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

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Abstract
Past empirical research on monetary policy in open economies has found evidence of the ‘delayed overshooting puzzle’ and the ‘forward discount puzzle’. 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. In a first step, we leave the response of the exchange rate agnostically open and find sizeable evidence of both puzzles. In a second step, we additionally rule out the delayed overshooting by construction. Our results indicate that the forward discount puzzle is robust even without delayed overshooting.

Keywords: vector autoregressions, agnostic identification, forward discount puzzle, delayed overshooting, exchange rates, monetary policy

JEL codes: C32, E58, F31, F42

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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. Following the lead of Sims (1972, 1980), however, empirical studies have found different results employing vector autoregressions (VARs) in open economy settings.

To study the effects of monetary policy on exchange rates, e.g. Eichenbaum and Evans (1995) and Grilli and Roubini (1995, 1996) use recursive identification strategies and find a persistent appreciation of the domestic currency for periods up to three years. This finding is known as the ‘delayed overshooting puzzle’, see figure 1, and is also reported in Leeper, Sims and Zha (1996), Clarida and Gali (1994), Kim (2001) and Kim (2005). It is often called the ‘forward discount puzzle’, see figure 2, since a violation of the uncovered interest parity (UIP) condition is implied. Note that there may be a forward discount puzzle even without delayed overshooting.

Recently, this conventional view has come under attack, see e.g. Cushman and Zha (1997), Kim and Roubini (2000) and Faust and Rogers (2003).1 Faust and Rogers (2003) argue, that one needs to “relax dubious identifying assumptions” stemming from e.g. recursive identifications and impose at most rather mild sign restrictions or shape restrictions a priori. In response to monetary policy shocks they find no robust results regarding the timing of the peak response of the exchange rate, but robust

1Other papers that do not find evidence of the delayed overshooting puzzle are e.g. Kalyvitis and Michaelides (2001) and Bjornland (2006).
evidence in favor of large deviations from UIP due to monetary policy shocks.

This paper re-examines these issues by using Uhlig’s (2005) identification strategy of imposing sign restrictions on selected impulse response functions for a certain period following the shock. We focus on two questions. First, is there robust evidence of a delayed overshooting of the exchange rate in response to monetary policy shocks? Second, the delayed overshooting may be interpreted as a sign of failing to appropriately identify monetary policy shocks, see Cushman and Zha (1997). Our identification methodology allows us to rule out the delayed overshooting by construction. The question then is: does the forward discount puzzle still survive or is it just a ‘twin appearance’ of delayed overshooting?²

To analyze the first question, our identification procedure assumes 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, by construction our identification procedure avoids the price puzzle that is often implied by recursive identification strategies. Note that at this stage we do not impose any restrictions on the exchange rate to leave the central question agnostically open. To analyze the second question, our second set of identifying sign restrictions additionally imposes that in response to an increase in domestic interest rates relative to foreign interest rates the exchange rate shows no delayed overshooting and follows the solid line in figure 1. We argue that these sign restrictions are plausible because they most directly reflect what economists have in mind when thinking about monetary policy shocks.

Similar to Faust and Rogers (2003), we apply our identification method to the

²We are grateful to an anonymous referee for this suggestion.
VAR specification used by Eichenbaum and Evans (1995). Following the arguments of Sims and Uhlig (1991) we use a thoroughly Bayesian procedure. 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. We also exploit the Bayesian perspective to ask questions concerning the risk a Bayesian investor faces when betting on violations of UIP. We calculate an implied Sharpe ratio conditional on the monetary policy shock and compare it to Sharpe ratios conventionally observed on e.g. equity markets.

