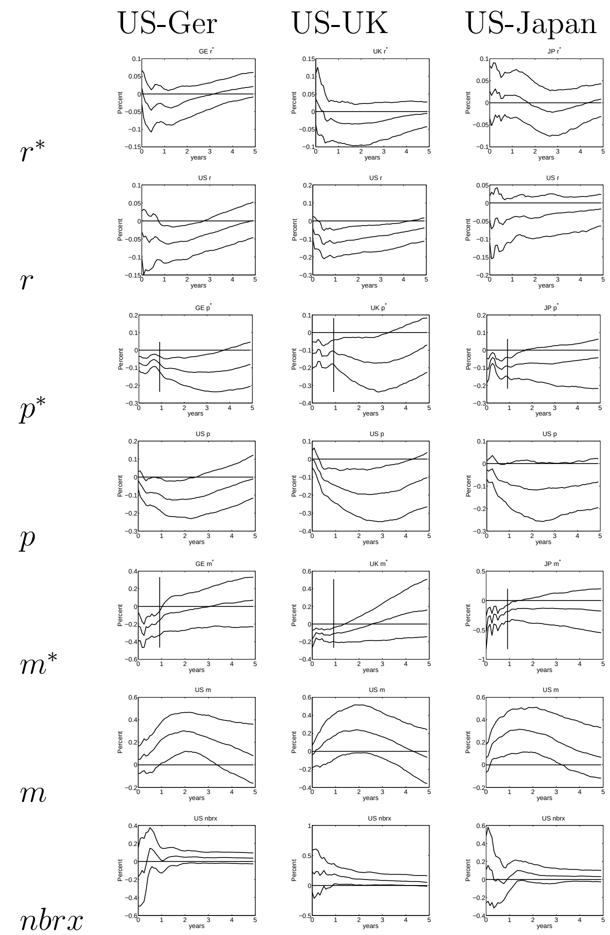

Figure 20, impulse responses to a foreign monetary contraction for the **BIG** VAR, continued.

6.4.4 Variance Decompositions

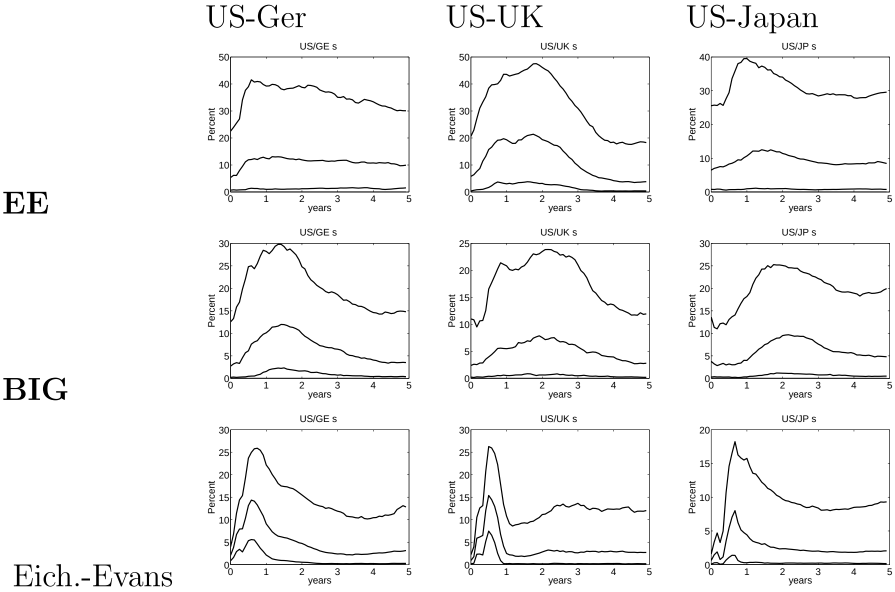

Figure 21: Variance decompositions: the contribution of US monetary policy shocks to the variance of the nominal exchange rate.

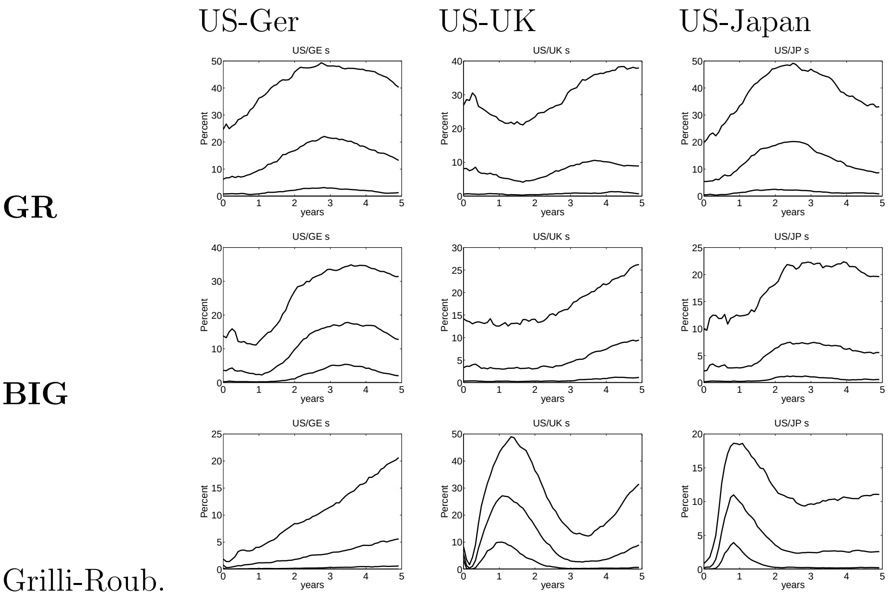

Figure 22: Variance decompositions: the contribution of foreign monetary policy shocks to the variance of the nominal exchange rate.