In contrast to Faust and Rogers (2003) who argue that delayed overshooting is a fragile finding, we restore the puzzle originally stated by Eichenbaum and Evans (1995). We find sizeable and robust evidence in favor of a delayed overshooting of the US-German, the US-UK and the US-Japanese bilateral exchange rates as well as confronting the US with an aggregate of the other G7 countries. In line with Eichenbaum and Evans (1995) and Faust and Rogers (2003) there is a robust forward discount puzzle implying a large risk premium conditional on monetary policy shocks. Taking into account the uncertainty, we find that the implied Sharpe ratio for a Bayesian investor is considerably higher than the Sharpe ratio for US stock markets. If we rule out the delayed overshooting puzzle by construction, we do find sizeable and statistically significant risk premia across all country-pairs, and hence, the forward discount puzzle seems to be robust even without delayed overshooting. Notably, our findings are in contrast to Cushman and Zha (1997) and Kim and Roubini (2000) who identify monetary policy shocks by explicitly taking into account the structure of the economy. They find no evidence of delayed overshooting and, in addition, no statistically significant evidence of large deviations from UIP.
As opposed to Faust and Rogers (2003), 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 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. 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). Like Faust and Rogers (2003), Farrant and Peersman (2006) use sign restrictions to identify shocks in an open economy framework. However, their focus lies on the sources of exchange rate movements and they do not address the delayed overshooting and forward discount puzzle.

The structure of the paper is as follows. In section 2 we present the theoretical framework. Section 3 describes our empirical approach. Section 4 reports the results and, finally, section 5 concludes.

2 Overshooting, the Forward Discount Premium and the Sharpe Ratio

The Dornbusch (1976) overshooting of the exchange rate follows from uncovered interest parity (UIP), long-run purchasing power parity (PPP) and a liquidity effect in response to monetary policy shocks, and is a feature that many theoretical models have in common. Recently, stochastic dynamic general equilibrium models with monetary policy shocks have been employed to explain the volatility and the persistence of the exchange rate, see e. g. Alvarez, Atkeson and Kehoe (2002, 2006), Chari, Kehoe and McGrattan (2000), Kollmann (2001) and Bergin (2006). In line with Dornbusch’s
prediction, these models imply that an increase in the interest rate differential leads to an impact appreciation of the exchange rate followed by depreciation, see the impulse responses in Bergin (2006). The Dornbusch overshooting model and UIP 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.

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 nominal 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 $k$,

$$i^*_{0 \rightarrow k} = \sum_{j=0}^{k-1} i^*_j, \quad i^*_{0 \rightarrow k} = \sum_{j=0}^{k-1} i^*_j.$$

Define the forward discount premium

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

which is the gain due to the monetary policy shock (compared to the baseline scenario without that shock) from borrowing foreign currency for $k$ 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, $\rho_j$ stays flat for $j \geq j^*$, if $\xi_j = 0$ for $j \geq j^*$.

(Conditional) UIP 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. In the empirical literature it is common to assume rational expectations and to employ regressions of exchange rate movements on interest rate differentials to study the (unconditional) forward discount premium, see Hansen and Hodrick (1980), Fama (1984), Froot and Thaler (1990) and Lewis (1995) who provides an excellent survey. Our study of forward premia and the associated empirical Sharpe ratios conditional on monetary policy shocks using VARs complements this literature.

The forward discount puzzle is often stated as an implication of delayed overshooting. The Dornbusch (1976) overshooting model results from adding to (conditional) UIP the liquidity effect of monetary policy shocks, $i_0 - i_0^* > 0$, and the assumption that in the long run the impulse response for the exchange rate must reach an appreciated value after the monetary contraction due to long run PPP. Under these assumptions, $s_0$ should appreciate on impact followed by a depreciation to the long-run value. However, the empirical literature has found delayed overshooting in response to US monetary policy contractions, see figure 1.\(^3\)

\(^3\)In reaction to the puzzling empirical findings, there have been some recent attempts to develop
There are a couple of things to note at this point. First, while exchange rate movements certainly depend on the policy rules of both countries, here the liquidity effect of monetary policy shocks, $i_0 - i_0^* > 0$, is essential. Regarding our empirical analysis, we argue that it is sufficient to study the interest rate differential since it reflects cross-country policy differences. 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 (explicitly or implicitly) significant violations of conditional UIP, this may not suffice for an investor contemplating exploiting this deviation at some date $t$. For an investor it is essential to measure the risk-adjusted performance of an investment strategy. For this we evaluate the Sharpe ratio that measures the reward to risk and is defined as the ratio of the mean excess return to the standard deviation. Let $SR_{\zeta_k}$ and $SR_{\rho_k}$ be the Sharpe ratios for the hedging strategy $\zeta_k$ and $\rho_k$, respectively, from the perspective of a Bayesian investor who remains uncertain about the precise impact of monetary policy shocks on the forward discount premium. This is due to uncertainty regarding the reduced-form dynamics of the economy as well as uncertainty regarding the precise nature of monetary policy shocks.

For equity (and the reverse strategy described here, i.e. for going long on equity theoretical explanations. Gourinchas and Tornell (2002, 2004) explain the delayed overshooting as the interaction of learning about the current state and the intrinsic dynamic response of interest rates to monetary shocks. Other recent papers like e.g. Moore and Roche (2002), Backus, Foresi and Telmer (2001), Verdelhan (2006) and Bacchetta and van Wincoop (2005) consider habit-persistence or rational inattention.
and borrowing at the short rate) and an investment horizon of one year, a Sharpe ratio of 0.5 is common as is well known from the literature. To compare this benchmark value to Sharpe ratios associated with investment horizons of one month, we assume that the monthly forward premium $\zeta_k$ is i.i.d. with mean $\tilde{\mu}$ and variance $\tilde{\sigma}^2$. Hence, $\rho_k$ is i.i.d. with mean $k\tilde{\mu}$ and variance $k\tilde{\sigma}^2$. Given an annual benchmark value of 0.5, the monthly benchmark Sharpe ratios for $\rho_k$ is then given by $\frac{0.5}{\sqrt{k}}$.

3 Our Approach

3.1 Methodology


Consider a vector autoregression

$$Y_t = B_{(1)}Y_{t-1} + B_{(2)}Y_{t-2} + \ldots B_{(n)}Y_{t-n} + u_t, \quad E[u_t u_t'] = \Sigma,$$

$t = 1, \ldots, T$, for some $\ell$-dimensional vector of variables $Y_t$, coefficient matrices $B_{(i)}$ of size $\ell \times \ell$ and a variance-covariance matrix for the one-step ahead prediction error $\Sigma$. The key to identification is to represent the one-step ahead prediction error $u_t$ as a linear combination of orthogonalized ‘structural’ shocks,

$$u_t = A v_t, \quad E[v_t v_t'] = I.$$

Traditional identification strategies impose a recursive ordering or structural restric-
tions 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 $a \in \mathbb{R}^\ell$ 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 $a = \tilde{A}\alpha$, where $\tilde{A}\tilde{A}' = \Sigma$ is the Cholesky decomposition of $\Sigma$ and $\alpha$ is an $\ell$-dimensional vector of unit length. Let $ir_i(k) \in \mathbb{R}^\ell$ be the vector response at horizon $k$ to the $i$-th shock in a Cholesky-decomposition of $\Sigma$. Then the impulse response $ir_a(k)$ for $a$ is given by

$$ir_a(k) = \sum_{i=1}^{\ell} \alpha_i \cdot ir_i(k).$$

The identifying restrictions we impose to identify an impulse vector characterizing monetary policy shocks are that $(ir_a(k))_j \geq 0, j \in J_+$ and $(ir_a(k))_j \leq 0, j \in J_-$ for some subsets of variables $J_+$ and $J_-$ and some horizon $k = 0, \ldots, K$.

We use a Bayesian prior for the 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 $\tilde{A}\tilde{A}' = \Sigma$ and can even be random as long as the choice of the decomposition is independent of $\alpha$. For further methodological details see Uhlig (2005).

### 3.2 Identification of Monetary Policy Shocks

The first choice to be made is the selection of variables. To assure comparability and similar to Faust and Rogers (2003), we use the specification chosen by Eichenbaum
Table 1: Identification of Monetary Policy Shocks

<table>
<thead>
<tr>
<th>Variables:</th>
<th>y, y*, p, nbrx, i, i*, s</th>
</tr>
</thead>
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Identification I: \( p(k_1) \leq 0, \ nbrx(k_1) \leq 0, \ i(k_1) \geq 0 \)

Identification II: \( p(k_2) \leq 0, \ nbrx(k_2) \leq 0, \ i(k_2) \geq 0, \ i(k_2) - i^*(k_2) > 0 \)
\( s(0) < 0, \ |s(j)| < |s(j - 1)|, \ |s(j)| < |s(0)| \)

Notes: Identification I is based on the horizon \( k_1 = 0, \ldots, K_1, \ K_1 = 11 \), i.e. one year. Identification II is based on horizon \( k_2 = 0, \ldots, K_2, \ K_2 = 5 \), \( j = 0, \ldots, J, \ J = 2 \) and \( j = J + 1, \ldots, \bar{J}, \bar{J} = 23 \). \( y \) and \( y^* \) denote domestic and foreign industrial production, \( p \) is the domestic price level, \( i \) and \( i^* \) are the domestic and foreign short-term rates, \( nbrx \) denotes the ratio of nonborrowed reserves to total reserves and \( s \) is the nominal exchange rate.

and Evans (1995). The specification investigates 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 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) and Strongin (1995).

Table 1 provides a detailed description of the VAR specification and the set of sign restrictions used. In order study the robustness of the delayed overshooting puzzle our first identification set imposes that prices as well as monetary aggregates like the ratio of nonborrowed to total reserves or the money stock do not rise in response to a monetary contraction. Thus, ‘price puzzles’ are avoided by construction. We furthermore match conventional wisdom by imposing that, in response to a domestic monetary policy contraction, domestic interest rates do not fall. Crucially, at this stage, we leave the exchange rate unrestricted to leave the question at hand agnostically open.

As also argued by Cushman and Zha (1997), the delayed overshooting may be a sign
of failing to appropriately identify the monetary policy shock. The methodology here allows us to impose the absence of delayed overshooting in the presence of a (relative) monetary tightening: this is what we do in the identification strategy II. We impose two additional assumptions. First, the US Dollar initially appreciates in response to a monetary policy shock and then persistently depreciates. Second, the interest rate differential is constrained to be positive. The question now is: does the forward discount puzzle still survive or is it just a ‘twin appearance’ of delayed overshooting? Put differently, we ask the following question: if there is an increase in the interest rate differential and the exchange rate shows no delayed overshooting, is then UIP fulfilled, i.e. is there a forward discount premium?

For the horizon of the sign restrictions, our first identification scheme uses $k_1 = 0, \ldots, K_1 = 11$, i.e. one year. Choosing shorter restrictions leaves too much room for spurious effects while imposing a longer horizon imposes an implausibly long duration of the liquidity effect. Since our second identification avoids the delayed overshooting puzzle by construction and therefore imposes very tight identifying assumptions, we choose a horizon of half a year, $k_2 = 0, \ldots, K_2 = 5$ for the responses of the interest rates, the price level and the monetary aggregate. Moreover, we restrict the exchange rate to appreciate on impact and to depreciate during the first quarter following the shock. In addition, to ensure the persistency of the depreciation, for two years following the shock, the response of the exchange rate is constrained to be smaller in absolute value compared to the impact response. Uhlig (2005) contains some discussions how results vary with the horizon when applying the methodology in the analysis of monetary policy shocks for US data. To save space we decided to omit results for other choice of the horizon length. We experimented with the set of identifying sign restrictions and
provide a discussion in our robust analysis.

We study three country pairs: the US and Germany, the US and the UK, and the US and Japan. To additionally take into account cross-sectional information we construct an aggregate of the six G-7 countries other than the US. For the three country pairs we employ monthly data from 1975:07 to 2002:07 while for the aggregate we have monthly data from 1977:04 to 2001:12 due to data limitations. For a detailed description of the variables, the data sources and the construction of the aggregate see table 2 in the appendix.

A Bayesian VAR with 6 lags in levels of the logs of the series has been fitted to the monthly data except for using interest rates directly. No constant or time trend are included. This may result in a slight misspecification but the results are more robust because of the interdependencies in the specification of the prior between these terms and the roots in the autoregressive coefficients, see Uhlig (1994). To ensure comparability, the choice of 6 lags follows the choices made in the literature.

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 and draws from the space of possible impulse vectors. Inference statements are based on the joint draws that satisfy the sign restrictions for the impulse responses. We typically show the median and the 16% and 84% quantiles of the distribution for the points on the impulse response functions.
4 Empirical Results

4.1 The Delayed Overshooting Puzzle

We start by considering the identification strategy by Eichenbaum and Evans (1995) who impose the recursive ordering \([y, p, y^*, i^*, nbrx, i, s]\) and define a monetary contraction by a fall in the ratio of nonborrowed reserves to total reserves. Figure 3 considers the US-UK as an example and plots the impulse responses of the nominal exchange rates and the price level considering two data sets. The first data set ends in 1990:05 and corresponds to the one in Eichenbaum and Evans (1995). The second data set we use for our analysis is updated and ends in 2002:07. For the original data set we find a persistent appreciation for periods up to three years as in Eichenbaum and Evans (1995). Using the updated data set instead, we find a slightly less strong reaction of the exchange rate but the pattern appears qualitatively to be very similar. Since the impulse responses of the US-German and the US-Japanese bilateral exchange rates follow a similar pattern we omit the figures to save space. The second line of figure 3 reveals that there emerges a huge price puzzle that is even more pronounced for the updated data set. 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. Instead we propose to employ sign restrictions to identify monetary policy shocks.

Our first set of identifying assumptions avoids the price puzzle by construction and restricts the US interest rate \(i\) not to be negative and the ratio of nonborrowed reserves to total reserves \(nbrx\) not to be positive. Figure 4 plots the impulse responses of the nominal and real exchange rate, \(s\) and \(q\), the interest rate differential \(i - i^*\), the forward
discount premium $\rho$ and the associated Sharpe Ratio $SR_\rho$.\(^4\) In response to monetary policy shocks the interest rate differentials between the domestic and foreign short-term interest rates increase which matches conventional wisdom. We thus turn to a discussion of the exchange rate response.

The first line of figure 4 shows the impulse responses of the nominal exchange rate and should be compared to figure 1. The results show that the US-German and the US-UK bilateral nominal exchange rates appreciate during the first two years and then depreciate, while the US-Japanese exchange rate seems to peak within the first year (however, the response is only weakly significant). Our results suggest that, in contrast to e.g. Kim and Roubini (2000) and Faust and Rogers (2003), there is robust evidence of a delayed overshooting, but that the delay is considerably shorter than the three-year horizon found by Eichenbaum and Evans (1995). Considering the US-G7 case we observe a quite persistent appreciation strongly confirming the delayed overshooting puzzle. The second line of figure 4 shows the impulse response of the real exchange rate. It turns out that the nominal and the real exchange rates respond very similar to monetary policy shocks which is in line with the results in Eichenbaum and Evans (1995). In the following we therefore focus on the nominal exchange rate.

When exactly do the responses of the exchange rate reach their lowest value, i.e. for how many periods do the exchange rates appreciate? To answer this question the first column of figure 5 shows the posterior distribution of the timing of the peak appreciation, i.e. the distribution of the month containing the lowest point of an impulse response drawn from the posterior. It could be that the posterior distribution of the peak months gives misleading results, if e.g. impulse responses showing mild apprecia-

\(^4\)The other variables like e.g. the price level are restricted and therefore are omitted to save space. US industrial production does not show a significant response.
tions peak early and those with strong appreciations peak late. We therefore analyze the size of the peak appreciation in addition to the timing. The second and third column of figure 5 show the posterior joint distributions of the timing of the peak and its altitude on a monthly and quarterly basis. It turns out that for the US-UK, US-Germany and the US-G7 case there is some mass on early peaks that occur within the first five months. The joint posterior distribution of the timing of the peak and its altitude reveals that the early peaks of the US-German and the US-G7 exchange rate are of intermediate size. However, in the US-UK case these early peaks are rather mild appreciations. For the US-UK and US-Germany most of the distribution is centered around months 12 to 26. In contrast, in the US-Japan case the distribution is centered around months 5 to 15. Between 8 and 12 percent of the responses do not reach their lowest value within the considered time horizon. The distribution of the peak appreciation is more dispersed for the US-G7: most of the distribution is found between months 15 to 45. Here, around 20 percent of the responses do not peak within the considered time horizon.

4.2 The Market Price for the Forward Discount Risk

How large are the forward premia conditional on monetary policy contractions? What risk does a Bayesian investor face when betting on violations of UIP and what is the reward of doing so?

To analyze these questions the last two lines of figure 4 focus on the impulse responses of the forward premia for k-month holding periods, $\rho_k$, as well as the corresponding Sharpe ratios $SR_{\rho}$. On stock markets annual Sharpe ratios of 0.5 are commonly observed. The dotted line corresponds to this benchmark value transformed to
a monthly basis as described in section 2. The impulse response of the accumulated
forward premium should be compared to figure 2.

In line with Eichenbaum and Evans (1995) and Faust and Rogers (2003), the em-
pirical results reveal that across all country pairs there is a forward discount premium
but there is also considerable uncertainty regarding its size. When taking this uncer-
tainty into account and calculating the Sharpe ratio $SR_\rho$, one finds values between 1
and all the way up to 2.5. Analyzing the one-period forward premium and the asso-
ciated Sharpe ratio yields similar conclusions. We conclude that the reward for risk
of betting on violations of UIP is higher than the ‘benchmark’ reward for holding risk
on the stock market by a factor somewhere between 2 and 5. Note that the Sharpe
ratio is calculated here conditional on a single monetary policy shock. To trade on
this comparably high reward requires a portfolio that enables investors to solely fo-
cus on the forward discount risk associated with an observed monetary policy shock,
hedging against all other disturbances. This implies the use of a portfolio with appro-
priate derivates and/or a trading strategy which is more sophisticated than a simple
carry trade. Developing the appropriate hedging strategies in light of existing financial
market instruments is beyond the scope of this paper.

4.3 No Delayed Overshooting and Forward Discounting

Alternatively, the delayed overshooting observation may be interpreted as a failure in
appropriately identifying a monetary policy shock, rather than a result. The interesting
question then is: is there nonetheless an economically significant violation of UIP on
which smart investors may take bets if they observe a surprise relative tightening of
monetary policy?
To analyze this question, we employ our second identification strategy and additionally impose that domestic interest rates rise relative to foreign interest rates and that nominal exchange rates appreciate on impact and do not exhibit delayed overshooting, see table 1. Figure 6 plots the impulse responses of the nominal exchange rate $s$, the interest rate differential $i - i^*$, the forward discount premium $\rho$ and the associated Sharpe Ratio $SR_{\rho}$. The horizontal lines indicate that the impulse responses of the exchange rate as well of the interest rate differential are restricted by assumption. Note that we used the same scaling as in figure 4 to make comparisons easier. By construction, the exchange rates appreciate on impact and then depreciate to a long-run appreciated value, in line with the Dornbusch model.

First, consider the US-Germany case. Though the increase in the interest rate differential is approximately of the same size as before, the response of the US-German nominal exchange rate is much weaker. There occurs a sizeable and persistent forward premium and UIP is violated. Similar patterns can be observed considering the US-UK, the US-Japanese and the US-G7 cases. There, the increases in the interest rate differentials are slightly lower than under identification I and, therefore, the forward premia are smaller and less persistent, but still significantly positive for the horizon of one year. The associated Sharpe ratios lie between 1 and 1.5 such that the reward for risk of betting on violations of the uncovered interest parity is high.

All in all, our results indicate that across all country pairs there are sizeable and significantly positive forward premia even when there is no delayed overshooting. Moreover, the market price for risk of betting on violations of UIP is high. Notably, our findings are in contrast to Cushman and Zha (1997) and Kim and Roubini (2000) who do not find statistically significant large deviations from UIP.
4.4 Monetary Policy and Exchange Rate Volatility

In this section we have a closer look at the identified monetary policy shocks associated with identification I. To do so we construct a stance of US monetary policy as a 12-month moving average window of the monetary policy shocks. The first panel of figure 7 plots the median of the US monetary policy stance considering the three country pairs US-UK, US-Germany and US-Japan. It is evident that the identified US monetary policy shocks follow the same pattern across country-pairs which makes us feel confident concerning our identification strategy. The second graph focuses on the US-G7 and plots the 14, 50 and 86 percent quantiles of the US monetary policy stance. Major episodes of US monetary tightening and loosening are reflected. First of all, between 1980 and 1983 the monetary policy stance shows a clear spike which corresponds to the Volcker Recession where interest rates increased to historical highs due to tight monetary policy. In 1987 the US economy was hit by the stock market crash such that the US Fed, now under the guidance of Alan Greenspan, decreased the fed funds rate target. Because of inflationary pressure after the crash the Fed increased the fed funds rate which is reflected by our monetary policy shocks. The effects of the Gulf War in 1990 are clearly shown by a sharp fall of our monetary policy stance. Moreover, the economic expansions in the 1990s as well as the 2001 recession are reflected by the monetary policy shocks. Overall, the largest spikes take place during the Volcker Recession and the Gulf War while during the 1990s shocks are not as large. This is line with Rubio-Ramirez, Waggoner and Zha (2005) who find only mild fluctuations in the interest rate during that time period.

To analyze how much US monetary policy shocks account for exchange rate fluctu-
iations, we start with a comparison of the data with the prediction and the projection. The projection is the prediction if one had also known the monetary policy shocks but not all the other shocks and, thus, can be interpreted as the prediction available to the Fed planning a particular sequence of shocks. We consider two special periods: first, the Volcker Recession of the early 1980s as an example of extreme monetary policy, and, second, the mid 1980s as an example of a strong depreciation of the US exchange rate. The last two panels of the first line in figure 7 compare the data with the median of the prediction and projection. Here, we omit the error bands but show them in the forecast error variance decompositions later on. The figures clearly show that monetary policy shocks improve the prediction and, hence, partly account for the movements in the exchange rates.

To analyze the impact quantitatively, the second line of figure 7 presents the forecast error variance decomposition for the exchange rate movements explained by a US monetary policy shock considering the four country pairs. US monetary policy shocks account for roughly 10 percent of the US-German, the US-UK and the US-Japanese and about 20 percent of the US-G7 nominal exchange rate fluctuations at the median estimate. 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, respectively, at lags 31 – 36 using data until 1990. Other examples for studies which find that nominal shocks have a substantial contribution in explaining exchange rate fluctuations are Clarida and Gali (1994), Rogers (1999) and Farrant and Peersman (2006). Our result that monetary policy shocks have a lower impact on 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.

4.5 Robustness Analysis

In our robustness analysis we considered alternative sign restriction sets, that e. g. involved additional restrictions on US industrial production. This did not change our findings. Considering the second identification set, it turned out to be crucial to restrict the interest rate differential since otherwise these responses were not significantly positive in all cases. Moreover, it was important to require the response of the exchange rate to be smaller in absolute value compared to the impact value for a long period following the shock to avoid strong appreciations in the medium run. Similar results could have been achieved by increasing the horizon $J$, however, this identification scheme implies much stricter assumptions and we decided to stick to the presented one.

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 our robustness analysis, we extended the benchmark VAR specification by adding possibly important variables. This extended VAR specification treats the US and the foreign country in a symmetric manner and takes into account some foreign monetary measure $m^*$, the foreign price level $p^*$ and long-term interest rates $r$ and $r^*$. Finally, we varied the lag length. With respect to these sensitivity checks the results appeared to be robust and the main conclusions were not altered.
5 Conclusions

This paper has estimated the effects of monetary policy shocks on 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 restore the delayed overshooting and the forward discount puzzle originally stated by Eichenbaum and Evans (1995) that recently have been claimed to be fragile. However, compared to Eichenbaum and Evans (1995) the quantitative properties change. With $2/3$ probability the peak appreciation happens during the first two years after the shock for the US-German and the US-UK, and during the first year for the US-Japanese exchange rate, while the US-G7 exchange rate exhibits a persistent delayed overshooting.

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 values than those found on asset markets. Even if we rule out the delayed overshooting puzzle by construction, we do find statistically significant evidence of sizeable forward premia.

We argue that our identifying sign restrictions are plausible because they generate monetary policy shocks that, firstly, match important episodes of US monetary tightening and loosening, and, secondly, most directly reflect what economists have in mind (or how economists informally evaluate empirical results) when thinking about monetary policy shocks. We view this identification strategy as a consequent pursuit of the agenda put forth by Sims (1980) of avoiding incredible identifying restrictions.
References


## A Appendix

Table 2: Data Sources

<table>
<thead>
<tr>
<th>Variable</th>
<th>Description</th>
<th>Source</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y$</td>
<td>index of US industrial production</td>
<td>IMF Washington, line 66</td>
</tr>
<tr>
<td>$p$</td>
<td>US consumer price index</td>
<td>IMF Washington, line 64</td>
</tr>
<tr>
<td>$nbrx$</td>
<td>US non-borrowed reserves + extended credit / total reserves</td>
<td>Fed. Reserve Bank St. Louis</td>
</tr>
<tr>
<td>$i$</td>
<td>US 3-months treasury bill rate</td>
<td>IMF Washington, line 60c</td>
</tr>
<tr>
<td>$r$</td>
<td>US 10-year bond yield</td>
<td>IMF Washington, line 61</td>
</tr>
<tr>
<td>$y^*$</td>
<td>foreign industrial production</td>
<td>IMF Washington, line 66</td>
</tr>
<tr>
<td>$m^*$</td>
<td>foreign money supply M1 (UK: M0)</td>
<td>IMF Washington, line 39 (UK: Bank of England)</td>
</tr>
<tr>
<td>$p^*$</td>
<td>foreign consumer price index</td>
<td>IMF Washington, line 64</td>
</tr>
<tr>
<td>$i^*$</td>
<td>foreign short-term interest rate</td>
<td>IMF Washington, line 60c</td>
</tr>
<tr>
<td></td>
<td>UK: 3-months treasury bill rate</td>
<td></td>
</tr>
<tr>
<td></td>
<td>Germany, Japan: short-term money market rate</td>
<td>IMF Washington, line 60b</td>
</tr>
<tr>
<td>$r^*$</td>
<td>foreign 10-year bond yield</td>
<td>IMF Washington, line 61</td>
</tr>
<tr>
<td>$s$</td>
<td>nominal exchange rate in dollar per foreign currency</td>
<td>Fed. Reserve Bank St. Louis</td>
</tr>
<tr>
<td>$q$</td>
<td>real exchange rate per foreign currency</td>
<td>derived</td>
</tr>
<tr>
<td>$\rho$</td>
<td>forward discount premium (accumulated)</td>
<td>derived</td>
</tr>
<tr>
<td>$\xi$</td>
<td>forward discount premium (for one period)</td>
<td>derived</td>
</tr>
</tbody>
</table>

**Notes:** Variables are in logs except for interest rates. $*$ refers to the foreign country. We consider Germany, UK and Japan. In addition we consider the aggregate of the G7 countries Canada, Germany, France, Italy, Japan and the UK.
A.1 Aggregation Method

To construct an aggregate of the G7 countries Canada, Germany, France, Italy, Japan and the UK we first take growth rates to remove national basis effects. As country specific weight we consider the country’s GDP relative to the total 6-country GDP calculated at purchasing power parity (PPP) values. The aggregated growth rates are then constructed as a weighted sum of the individual growth rates. To calculate levels the aggregate growth rates are cumulated starting from the initial base year.

A.2 Figures

Figure 1: A Stylized Representation of the Delayed Overshooting Puzzle

![Figure 1](image1)

Figure 2: A Stylized Representation of the Forward Discount Puzzle

![Figure 2](image2)
Figure 3: The US-UK: Recursive Identification

Notes: The identification scheme proposed by Eichenbaum and Evans (1995) is applied. The median impulse response and the 16 % and 84 % quantiles of the distribution are shown.
Figure 4: Evidence on the Delayed Overshooting Puzzle

Notes: Results refer to identification I (see table 1). The median impulse response and the 16 % and 84 % quantiles of the distribution are shown.
Figure 5: Properties of the Peak of the Nominal Exchange Rate $s$

Notes: Results are conditional on a US monetary policy contraction and refer to identification I (see table 1) based on 10000 draws. The marginal distribution as well as the posterior distribution for the size and timing of the peak appreciation $s(k)$ are shown. The first and second columns show the distribution on a monthly scale while the third columns shows the results over the first 20 quarters.
Figure 6: Evidence on the Forward Discount Puzzle

Notes: Results refer to identification II (see table 1). The median impulse response and the 16% and 84% quantiles of the distribution are shown.

Figure 7: Monetary Policy Shocks and Exchange Rate Volatility

Notes: Results refer to identification I (see table 1). The panels in the first row refer to median values. For the US-G7 monetary policy stance the 16% and 84% quantiles of the distribution are additionally shown. The second row reports the forecast error variance decompositions referring to a US monetary policy shock. The median and the 16% and 84% quantiles of the distribution are shown.